Cohort Size and The Marriage Market:

Explaining Nearly a Century of Changes in U.S. Marriage Rates*

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Current Draft, June, 2012. First draft, February, 2012

Abstract

In the paper, we propose an explanation for almost a century of changes in the U.S. marriage rates. We do this in three stages. In the first stage, we use reduced-form methods to provide evidence on the following two results. First, a single variable, cohort size, can account for almost all the variation in marriage rates since the early 1930s for both blacks and whites. Second, an increase in cohort size reduces marriage rates, whereas a decline in cohort size has the opposite effect. The most convincing evidence on the relationship between cohort size and marriage rates is obtained by using as a source of exogenous variation differences across states in mobilization rates during World War II. In the second stage, we develop a dynamic search model of the marriage market that can generate the observed patterns. Using the model we show that, qualitatively, it can generate the negative relationship between cohort size and marriage rates. We then derive a testable implication from the model: an increase in cohort size reduces the age difference at marriage and vice versa. In the last stage, we investigate whether the model is consistent with the data. We first use the derived implication to test the model and fail to reject it. We then estimate the model and evaluate

*We are grateful to Guillermo Beylis, Pierre-André Chiappori, Edgar Cortes, Kenneth Mirkin, Dmitry Plotnikov, and Nico Voigtlaender for helpful comments. Research reported in this publication was supported by the National Institutes of Health under Award Numbers 2T32HD7545-11 and 5T32HD7545-10. The content is solely the responsibility of the authors and does not necessarily represent the official views of the National Institutes of Health.

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whether it can quantitatively match the documented variation in marriage rates. We find that the search model can explain most of the observed variation.

1 Introduction

Changes in marital patterns have major implications for many variables that are of interest to economist and policy makers. They affect, to name a few, fertility rates, children's welfare, children's education, labor force participation, hours of work, income inequality, the fraction of individuals on welfare, population growth, and workers' productivity. In spite of this, there is no general explanation for the variation in marriage formation over time and across races. The existing theories, which will be discussed in the next section, only apply to specific periods or specific groups of individuals.

The main contribution of this paper is to provide an explanation for nearly a century of changes in U.S. marriage rates. We show that changes in cohort size on their own explain virtually the entire variation in marriage rates since the early 1930s for both blacks and whites. The paper is divide into three parts.

In the first part, we present reduced-form evidence which indicates that an increase in cohort size generates a decline in marriage rates and that a reduction in cohort size has the opposite effect. The reduced-form evidence is provided in three steps. We first employ time-series variation to study the relationship between changes in marriage rates and changes in cohort size. Our results show that, for both blacks and whites, there is a strong and negative relationship between these two variables. They also indicate that changes in cohort size account for around 70 to 80 percent of the variation in marriage rates for both black and white populations. We then use cross-state variation to provide additional evidence on the link between cohort size and marriage rates. The variation across states confirms the time-series finding that there is a strong and negative relationship between the two variables. In the third step, we provide what is arguably the most convincing evidence on the hypothesis that there is a causal relationship between changes in cohort size and variation in marriage rates. Using an idea proposed by

Acemoglu, Autor, and Lyle (2004), we employ as a plausible source of exogenous variation in cohort size differences in mobilization rates across states during World War II. The results show that an increase in cohort size generates a decline in marriage rates whereas a decline has the opposite effect. The exogenous variation, therefore, gives results that are consistent with the time-series and cross-state variation.

In the second part of paper, we propose a model with the potential of generating the relationship observed in the data between changes in cohort size and changes in marriage rate. We develop a dynamic search model of the marriage market with the following key feature. Women can marry only when young whereas men can marry when young and old. This modeling choice is based on the fact that women are fertile only when young but men can have children also when old, and on the common insight that the main reason for the existence of marriage is that it is an effective arrangement for the upbringing of children. This key feature of the model implies that men search longer than women for a spouse, which is the pattern that enables us to explain the link between changes in cohort size and changes in marriage rates. Using the model, we prove two results. First, we show that a positive change in cohort size has the effect of reducing the marriage rate and a decline in cohort size has the opposite effect. The model can therefore qualitatively explain the negative relationship between those two variables. We then derive an implication that can be used to test the model. We show that in our dynamic search model, an increase in cohort size has the effect of reducing the age difference between spouses.

In the last part of the paper, we test the ability of the proposed model to explain the observed data. We use the implication derived in the theory part of the paper as our first test. We find that a positive change in cohort size generally reduces the age difference at marriage. The search model is therefore consistent with the data and cannot be rejected. We then estimate the model and evaluate whether it can quantitatively explain the changes in marriage rates observed in the data. We find that the estimated model can explain almost all the variation in marriage rates across cohorts. This result provides additional support in favor of the mechanism we propose as an explanation for the patterns documented in this paper.

The paper proceeds as follows. In Section 2, we discuss related papers. In section 3, we describe the data sets used to derive the empirical results. Section 4 documents our reduced-form findings. In section 5, we develop the search model and derive the two theoretical results. In section 6, we test and estimated the search model. Section 7 concludes.

2 Existing Explanations

In this section, we will describe the existing explanations for the variation in the U.S. marriage rates. The discussion will emphasize that the existing explanations for which empirical evidence is provided only apply to particular periods or to a subset of individuals. As a consequence, they cannot account systematically for the historical variation in U.S. marriage rates.

One set of explanations for changes in marriage rates over time focuses on changes in income, as reviewed in Cherlin (1981). These studies emphasize the correspondence between rising incomes during a period of postwar prosperity and the associated marriage boom after World War II. However, a positive relationship between income and marriage rates has not been successfully tested over different periods. A related explanation about income has been advanced by several researchers to account for the marriage decline during the Great Depression. It is argued that the decline in income was the main factor behind the reduction in marriage rates during this period. There is evidence, however, that this hypothesis can be rejected if one investigates a longer time horizon. For instance, Wolfers (2010) looks at the relationship between marriage rates and recessions for the past 150 years and finds no clear pattern between marriage and periods of economic decline. This suggests that an explanation about income or recessions cannot be used as a general theory that can account for the overall variation in marriage rates. In addition, there is a reverse causality problem with this theory that is not addressed in the studies reviewed by Cherlin (1981). It is well documented that married individuals have higher income. It is therefore difficult to determine whether an increase in income causes a rise in the marriage rate or whether an increase in the marriage rate generates higher income levels.

An alternative explanation for the variation in household formation is proposed by Akerlof, Yellen, and Katz (1996). They suggest that one can account for the marriage decline in the seventies using the adoption of new fertility technologies such as the pill and abortion. A potential direct effect could be that the adoption of the new technologies mechanically reduced the number of shut-gun marriages. Akerlof, Yellen, and Katz argue that the adoption of these new methods had also an indirect effect on shotgun marriages. The indirect effect was that the new technologies generated a decline in the competitive position of women relative to men. Under the assumption that a larger fraction of women wish to marry relative to men, this event should reduce marriage rates. It is immediately evident that this explanation only applies to the seventies when the new fertility technologies were introduced. It cannot account for the large variation in marriage rates observed before and after the seventies.

Greenwood and coauthors have suggested that the decline in the price of appliances explains patterns for several household outcomes and marriage is one of them. In particular, Greenwood and Guner (2009) argue that labor-saving technological progress in the household sector can explain the decline in marriage rates observed during some periods in the past century. Their idea is that the technological progress in the household sector makes it easier for singles to maintain their own home, which increases the value of being single and therefore reduces the marriage rate. This theory, however, cannot be used to systematically explain the historical variation in marriage rates. Since technological progress has constantly improved household appliances in the past 100 years, the proposed theory predicts that marriage rates should decline throughout this period. But this is not the case in the data. We will document later in the paper that marriage rates experience large fluctuating during this period, with times in which the fraction married increased sharply and times in which the fraction married declined at a fast rate.

Wilson (1987) proposes a theory of the variation in marriage rates which is based on fluctuations in labor market opportunities. Under the assumption that men are the primary earners, in his book, he argues that in periods with limited labor market opportunities there are fewer marriageable men. As a result, the marriage rate declines. Ellwood and Crane (1990) review papers that have tested Wilson's hypothesis and find conflicting evidence. Some papers provide evidence in support of this theory. But there are papers that reject this hypothesis. A couple of examples are Plotnick (1990) and Lerman (1989).

Some papers point at welfare programs and incarceration rates as possible explanations of the changes in the fraction married. In their review, Ellwood and Crane (1990) also evaluate papers that test the link between welfare aid and marriage and they conclude that there is very little empirical support for the proposition that welfare benefits played a major role in the marriage trends in blacks. In particular, they argue that the time-series patterns in welfare benefits are inconsistent with the hypothesis that higher welfare benefits accounted for family changes. Moreover, this explanation applies only to the black population for the majority of the last 100 years since, during those years, the fraction of white families making use of welfare programs was relatively small. Charles and Luoh (2010) study the relationship between incarceration and marriage rates and they find that higher incarceration rates decrease the fraction married in the black population. This explanation, however, can only explain the fluctuations in marriage rates after 1980 when large scale incarceration in particular of black men began.

The explanation that has some communality with the one we propose is Easterlin's hypothesis. Easterlin (1987) argues that the relative size of a cohort can explain all the variables that determine the economic and social fortunes of a birth cohort: earnings and unemployment rates, college enrolment rates, divorce, fertility, crime, suicide rates, and marriage. The idea behind this claim is that when income is above the aspiration level for a given cohort, the individuals in a given cohort will be optimistic and therefore will have better economic and social outcomes. If the distribution of income of a cohort is affected by its size, then the size will affect its economic fortunes. Easterlin, however, has provided only indirect evidence in support of his hypothesis and researchers that have attempted to test the general idea behind it have found mixed results. For instance, Pampel and Peters (1995) review papers that have investigated the mechanism outlined by Easterlin and conclude that "the evidence for the Easterlin effect proves mixed at best". More importantly, Easterlin provides no direct evidence on the link between cohort and marriage and on the mechanism behind it, which is the main objective of this paper.

3 Data

We use the total births recorded in the Vital Statistics of the United States (1909-2011) to measure cohort size, nationally and by state for white and black populations. For cohorts born before 1960, exact classification by race is not available. Therefore, our series for black cohort size includes all non-white individuals for these pre-1960 cohorts. This is not a great limitation, as more than 95 percent of non-white individuals during this time were black. While in-migration, out-migration, and mortality before age 30 affects the final number of marriageable-age adults in each cohort, cohort size provides a very good approximation for this number. Since the time frame for the analysis begins after the end of large migration waves at the end of the nineteenth century and beginning of the twentieth century, for most of the period that we consider the annual immigration rates are below 1 percent of the U.S. population.

We construct our measures of the share never married using the IPUMS Current Population Survey (1962-2011). Our most commonly used measure, the share never married by 30, can be constructed from the CPS for all cohorts born after 1930. For younger cohorts, we use the IPUMS Census 1960-1980, which provides information on the age at first marriage for all ever-married men and women. We record the share of all individuals between ages 30 and 45 who are married and report an age at first marriage of 30 or younger. We use a cutoff age of 45 to avoid potential measurement errors due to differential mortality rates of married and nonmarried persons. For the state-level cross-sectional results, we use only the IPUMS Census.

Finally, for the mobilization rates used in the instrumental variables regressions we use data made available by Acemoglu, Autor, and Lyle (2004) from the Selective Service System (1956) monographs.

4 Empirical Results

4.1 Change in Marriage Rates Over Time

As a first contribution, we show that the variables used in the literature to study the evolution of the marriage market are problematic. The U.S. marriage market saw significant change over the course of the twentieth century, as documented in Figures 1 and 2 using the two measures most commonly employed in the literature. Figure 1 records the number of marriages per 1,000 individuals and per 1,000 unmarried women. The cross-sectional measure in Figure 2 in turn records the share of women within an age range, e.g., between the ages of 18 and 30, who ever married.

The two measures portray somewhat different patterns, and both suffer from important drawbacks when they are used to study the evolution of marriage choices over time. The most serious problem with the first measure, the number of marriages normalized by population, is that it conflates changes in the numerator (marriages) with changes in the denominator (population). During periods of population growth or decline, e.g., due to large-scale demographic changes in fertility or migration, one may draw the wrong inference about the changes in marriage rates. The variable in Figure 2, in turn, conflates a different set of effects, the change in the number of people who will ever marry, and the age at first marriage. If the age at first marriage changes over time, one may similarly draw the wrong inference about variation in marriage rates.

To employ a consistent and comparable measure of marriage behavior over time, we calculate the share of men or women in a particular cohort ever married by a specific age (e.g., 30, 40, etc.). To compare the performance of this and the previous measures, Figure 3 plots marriages per 1,000 people, the cross-sectional measure (share of women ages 18-30 ever married in a given year), and our measure, the share of women in a cohort married by age 30. We shift the measure forward in time (year of birth plus 25 years) to make it comparable with the other two measures, which are based on the calendar year.

As one may expect, the cross-sectional and cohort-based measures evolve fairly similarly, except after the mid-1980s. While the cross-sectional measure would suggest a sustained drop in marriage rates until the present, the cohort-based measure documents an actual increase in the share ever married by age 30 after 1985, even as the marriage age continued to increase. The cohort-based measure contrasts even more dramatically with the number of marriages per 1,000 people. The observed drop after the mid-1940s and the recovery in the mid-1960s for this measure are accounted for primarily by the changes in overall population associated with the start and end of the baby boom (1945-1960). By contrast, the cross-sectional and cohort-based measures show a sustained increase in marriage rates until the mid-1960s, after which they begin to flatten and drop somewhat. The cohort-based measure shows a subsequent steeper drop beginning in the early 1970s and, looking back again at birth years, suggests that this sustained drop occurs for the first cohorts of the baby boom generation.

Figure 4 plots our measure, the share of men and of women never married by age 30, by cohort, against cohort size. The figure documents not only the drop in marriage rates starting with the baby boom, but in fact shows a striking co-movement between a cohort's size and the share never married by age 30 in that cohort over virtually the entire period. The decrease in the share never married for the small cohorts born in the 1920s and 1930s corresponds to the well-documented marriage boom that starts in the mid-1940s and lasts through the early 1960s. A sharp rise in the share never married begins in 1946 with the first of the post-war baby boomer generations. Between 1945 and 1955, total births for whites increased from around two million to nearly three and a half million, with a population increase of similar scale for blacks. The share never married by age 30 tripled over this period (more than tripled for blacks), until cohort size stabilized and began to drop in the 1960s. Notably, black and white marriage behaviors follow exactly the same pattern, contrasting with previous work that has emphasized fundamental differences across race in marriage behaviors.

Starting with cohorts born in the 1970s, the relationship in Figure 4 between cohort size and the share never married begins to weaken, for both black and white population. Figure 5 plots the same relationship again, and accounts additionally for the increase in cohabiting households during this period. Remarkably, once cohabitation is taken account of, the relationship between cohort size and household formation resembles again that of pre-1970 cohorts. Falling cohort sizes in the 1960s and 1970s correspond to a fall in the share never married (and not cohabiting) by 30, and the measure responds even to medium-run fluctuations in cohort size. Using a different cutoff age (e.g., share never married by 40) replicates the same pattern (Figure 6).

As Table 1 documents, the relationship between cohort size and the share never married is tightly estimated, and the R-squared of regressions of the log share never married and not cohabiting by age 30 or 40 in a cohort on the log of the cohort's size, without any other controls, ranges between 63 percent and 86 percent. The elasticity of marriage and household formation rates by 30 in response to cohort size growth ranges between -0.23 and -0.96, and is highest for black women. The coefficients are smaller for a later cutoff age of 40 (columns (5) and (6)). Increasing cohort size by 10 percent, in other words, decreases the share ever married by 30 on average by about 2.3 to 9.6 percent, a very strong effect. In percentage points, this amounts to an approximately 2.0 to 6.6 percentage point decrease on average in the share of individuals ever married.

Because cohabitation has become an important pre-cursor to or close substitute for marriage since the early 1990s, throughout the rest of the study, we will use this adjusted measure of household formation, the share never married and not cohabiting by a given age. Unless specifically noted otherwise, henceforth when we use the shorthand "ever married" we refer to those ever married or cohabiting.

4.2 Change in Marriage Rates Across States

If differences in the growth of cohorts causally generate variation in marriage or household formation rates, we should observe such an effect not just across time but also across geography. For example, one test of a causal relationship is that states with large increases in cohort sizes in the relevant marriageable-age groups should have lower marriage rates relative to other states, and vice versa. In general, making inferences from a simple cross-sectional comparison can be problematic, in particular because population growth of marriageable-age individuals will in part be driven by migration, which is not exogenous to economic and social conditions and which will also affect marriage rates. To exacerbate the problem, migration can be sex-biased, which can skew sex ratios and affect marriage rates directly.

To avoid these potential sources of endogeneity, to conduct a cross-sectional test we use total births by state in a given year to approximate changes in the cohorts of adults in the marriage market, rather than using a measure of the number of adults in a particular cohort and state directly. Specifically, for the decennial years between 1940 and 1970, we record two variables: the total births in a state, by race (from the U.S. Vital Statistics); and our measure of interest, the share ever married in those cohorts by state thirty years later. We assign individuals in the Census to a particular state based on their birthplace. We end the analysis with 1970 cohorts because they are the oldest cohorts for which we have data (they are of age 30 in the 2000 Census). We start with the 1940 cohort because it is the first year for which we have birth data for whites and non-whites without a substantial number of missing data for a large number of states. Finally, for states in which we record total births of less than 500 individuals (this is true mainly for blacks in states that are predominantly white), we drop the state from the analysis, since the outcome variable (share of men or women not married by age 30) is estimated on a very small sample. Hawaii and Alaska are excluded as data is not available for those states in 1940 and 1950.

Table 2 presents the results of the test. We regress the ten-year log difference in the ever married for thirty-year-olds in a particular state on the ten-year growth rate of those cohorts. Because of the small sample size, we pool the cross-sections and control for general time trends with year fixed effects. We report the results separately by race and gender. Without any other controls, the size of the coefficient is between -0.55 and -0.74 for men and slightly lower for women, between -0.29 and -0.35. All results are significant and fall within the range of the longitudinal results (-0.23 to -0.96). Evaluated at the mean, a 10 percent increase in cohort size over 10 years reduces the share ever married by 2.9 to 7.4 percentage points. The results in the individual, unpooled cross-sections yield R-squared values ranging between 13 and 60 percent (see, for example, columns (5) and (6)).

4.3 Potential Endogeneity Concerns

The test using cross-state regressions strongly corroborates the relationship between changes in cohort size and changes in marriage rates. However, there may be remaining reasons that prevent a causal interpretation of the relationship. While using births by state as our independent variable allows us to reasonably avoid reverse causality problems as well important endogeneity concerns due to migration, one may nevertheless worry that certain state-level characteristics could drive both changes in birth rates for cohorts born 30 to 40 years ago, as well changes in current marriage decisions. Such an omitted variable would have to be a highly persistent shock generating a differential trend over time. In particular, it would have to affect growth in births as well as growth in subsequent marriage rates about thirty years later.

It is not easy to think of variables that fit this description. Potential examples could be highly persistent productivity shocks which cause wages to grow more rapidly in some states over time, affecting both birth rates at the time of the initial shock and marriage rates two or three decades later. For example, rising men's incomes in some states could both increase births in 1950 relative to 1940 (if children are a normal good) in those states, as well as make men more eligible husbands in 1980 relative to 1970. Alternatively, improving fertility technologies may have a differential effect in states that are strongly religious or have a preference for family compared to states that do not. One might again expect depressed birth rates in states with lower preferences for family, and potentially lower marriage rates in the future as well.

Note that in these and most tenable cases we would typically expect an increase (or decrease) in *both* births and subsequent marriage rates, in as far as fertility and marriage are typically positively associated. This bias works against our favor and would result in a coefficient of the opposite sign. In fact, cross-sectionally we observe that increasing births generate decreasing marriage rates 30 years later, consistent with the time series evidence. Nevertheless, without exogenous variation in cohort size we cannot entirely eliminate the possibility that some biases could work in our favor.

4.4 Instrumental Variables Strategy

To isolate a plausibly exogenous source of variation in births, we employ an idea proposed originally by Acemoglu, Autor, and Lyle (2004) and exploit differences in mobilization rates across states during World War II. Mobilization rates between states differed substantially during the war, ranging overall from approximately 41.2 percent to 54.5 percent of young men ages 18-44. From 1940 to 1947, over 10 million men were inducted into the military, with the largest bulk of American soldiers deployed starting in 1943 and 1944, most for the duration of the war. As we show below, this introduced substantial variation in birth rates across states as prime-age men deployed for the war effort left behind wives and potential partners.

In this section, we focus on white individuals only for two reasons. Firstly, black populations in the 1940s were not sufficiently large in many states outside the South to construct reliable marriage rates by cohort using Census samples. Secondly, mobilization rates for black men were very low as discussed in Acemoglu, Autor, and Lyle (2004). As a result, we use birth data from the National Vital Statistics for white individuals only by state from 1938, five years before large-scale mobilization, to 1950, five years after the end of the war. Similarly, our mobilization rate measure is constructed for white individuals only. Our data from the Selective Service's Special Monograph (1956) for men ages 18-44 provides racial breakdowns of draft registrations and inductions, but not of enlistments. Therefore, we construct the mobilization rate measure as the number of white men ages 18-44 inducted into the armed forces, divided by the number of white men ages 18-44 registered. The latter includes virtually all men ages 18-44 between 1940 and 1945. A benefit of constructing the mobilization measure in this particular way is that the key source of variation in participation in the armed forces stems from randomness generated by the differences in deferral processes of draft boards across states leading to inductions, rather than by differences across states in men's voluntary enlistment rates. For this reason, the mobilization rates we use are on average lower than those constructed by Acemoglu, Autor, and Lyle (2004).

Figures 8, 9, and 10 illustrate that the effect of higher mobilization rates on births can be divided into three parts. First, during the first period of mobilization, fathers were granted a delay in deployment. As a consequence, during the first phase of World War II, we would expect higher birth rates for men with higher probabilities of being deployed and consequently for states with a larger fraction of them. Figure 8 describes this "anticipation effect" for the years 1940 and 1942. As predicted, there is a positive correlation between the mobilization rates and changes in log births, with a regression coefficient of about 0.8.[1] Second, during the second phase of the war when the majority of U.S. men were deployed, in states with higher rates of mobilization, the absence of prime-age men should be associated with a larger drop in births. We will refer to this second effect as the "incapacitation effect". Figure 9 shows the correlation between mobilization rates and changes in log births from 1943 to 1945. As expected, the correlation is negative. In states with high mobilization, the greater absence rate of potential partners leads to a greater drop in births. The coefficient on the regression line suggests that a one percentage point increase in the mobilization rate is associated with 1.26 percent lower cohort growth. Finally, we would expect an increase in birth rates when the deployed men who survived the war returned home and started to make up for the missing years. We will refer to this effect as the "return effect". Figure 10 plots the correlation between mobilization rates and changes in log births for 1945 and 1946. As expected, the post-war recovery in births in states with high mobilization rates is greater than the recovery in low-mobilization states. A one percentage point increase in mobilization rates is associated with a 0.58 percent increase in the growth in births.

We will use this source of arguably exogenous variation to identify a causal effect on wartime-induced differential changes in cohort size across states. Specifically, we will estimate in a second stage a model similar to the one we introduced in the previous subsections:

$$y_{c,s} = \sum_{s} \pi_s + \sum_{c} \gamma_c + \phi \cdot logcohortsize_{c,s} + X_{c,s}\beta + \epsilon_{c,s} \tag{1}$$

where $y_{c,s}$ is the outcome variable of interest, the share of individuals never married who were born in year c and state s, and ϕ is our coefficient of interest is. Additionally, π_s is a set of

[1]To show that the anticipation effect is not a cross-state difference that existed before the war and is therefore independent of mobilization, in Figure 7 we report the same correlation for the years 1935 to 1940. Our results indicate that before the war the correlation between mobilization rates and changes in log births, while insignificant, was of the opposite sign than what the anticipation effect predicts.

state fixed effects; γ_c is a set of birthyear fixed effects, which are also allowed to vary by region in some specifications; $X_{s,t}\beta$ is an additional set of 1940 state demographic characteristics (e.g., educational attainment) interacted with time that allows for potentially different time trends in states with different baseline characteristics. In the first stage, we then instrument log cohort size using the interactions between a state's mobilization rate and year dummies, to allow for a flexible effect of the mobilization rate in different years:

$$logcohortsize_{c,s} = \sum_s \pi_s + \sum_c \gamma_c + \sum_c \gamma_c \cdot mobilizationrate_s + X_{c,s}\beta + \nu_{c,s} \tag{2}$$

Reasons for differences in mobilization rates across states are discussed extensively in Acemoglu, Autor, and Lyle (2004). To summarize, they include exemptions for farmers, as food production was a priority for the U.S. at this time; differences in ethnic composition, in particular concentration of Germans and Japanese in a state; differences in age; differences in occupational structures, where workers employed in industries central to the war effort were more likely to be exempted; and idiosyncratic differences in the behavior of local draft boards. Table 3 groups states into three groups, based on armed forces induction rates that vary from 26.4 to 35.9 percent: low-mobilization (with induction rates below 28.8 percent), medium-mobilization (induction rates between 28.8 and 31.2 percent), and high-mobilization states (induction rates greater than 31.2 percent). The results in Table 3 suggest that states in these three groups were not significantly different along all 1940 measures which include income, employment levels, family size, and share never married by 30, with one exception. The difference in mean years of education between high- and low-mobilization states is statistically significant at the five-percent level, with low-induction rate states recording about 0.7 more years of education.

Because we can control for state fixed effects, differences in the levels of state-specific baseline characteristics between high- and low-mobilization states, like income or education, will not on their own affect our estimates. However, it is possible that states may also have differing trends in these independent variables. To address this concern, we interact a set of baseline observables that are potentially correlated with the scale of mobilization with functions of time (up to fourth-order polynomials). We present the results with such time controls for two key 1940 baseline variables, education and income. This allows us to control flexibly for potential differences in the time paths of these variables.

The data we use to run the regressions of interest is for cohorts born between 1940, three years before the large-scale mobilization began in 1943, and 1948, three years after the end of the war. We construct the share ever married by 30 for each state and cohort by using the IPUMS Census 1970 1% and 1980 5% samples. For individuals above age 30, the reported state of birth, marital status, and age at first marriage allows us to construct our outcome marriage measure by state. Because the age at first marriage variable is not available in subsequent Censuses, the 1950 cohort is the most recent cohort for which we can construct the share ever married by 30.

Table 4 reports the outcome of the first-stage regression. We consider two different specifications. In both of them, we allow for different time trends based on baseline state characteristics. In the specification reported in column (1) we also control for year fixed effects. To control for the possibility that part of the cross-state variation in mobilization rates is generated by underlying differences across regions, in the second specification we add region fixed effects. This specification is the most conservative since we only use variation in mobilization rates across states that belong to the same region. If differences across regions in mobilization rates are an important source of exogenous variation, the effect of changes in mobilization rates on birth rates will be underestimated in the second specification.

In table 4, we report the coefficients on the interactions between the mobilization rate in a particular state and a dummy for each year from 1940 to 1948. We start by describing the results obtained using the first specification. Our estimates indicate that, after controlling for underlying differences across states, only the anticipation and incapacitation effect are present in the data. A positive and statistically significant coefficient estimate on the mobilization rate for 1942 suggests that men attempted to delay their deployment by becoming fathers. As expected, the coefficient estimate becomes zero for 1943 and turns negative and statistically significant for 1944 and 1945. The deployment of prime-age men had the effect of reducing the birth rate. We do not observe the return effect. The coefficients on the mobilization rate after the end of the war are all negative and statistically insignificant. This results can be justified by the fact that high-mobilization states lost a larger portion of their young men to casualties. As a consequence, the return effect may be counteracted by a reduction in the number of prime-age men. In the second, more conservative specification, where we control for region fixed effects, only the anticipation effect remains statistically significant, with coefficients for 1942 and 1943 that are positive and large. The coefficients for 1944 and 1945 are now statistically equal to zero. This finding indicates that the incapacitation effect is mostly the result of differences across regions.

Table 5 reports the results of the second stage regression of log share ever married by 30 on log cohort size, for both specifications. The effect we measure is negative and strongly significant. We find that a one percent increase in cohort size generates a 0.041% to 0.055% reduction in the share of women ever married and a 0.061% to 0.100% decrease in the share of men ever married by 30. The small size of the measured effect is not surprising. Even though the reduction in births due to mobilization measured in our first stage regression is potentially large during the key years of the war - according to our previous calculations, a 10 percentage point increase in mobilization reduces cohort size by more than 10 percent between 1943 and 1945 - the period during which cohort size could be affected during the war is very short, potentially only three or four years. Changes in cohort size over a longer time horizon will generally have stronger effects that cumulate over time. The difference between the short and long run effects should account for the larger effects measured in the previous section. Nevertheless, we believe that the fact that such a short-term reduction in cohort size has a noticeable and strongly significant effect provides compelling evidence for the causal effect of changes in cohort size on changes in marriage patterns.

5 A Dynamic Search Model of the Marriage Market

We start with a simply dynamic search model of the marriage market and evaluate whether it can generate the patterns observed in the data. As an alternative, one could have used a matching model of the type employed by Gale and Shapley (1962), Becker (1973), Becker (1974), Mortensen (1988), Bergstrom and Bagnoli (1993), Peters and Siow (2002), Choo and Siow (2006), Chiappori, Iyigun, and Weiss (2009), Hitsch (2010), and Iyigun and Walsh (2007). We choose the search model because it allows us to better capture the dynamic nature of the marriage market which is essential to explain the link between cohort size and marriage rates.

The economy is populated by $T + 1$ overlapping generations of men and women. In each period t, a new generation is born and lives for $T + 1$ periods. Men and women can be either single or married. If an individual is married she or he makes no choice. If in period t an individual of gender i and age a is single, she or he meets a potential spouse with probability $\theta^i_{a,t}$. The two spouses then decide whether to marry with the objective to maximize their lifetime utility.

To capture the fact that women are fertile for only part of their lives whereas man are fertile for most of their life, we will assume that women meet a man with a positive probability only in their first period of life while men meet a potential spouse with a positive probability in their first two periods of life. This is equivalent to assuming that the only reason for marriage is children, that after the first period of their life women become infertile, and that men cannot have children after two periods. This assumption has two implications. First, the marriage market is populated by women of age 0 and by men of age 0 and 1. Second, women cannot change their marital status after the first period and men cannot change it after the second period. Allowing women to marry for more than one period and men for more than two periods makes the model more complicated without changing the main results.

The within-period utility of being single will be denoted by δ, whereas the within-period utility of being married for the couple will be denoted by η. The utility from future periods is discounted at the discount factor $\beta \leq 1$. We will assume that the value of being single is constant across individuals and over time. As a consequence, if an older man chooses to be single in the second period, his lifetime utility takes the following form:

$$\sum_{t=0}^{T-1} \beta^t \delta = \frac{1-\beta^T}{1-\beta}\delta.$$

Similarly, if a woman decides to stay single in her first period of life, her lifetime welfare can be computed as follows:

$$\sum_{t=0}^{T} \beta^t \delta = \frac{1-\beta^{T+1}}{1-\beta}\delta.$$

The value of being married is drawn from a distribution $F(\eta)$ which does not depend on the age of the couple or on time. If two potential partners decide to marry, the within-period utility they have drawn is also the utility they will experience in each period for the rest of their life. The lifetime utility of a couple of individuals who are both of age 0 and have drawn a value η in period t can therefore be written as follows:

$$v_{0,0,t} = \sum_{\tau=0}^{T} \beta^\tau \eta = \frac{1-\beta^{T+1}}{1-\beta}\eta.$$

If the couple is composed of a woman and on older men, the lifetime utility takes the following form:

$$v_{0,1,t} = \sum_{\tau=0}^{T-1} \beta^\tau \eta + \beta^T \delta = \frac{1-\beta^T}{1-\beta}\eta + \beta^T \delta.$$

We will assume that the couple can freely divide the gains from marriage and that its lifetime utility is split between the two spouses using a Nash bargaining solution. For a couple composed of a woman of age 0 and a man of age 1, the share received by the man in period t is

$$w_{1,t}^m(\eta) = v_{1,t}^m + \gamma\left[v_{0,1,t} - v_{1,t}^m - v_{0,t}^w\right] = v_{1,t}^m + \gamma_m\left[\frac{1-\beta^T}{1-\beta}\eta + \beta^T\delta - v_{1,t}^m - v_{0,t}^w\right], \qquad (3)$$

where the parameter $\gamma \in [0,1]$ captures the bargaining power of the man and $v_{1,t}^m$ and $v_{0,t}^w$ are the value of being single in this period and of re-optimizing from next period onward for the man and the woman. A similar equation can be derived for the woman. To make it harder for our model to explain the variation observed in the data, we will assume that γ is independent of market conditions.

We will now solve the model starting with the decisions of a man of age 1 in period t. With probability $\theta^m_{1,t}$, he meets a woman of age 0 and they marry if their joint lifetime utility from being married is greater than the sum of their lifetime utilities if they choose to stay single. To determine the match quality η above which the couple will choose to marry, observe that if a man of age 1 and a woman of age 0 decide to remain single, they will be single for the rest of their life. As a consequence, they will marry if and only if

$$\eta\frac{1-\beta^T}{1-\beta}+\delta\beta^T \geq \delta\frac{1-\beta^T}{1-\beta}+\delta\frac{1-\beta^{T+1}}{1-\beta}=2\delta\frac{1-\beta^T}{1-\beta}+\delta\beta^T.$$

This implies that the reservation value for marriage for a man of age 1 is

$$\underline{\eta}_{1,t}=2\delta.$$

We can now derive the expected value function for an older man before he enters the marriage market. If in period t this man meets a woman and draws a value η, Nash-bargaining implies that he receives the following share of the couple's lifetime utility:

$$w^m_{1,t}(\eta)=\delta\frac{1-\beta^T}{1-\beta}+\gamma\left[\eta\frac{1-\beta^T}{1-\beta}+\delta\beta^T-2\delta\frac{1-\beta^T}{1-\beta}-\delta\beta^T\right]=\left[\delta+\gamma\left(\eta-2\delta\right)\right]\frac{1-\beta^T}{1-\beta}.$$

As a consequence, the expected value function of an older man can be written in the following form:

$$v^m_{1,t}=E\left[\delta+\gamma\left(\eta-2\delta\right)|\eta\geq\underline{\eta}_1\right]\frac{1-\beta^T}{1-\beta}\left(1-F\left(\underline{\eta}_1\right)\right)\theta^m_{1,t}+\delta\frac{1-\beta^T}{1-\beta}F\left(\underline{\eta}_1\right)\theta^m_{1,t}+\delta\frac{1-\beta^T}{1-\beta}\left(1-\theta^m_{1,t}\right)$$

It is composed of three parts. The first part describes the value for the older man of meeting a woman with a match quality η sufficiently high that the couple will choose to marry. The second part characterizes the value of meeting a woman with a match quality η that is below the reservation value $\underline{\eta}_1$. Finally, the last part captures the value of not meeting a woman in the current period.

By replacing $\underline{\eta}_1 = 2\delta$, by dividing both sides of the equation by $\dfrac{1-\beta^T}{1-\beta}$, and by simplifying some of the terms, we obtain the following equation for the value function:

$$v_{1,t}^m \frac{1-\beta}{1-\beta^T} = \gamma \left\{ E\left[\eta \left| \eta \geq 2\delta \right.\right] - 2\delta \right\} \left(1 - F\left(2\delta\right)\right) \theta_{1,t}^m + \delta.$$

We are now in the position to consider the decision of a young man. This individual meets a potential spouse with probability $\theta_{0,t}^m$ and they marry if their joint lifetime utility is greater than the joint lifetime utility if they choose to be single in this period, i.e. if

$$\eta \frac{1-\beta^{T+1}}{1-\beta} \geq 2\delta + \beta v_{1,t}^m + \beta\delta \frac{1-\beta^T}{1-\beta},$$

where the first term on the right hand side is the joint value of being single in this period, the second term is the man's discounted expected value function for next period if he chooses to stay single today, and the third term is the woman's discounted value from next period onward if she chooses to stay single today. The reservation value for a man of age 0 can therefore be written as follows:

$$\underline{\eta}_{0,t} = 2\delta \frac{1-\beta}{1-\beta^{T+1}} + \beta v_{1,t}^m \frac{1-\beta^T}{1-\beta^{T+1}} + \beta\delta \frac{1-\beta^T}{1-\beta^{T+1}}$$

We can now substitute for the expected value function of an older man and simplify some of the terms to obtain the following equation for the reservation value of a young man:

$$\underline{\eta}_{0,t} = 2\delta + \beta \frac{1-\beta^T}{1-\beta^{T+1}} \gamma \left\{ E\left[\eta \left| \eta \geq 2\delta \right.\right] - 2\delta \right\} \left(1 - F\left(2\delta\right)\right) \theta_{1,t}^m. \tag{4}$$

Using the reservation value of a young man, we can now derive the expected value for a young man in period t. It takes the following form

$$\begin{aligned} v_{0,t}^{m} &= \theta_{0,t}^{m}\left(1-F\left(\underline{\eta}_{1,t}\right)\right)\left\{\delta+\beta v_{1,t}^{m}+\gamma\left\{\frac{1-\beta^{T+1}}{1-\beta}E\left[\eta\left|\eta\geq\underline{\eta}_{0,t}\right.\right]-\left(\delta+\beta v_{1,t}^{m}\right)-\frac{1-\beta^{T+1}}{1-\beta}\delta\right\}\right\} \\ &+ \theta_{0,t}^{m}F\left(\underline{\eta}_{0,t}\right)\left(\delta+\beta v_{1,t}^{m}\right)+\left(1-\theta_{0,t}^{m}\right)\left(\delta+\beta v_{1,t}^{m}\right) \end{aligned}$$

where the first term represents the value of meeting a woman with a match quality η higher than the reservation value, the second term describes the value of meeting a women characterized by an η lower than the reservation value, and the third term measures the value of not meeting a woman when young.

To derive the woman's expected value function we have to take into account that she can meet both young and old men. As a consequence, it takes the following more complex form:

$$\begin{aligned} v_{0,t}^{w} &= \theta_{0,t}^{m}\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\left\{\frac{1-\beta^{T+1}}{1-\beta}\delta+(1-\gamma)\left\{\frac{1-\beta^{T+1}}{1-\beta}E\left[\eta\left|\eta\geq\underline{\eta}_{0,t}\right.\right]-\left(\delta+\beta v_{1,t}^{m}\right)-\frac{1-\beta^{T+1}}{1-\beta}\delta\right\}\right\} \\ &+ \theta_{0,t}^{m}F\left(\underline{\eta}_{0,t}\right)\frac{1-\beta^{T+1}}{1-\beta}\delta \\ &+ \theta_{1,t}^{m}\left(1-F\left(2\delta\right)\right)\left\{\frac{1-\beta^{T+1}}{1-\beta}\delta+(1-\gamma)\left\{\frac{1-\beta^{T}}{1-\beta}E\left[\eta\left|\eta\geq 2\delta\right.\right]+\beta^{T}\delta-v_{1,t}^{m}-\frac{1-\beta^{T+1}}{1-\beta}\delta\right\}\right\} \\ &+ \theta_{1,t}^{m}F\left(2\delta\right)\frac{1-\beta^{T+1}}{1-\beta}\delta \end{aligned}$$

In the next subsection, we will use the reservation values derived above to solve for the steady state equilibrium in the marriage market.

5.1 Steady State

In this subsection we determine the steady state equilibrium of our economy. We will use it to derive a couple of theoretical results and as an input for the structural estimation of the model which will be discussed in the next section.

To solve for the steady state equilibrium, we have to derive the probability that a younger man meets a woman $\theta_{0,t}^{m}$ and the corresponding probability for an older man $\theta_{1,t}^{m}$. Let N_{t}^{i} be the number of individuals of gender i and age t who are present in the marriage market. Then $\theta_{0,t}^{m}$

and $\theta^m_{1,t}$ can be derived by noting that

$$\theta^m_{0,t} = \theta^m_{1,t} = \frac{N^w_{0,t}}{N^m_{0,t} + N^m_{1,t}}. \tag{5}$$

The number of individuals of age 0 is exogenously given by the cohort size of a given generation. However, the number of older men in the marriage market $N^m_{1,t}$ is endogenously determined by the decisions of younger men. As a consequence, to derive $\theta^m_{1,t}$ we need to solve for $N^m_{1,t}$. This variable can be computed as the number of younger men who did not meet a woman at $t-1$ plus the number of younger men who met a woman at $t-1$ but draw a match quality η lower than the reservation value, i.e.

$$N^m_{1,t} = N^m_{0,t-1}\left(1-\theta^m_{0,t-1}\right) + N^m_{0,t-1}\theta^m_{0,t-1}F\left(\underline{\eta}_{0,t-1}\right) = N^m_{0,t-1}\left(1-\theta^m_{0,t-1}\left(1-F\left(\underline{\eta}_{0,t-1}\right)\right)\right). \tag{6}$$

We can now replace for $\theta^m_{0,t-1}$ using (5) and obtain the following equation for $N^m_{1,t}$:

$$\begin{aligned} N^m_{1,t} &= N^m_{0,t-1}\left(1-\frac{N^w_{0,t-1}}{N^m_{0,t-1}+N^m_{1,t-1}}\left(1-F\left(\underline{\eta}_{0,t-1}\right)\right)\right) \\ &= N^m_{0,t-1}\left(\frac{N^m_{0,t-1}+N^m_{1,t-1}-N^w_{0,t-1}\left(1-F\left(\underline{\eta}_{0,t-1}\right)\right)}{N^m_{0,t-1}+N^m_{1,t-1}}\right). \end{aligned}$$

In a steady state equilibrium, the cohort size $N^w_{0,t}$ and $N^m_{0,t}$ and the number of older men in the marriage market $N^m_{1,t}$ are constant over time. We therefore have that

$$N^m_1 = N^m_0\left(\frac{N^m_0+N^m_1-N^w_0\left(1-F\left(\underline{\eta}_{0,t-1}\right)\right)}{N^m_0+N^m_1}\right).$$

We can now solve for N^m_1 and obtain

$$N^m_1 = \sqrt{\left(N^m_0\right)^2 - N^m_0 N^w_0 + N^m_0 N^w_0 F\left(\underline{\eta}_0\right)}.$$

Observe that generally men and women have identical cohort size, $N_{1,t}^m = N_{1,t}^w = N_{1,t}$.[2] In this case the solution for N_1^m simplifies to

$$N_1^m = N_0 F\left(\underline{\eta}_0\right)^{\frac{1}{2}}.$$

If we substitute N_1^m back into θ_j^m, we have

$$\theta_0^m = \theta_1^m = \frac{N_0^w}{N_0^m + \sqrt{\left(N_0^m\right)^2 - N_0^m N_0^w + N_0^m N_0^w F\left(\underline{\eta}_0\right)}}.$$

If men and women have identical cohort size, θ_j^m simplifies to

$$\theta_0^m = \theta_1^m = \frac{N_0}{N_0 + N_0 F\left(\underline{\eta}_0\right)^{\frac{1}{2}}} = \frac{1}{1 + F\left(\underline{\eta}_0\right)^{\frac{1}{2}}}.$$

To determine the reservation value of younger men outside the steady state, we will assume that an individual uses the current meeting probabilities to compute the reservation value. We can therefore substitute for θ_1^m in the equation that determines the reservation value (4) and derive, for the case in which $N_0^m \neq N_0^w$, the following equation for the steady state reservation value:

$$\underline{\eta}_{ss} = 2\delta + \beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\left\{E\left[\eta\left|\eta \geq 2\delta\right.\right] - 2\delta\right\}\left(1 - F\left(2\delta\right)\right)\frac{N_0^w}{N_0^m + \sqrt{\left(N_0^m\right)^2 - N_0^m N_0^w + N_0^m N_0^w F\left(\underline{\eta}_{ss}\right)}},$$

If $N_0^m = N_0^w$, the equation simplifies as follows:

$$\underline{\eta}_{ss} = 2\delta + \beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\left\{E\left[\eta\left|\eta \geq 2\delta\right.\right] - 2\delta\right\}\left(1 - F\left(2\delta\right)\right)\frac{1}{1 + F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}}.$$

Observe that $F\left(\underline{\eta}\right)$ is monotonically increasing in $\underline{\eta}$. As a consequence, there is a unique

[2] This is not the case if men or women are more likely not to be in the marriage market for particular reasons. For instance, African-American men are more likely than African-American women to be in jail during their marriage years. As a consequence, the relevant cohort size for African-American men is smaller than the corresponding cohort size for women.

solution for $\underline{\eta}_{ss}$. Moreover, if men and women have identical cohort sizes, the steady state reservation value is independent of N_1^m and N_1^w. The following Proposition summarizes the result.

Proposition 1 *In steady state, there is a unique reservation value for marriage $\underline{\eta}_{ss}$. It does not depend on cohort size if $N_0^m = N_0^w$.*

5.2 An Unexpected Shock to Cohort size

We will now consider the effect of a shock to cohort size on the fraction of individuals that choose to marry. We will focus on the case in which the shock is unexpected. Similar results apply if the shock is known with certainty. We will show two results. The first result is that a positive shock to cohort reduces the fraction of women in a given cohort who choose to marry. The second result provides a testable implication for the model and it establishes that an increase in cohort size reduces the average age difference between spouses. In the remaining part of the section we will consider the case in which $N_0^m = N_0^w = N_0$. As a consequence, the results do not apply to African-Americans since for this population the jail rate and mortality rate are higher for men of marriage age than for women. Hence, for this marriage market $N_0^m < N_0^w$.

Suppose the economy in steady state when it is hit by an unexpected shock in period $t = \tau$ that changes permanently the women's cohort size from N_1^w to $N_1^w + \Delta$ and the men's cohort size from N_0^m to $N_0^m + \Delta$.

According to equation (5), the probabilities $\theta_{j,t}^m$ take the following form:

$$\theta_{0,t}^m = \theta_{1,t}^m = \frac{N_{0,t}}{N_{0,t} + N_{1,t}^m} \qquad \text{if } t < \tau$$

and

$$\theta_{0,t}^m = \theta_{1,t}^m = \frac{N_{0,t} + \Delta}{N_{0,t} + \Delta + N_{1,t}^m} \qquad \text{if } t \geq \tau.$$

Consider the period in which the shock is realized and notice that $N_{1,\tau}^m$ are the men born in period $\tau - 1$ who did not marry when young. As a consequence, $N_{1,\tau}^m$ equals the number of older men in steady state, i.e. $N_{0,\tau-1}F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}} = N_0 F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}$. Substituting for $N^m_{1,\tau}$ in the probabilities $\theta^m_{j,t}$, we have that in period τ

$$\theta^m_{0,\tau} = \theta^m_{1,\tau} = \frac{N_0+\Delta}{N_0+\Delta+N_0F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}} = \frac{1}{1+\frac{N_0}{N_0+\Delta}F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}}.$$

The previous equation implies that a positive cohort shock Δ increases the probability that a man of any age meets a woman, whereas a negative cohort shock has the opposite effect.

We can now determine the effect of a shock to cohort size on the reservation value of young men $\underline{\eta}_{0,\tau}$. Notice that in the determination of $\underline{\eta}_{0,\tau}$ a young men compares the value of getting married at τ with the value of waiting until the next period. The value of waiting depends on the probability that he will meet a woman in period $\tau+1$. This probability depends on the number of old men at $\tau+1$, $N_{1,\tau+1}$ and it can be written as follows:

$$\theta^m_{0,\tau} = \theta^m_{1,\tau} = \frac{N_0+\Delta}{N_0+\Delta+N_{1,\tau+1}}.$$

Using equation (6), we can substitute for $N_{1,\tau+1}$ to obtain the following expression:

$$\theta^m_{0,\tau} = \theta^m_{1,\tau} = \frac{N_0+\Delta}{N_0+\Delta+\left(N_0+\Delta\right)\left(1-\theta^m_{0,\tau}\left(1-F\left(\underline{\eta}_{0,\tau}\right)\right)\right)} = \frac{1}{1+\left(1-\theta^m_{0,\tau}\left(1-F\left(\underline{\eta}_{0,\tau}\right)\right)\right)}.$$

We can now substitute for $\theta^m_{1,\tau}$ in the equation that determines $\underline{\eta}_{0,\tau}$ to obtain

$$\underline{\eta}_{0,\tau} = 2\delta+\beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\left\{E\left[\eta\,|\eta\geq 2\delta\right]-2\delta\right\}\left(1-F\left(2\delta\right)\right)\frac{1}{1+\left(1-\theta^m_{0,\tau}\left(1-F\left(\underline{\eta}_{0,\tau}\right)\right)\right)}. \quad (7)$$

The same equation for the reservation value in steady state can be derived as follows:

$$\underline{\eta}_{0,ss} = 2\delta+\beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\left\{E\left[\eta\,|\eta\geq 2\delta\right]-2\delta\right\}\left(1-F\left(2\delta\right)\right)\frac{1}{1+\left(1-\theta^m_{0,ss}\left(1-F\left(\underline{\eta}_{0,ss}\right)\right)\right)}. \quad (8)$$

In the previous paragraph, we have shown that, with a positive shock to cohort size, $\theta^m_{0,\tau} > \theta^m_{0,ss}$.

As a consequence, a simple comparison of the last two equations implies that an increase in cohort size has the effect of increasing the reservation value of young men. This result is summarized in the following proposition.

Proposition 2 *A positive shock to cohort size in period τ increases the reservation value $\underline{\eta}_{0,\tau}$. A negative shock has the opposite effect.*

Proof. In the appendix. ■

It is left to show that a positive cohort shock and the consequent increase in the reservation value reduces the fraction of individuals in a given cohort who marry and that a negative shock has the opposite effect. The following Proposition proves this result.

Proposition 3 *A positive shock to cohort size in period τ reduces the fraction of cohort τ's women who get married by increasing the reservation value of marriage of young couples. A negative shock in period τ has the opposite effect.*

Proof. In the appendix. ■

To provide the insight behind this result, consider an increase in cohort size. After this event, older men become a scarce resource. This change has two effects. First, the fraction of women who marry mechanically declines because more women will not meet a potential spouse. Second, the fraction of women who marry decreases because young men become more selective since they will have a larger group of women to choose from when they will be old. The total impact of an increase in cohort size is therefore a reduction in the fraction of women who marry. This result indicates that the search model developed in this paper has the potential of explaining the relationship observed in the data between changes in cohort size and changes in marriage rates.

We will now derive an implication of the model that will be used in the next section as a test. It is based on an intuition which is similar to the one underlying Proposition 3. The two effects of an increase in cohort size mentioned above can be reworded in the following way. First, women are less likely to meet older men because there are relatively fewer of them. Since the reservation value of older men does not depend on cohort size, a consequence of this is that the average age difference between spouses should decline. Second, younger men become more selective in their decision to marry. As a result, fewer women will marry younger men and the average age difference between spouses should drop. As a whole, the increase in cohort size should therefore reduce the average age difference at marriage. The following Proposition shows that this intuition is correct.

Proposition 4 *In the model, an increase in cohort size reduces the average age difference between spouses. A reduction in cohort size has the opposite effect.*

Proof. In the appendix. ■

The result described in Proposition 4 will be used in the next section to test whether the model is consistent with the data.

6 Test and Estimation of the Search Model

This section will be divided into two parts. We will first test the model developed in the previous section using the result contained in Proposition 4. We will then estimate the model with the objective of evaluating whether the model can quantitatively match the variation in marriage rates observed in the data. In both parts we will only consider the population of whites because, as argued in the previous section, the assumption $N_0^m = N_0^w = N_0$ is not satisfied in the data for blacks.

6.1 A Test of the Search Model

Proposition 4 establishes an implication of the search model we consider in this paper. If the model is correct, an increase in cohort size should reduce the average age difference between spouse, whereas a reduction in cohort size should have the opposite effect. In this subsection, we will use this result to evaluate whether our model is consistent with the patterns observed in the data.

To implement the test, for each cohort we construct the variable age difference between spouses using the IPUMS CPS years 1960-2010 and the 1950 wage of the Census. Specifically, for each cohort we consider all women between the ages of 30 and 35 who are married. Then we compute the difference between their age and the age of their spouse. Finally, we calculate the average for each cohort. It should be remarked that the cohort size is measured at the time the individuals in a given cohort are born, whereas the age difference variables is computed three decades later when most of the women in that cohort had made their marriage decisions.

We perform the test in three stages. We first provide some evidence on the relationship between cohort size and average age difference by plotting the time series of these two variables. We then use the time-series variation to regress the logarithm of mean age difference between spouses on the logarithm of cohort size. Finally, we run the same regression using cross-state variation. We do not perform the test using the mobilization rates because we do not have enough variation to obtain precise estimates. Figure 11 describes two variables for the white population: the variation in age differences between the wife and husband across cohorts and the variation in cohort size. With the exception of the first ten cohorts, the Figure indicates that there is a tight relationship between age difference at marriage and cohort size. As the model predicts, when the size of a given cohort increases, the age difference between women in that cohort and their spouses becomes less negative and therefore decline. When the cohort size drops, the age difference between spouses becomes more negative and therefore increases.

Table 6 reports the coefficients and R-squared obtained by regression the logarithm of the average age difference between spouses on the logarithm of cohort size for whites using time-series variation. This regression enables us to determine whether the link between these two variables is statistically significant and how much of the variation in age difference at marriage is explained by cohort size. The estimated coefficient on cohort size is around -0.6 and strongly statistically significant. The estimated coefficients indicate that a 10% increase in cohort size generates a decline in age difference at marriage of approximately 6%. The size of the effect is therefore large. Finally, the R-squared suggests that cohort size can explain a significant fraction of the variation across cohorts in age difference between spouses. Our results indicate that about

81% of the variation in this variable can be explained by changes in cohort size. Finally, in the same Table we report the coefficient estimates obtained from regressing differences across states in the logarithm of age differences at marriage on differences across states in the logarithm of cohort size. At the state level, we observe the variables of interested in 1940, 1950, 1960, and 1970. We can therefore run three different regressions. The estimated coefficients are consistent with the search model. Except for the regression that uses differences between 1940 and 1950 for which our estimated coefficient is statistically equal to zero, our estimates are negative and statistically significant as required by our model. We therefore cannot reject the model developed in the previous section.

6.2 Estimation of the Search Model

In this subsection, we first estimate the dynamic search model developed in this paper. We then evaluate whether the estimated model can quantitatively explain the changes in marriage rates observed in the data in response to changes in cohort size. This exercise is important because it represents an additional and more thorough test of the mechanism behind the relationship between marriage rates and cohort size.

To structurally estimate the model, we have to make additional assumptions. The first assumption is about the distribution of the match quality η. We assume that it is distributed according to a beta distribution with shape parameters α_1 and α_2 defined on the interval $(0, 1)$. We have chosen the beta distribution for two reasons. First, it is of of the most flexible distributions. Evidence of this is that many popular distributions like the uniform, the exponential, and the gamma are a special cases of the beta distribution and that the normal distribution can be well approximated by it. The second reason is that the beta distribution is parsimonious with only two parameters to estimate.

A second assumption is required to be able to estimate the model. In the version developed in section 5, there is no source of uncertainty. To address this issue, we assume that the value of being single δ varies over time according to the following equation:

$$\delta_t = \delta + \nu_t,$$

where ν_t is drawn from a uniform distribution defined on the interval $[-0.2, 0.2]$.

The third set of assumptions we make are relative to the lifespan of an individual. Observe that individuals born in a given cohort marry over many years. Some of them find a spouse the first time they enter the marriage market, whereas others marry after having searched for many years. This implies that, in any given year, individuals from different cohorts compete in the marriage market. To model this feature, we assume that each period in our model corresponds to 10 years of an individual's life, that an individual starts making decisions at age 20, and that she or he lives for 50 years or, equivalently, 5 periods. To implement the assumption that each period corresponds to ten years, in a period we allow each individual to meet sequentially as many as ten potential spouses, one for each year. The individual leaves the marriage market when she or he marries one of the potential spouses. With this additional feature, an increases in cohort size that follows a previous rise will have a larger effect than a single increase, because the newcomers compete in a marriage market that is already crowded by the first rise. A similar argument applies to declines in cohort size. Notice that the reduced-form results cannot capture this aspect of the marriage market. The implicit assumption is that a change in cohort has the same effect on marriage rates independently of whether it is the first increase or a rise that follows a previous one.[3]

As additional assumptions, we set the annual discount factor equal to 0.98 and consider a symmetric Nash-barganing by setting γ equal to 0.5. Finally, we augment the model to allow for a fraction of men that are unwilling to marry no matter the value of match quality $1 - \phi$. This parameter only affects the probability that a younger or older man meets a woman. Specifically,

[3]We decided not to include this feature in the theory part because its addition does not change the insight provide by the model. The description of the model, however, would be more complicated.

these probabilities now take the following form:

$$\theta_{0,t}^{m} = \theta_{1,t}^{m} = \frac{N_{0,t}^{w}}{\left(N_{0,t}^{m} + N_{1,t}^{m}\right)\phi}.$$

The probability that a woman meets a younger or older man does not change because the parameter ϕ appears at the numerator as well as denominator.

Given these assumptions, the model has four parameters that must be estimated: the value of being single δ, which is assumed to be identical across gender and over time; the two shape parameters of the beta distribution α_1 and α_2; the fraction of individuals that are unwilling to marry ϕ. These parameters we be estimated using Simulated Method of Moments (McFadden (1989), Pakes and Pollard (1989), Lee and Ingram (1991), and Duffie and Singleton (1993)). Specifically, the estimation is performed in two steps. For a given set of parameters that characterize the model, we first simulate the individual decisions. We then match some of the statistical moments that characterize the data with the corresponding moments obtained using the simulated data. The estimated parameters are obtained by minimizing a function of the distance between the simulated and data moments required by indirect inference.

In the estimation, we use as our set of moments the fraction of women never married in a cohort starting from the cohort born in 1930 and ending with the cohort born in 1980. We therefore have 51 moments that will be matched using 5 parameters. Before presenting the results it is important to remark that these moments enable us to identify the value of being single δ and the parameter that determines the fraction of men who are unwilling to marry ϕ. To see this observe that the value of being single δ is linked to the fraction of individuals never married in a given cohort. Specifically, everything else equal, a higher value of δ increases the share of individuals who choose to stay single for each cohort and shifts up the curve that characterizes the fraction of never married as a function of cohort. The parameter ϕ enters equation (4) which defines the reservation value of young men. That equation can be viewed as a linear relationship between the reservation value of young men and the probability that when old they will meet a potential spouse. The slope of that equation is affected by ϕ: a larger value for ϕ generates a smaller slope. As a consequence, changes in the meeting probabilities will have smaller effects on the reservation value of younger men and therefore on marriage rates. In the model, changes in meeting probabilities are mainly generated by variations in cohort size. The parameter ϕ can, therefore, be identified by how the fraction never married varies in response to changes in cohort size. However, there is no reason to believe that the remaining two parameters can be identified using the selected set of moments. As consequence, the exercise performed in this subsection should be seen as a test of whether there exist parameter values for δ, α_1, α_2, and ϕ such that the dynamic search model can quantitatively explain the variation in marriage rates observed in the data. We believe that this exercise is the most interesting to perform, since in this paper we are not interested in performing policy evaluation or predictions.[4]

The estimated parameters are reported in Table 7. Our estimates of α_1 and α_2 are 0.020 and 0.072. These values imply that the mean of the distribution is equal to 0.217, the mode is equal to 0.512, and the standard deviation is equal to 0.036. The value of being single is estimated to be 0.107. This value indicates that only couples with relatively high match quality will choose to marry. To see this, remember that a woman and an older man marry only if $\eta > 2\delta$. Given our estimate of δ, this means that the couple will decide to marry only if the drawn match quality value is higher than 0.214. Since, this number is approximately equal to the mean, women matched to older men marry only if they draw a relatively high value for match quality. Finally, observe that younger men have a higher reservation value. Consequently, women paired with younger men will also marry only if their match quality is relatively high. Finally, our model rationalized the data by estimating that 13.3% of men are unwilling to marry independently of match quality.

We will now evaluate whether the model can quantitatively match the relationship between changes in cohort size and changes in marriage rates. We do this by plotting in Figure 12 the

[4]We have proven that the parameter α_1 can be identified using as a moment the probability that an older man marries plus the corresponding probability for a woman; α_2 can be identified using as a moment the probability that a younger man marries divided by the probability that a woman marries times the previous moment; the parameter ϕ can be identified using as a moment the probability that a woman does not marry divided by the sum of the probability that a young man does not marry and of the probability that an older man does not marry; the parameter γ can be identified using the fact that it changes the slope of the equation characterizing the reservation value. We will use these moments for the estimation of the parameters in future research.

marriage rates observed in the data jointly with the marriage rates simulated using the estimated model, both as a function of cohort. The Figure shows that the model can quantitatively replicate the variation in the share of never married across cohorts observed in the data. This outcome is particularly remarkable since our model is very parsimonious. We only have four parameters to match 51 moments.

7 Conclusions

In this paper we provide an explanation for the variation in U.S. marriage rates over the past century. Using time-series variation, cross-state variation, and variation in differences across states in mobilization rates during World War II we provide evidence in support of the following two results. First, cohort size can explain on its own almost all the variation in U.S. marriage rates. Second, an increase in cohort size reduces marriage rates and vice versa. We then develop a dynamic search model of the marriage market that has the potential of explaining the patterns observed in the data. Using the model, we first show that qualitatively it can generate the relationship between cohort size and marriage rates. We then derive the following testable implication: in the model an increase in cohort size reduces the age difference between spouses and vice versa. Finally, we test the model in two different ways. We first test whether the derived implication can be rejected. In the data, a rise in cohort size reduces the age difference at marriage and a decline in cohort size has the opposite effect. We therefore cannot reject our search model. We then estimate the model and evaluate whether it can quantitatively match the link between cohort size and marriage rates we document. The estimated model can easily match most of the variation in marriage rates observed in the data.

References

Acemoglu, Daron, David H. Autor, 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): pp. 497–551.

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): pp. 277–317.

Becker, Gary S. 1973. “A Theory of Marriage: Part I.” *Journal of Political Economy* 81 (4): pp. 813–846.

———. 1974. “A Theory of Marriage: Part II.” *Journal of Political Economy* 82 (2): pp. S11–S26.

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

Charles, Kerwin Kofi, and Ming Ching Luoh. 2010. “Male Incarceration, the Marriage Market, and Female Outcomes.” *The Review of Economics and Statistics* 92 (3): pp. 614–627.

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

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

Duffie, Darrell, and Kenneth J. Singleton. 1993. “Simulated Moments Estimation of Markov Models of Asset Prices.” *Econometrica* 61 (4): pp. 929–952.

Easterlin, Richard A. 1987. *Birth and Fortune: The Impact of Numbers on Personal Welfare.* Chicago: Chicago University Press.

Ellwood, David T., and Jonathan Crane. 1990. “Family Change Among Black Americans: What Do We Know?” *The Journal of Economic Perspectives* 4 (4): pp. 65–84.

Gale, D., and L. S. Shapley. 1962. "College Admissions and the Stability of Marriage." *The American Mathematical Monthly* 69 (1): pp. 9–15.

Greenwood, Jeremy, and Nezih Guner. 2009. "Marriage and Divorce since World War II: Analyzing the Role of Technological Progress on the Formation of Households." *NBER Macroeconomics Annual* 23 (231-276): pp.

Hitsch, Gunter J. 2010. "Matching and Sorting in Online Dating." *American Economic Review* 100 (1): 130–63 (March).

Iyigun, Murat, and Randall P. Walsh. 2007. "Building the Family Nest: Premarital Investments, Marriage Markets, and Spousal Allocations." *The Review of Economic Studies* 74 (2): pp. 507–535.

Lee, Bong-Soo, and Beth Fisher Ingram. 1991. "Simulation Estimation of Time-series mMdels." *Journal of Econometrics* 47 (2-3): pp. 197–205.

Lerman, Robert I. 1989. "Employment Opportunities of Young Men and Family Formation." *The American Economic Review, Papers and Proceedings* 79 (2): pp. 62–66.

McFadden, Daniel. 1989. "A Method of Simulated Moments for Estimation of Discrete Response Models Without Numerical Integration." *Econometrica* 57 (5): pp. 995–1026.

Mortensen, Dale T. 1988. "Matching: Finding a Partner for Life or Otherwise." *American Journal of Sociology* 94:pp. S215–S240.

Pakes, Ariel, and David Pollard. 1989. "Simulation and the Asymptotics of Optimization Estimators." *Econometrica* 57 (5): pp. 1027–1057.

Pampel, Fred C., and H. Elizabeth Peters. 1995. "The Easterlin Effect." *Annual Review of Sociology* 21:pp. 163–194.

Peters, Michael, and Aloysius Siow. 2002. "Competing Premarital Investments." *Journal of Political Economy* 110 (3): pp. 592–608.

Plotnick, Robert D. 1990. "Welfare and Out-of-Wedlock Childbearing: Evidence from the 1980s." *Journal of Marriage and Family* 52 (3): pp. 735–746.

Wilson, William J. 1987. *The truly disadvantaged: the inner city, the underclass, and public policy*. Chicago: Chicago University Press.

A Proofs

A.1 Proof of Proposition 2

Consider a positive change to cohort size. According to equation (8), in steady state the reservation value of a young man is the solution to the following equation:

$$\underline{\eta}_{0,ss} = 2\delta + \beta \frac{1-\beta^T}{1-\beta^{T+1}} \gamma \left\{ E\left[\eta \left| \eta \geq 2\delta \right.\right] - 2\delta \right\} \left(1 - F\left(2\delta\right)\right) \frac{1}{1 + \left(1 - \theta^m_{0,ss}\left(1 - F\left(\underline{\eta}_{0,ss}\right)\right)\right)}.$$

By substituting $\theta^m_{0,ss}$ with $\theta^m_{0,\tau}$ and by using the result that $\theta^m_{0,\tau} > \theta^m_{0,ss}$, we obtain the following inequality:

$$\underline{\eta}_{0,ss} < 2\delta + \beta \frac{1-\beta^T}{1-\beta^{T+1}} \gamma \left\{ E\left[\eta \left| \eta \geq 2\delta \right.\right] - 2\delta \right\} \left(1 - F\left(2\delta\right)\right) \frac{1}{1 + \left(1 - \theta^m_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,ss}\right)\right)\right)}.$$

Since the left hand side of the inequality is increasing in $\underline{\eta}_0$ and the right hand side is decreasing in $\underline{\eta}_0$, equation (7) implies that $\underline{\eta}_{0,\tau} > \underline{\eta}_{0,ss}$.

A.2 Proof of Proposition 3

The total number of women that marry in a particular cohort is given by the total number of women in the cohort time the probability that a woman in that cohort marries. As a consequence, the fraction of women in a cohort that marries is simply the probability of marriage for those women. The probability that a woman marries can be written as the probability that she meets a younger man times the probability she marries him plus the probability she meets an older man times the probability she marries him, i.e.

$$P\left(\text{woman marries at } \tau\right) = \theta^w_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) + \left(1 - \theta^w_{0,\tau}\right)\left(1 - F\left(2\delta\right)\right)$$

Define $1+\lambda_\tau = \dfrac{F\left(\underline{\eta}_{0,\tau}\right)}{F\left(\underline{\eta}_{0,ss}\right)}$ and $1+\phi_\tau = \dfrac{\theta^w_{0,\tau}}{\theta^w_{0,ss}}$, where $\lambda_\tau > 0$ and $\phi_\tau > 0$ because $\dfrac{\partial \underline{\eta}_{0,\tau}}{\partial N_0} > 0$ and $\dfrac{\partial \theta^w_{0,\tau}}{\partial N_0} > 0$. We then have

$$
\begin{aligned}
& P\left(\text{woman marries at } \tau\right) = \\
= \;& \theta^w_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) + \left(1 - \theta^w_{0,\tau}\right)\left(1 - F\left(2\delta\right)\right) \\
= \;& \theta^w_{0,ss}\left(1+\phi_\tau\right)\left(1 - F\left(\underline{\eta}_{0,ss}\right)\left(1+\lambda_\tau\right)\right) + \left(1 - \theta^w_{0,ss}\left(1+\phi_\tau\right)\right)\left(1 - F\left(2\delta\right)\right) \\
= \;& \theta^w_{0,ss}\left(1 - F\left(\underline{\eta}_{0,ss}\right)\right) + \left(1 - \theta^w_{0,ss}\right)\left(1 - F\left(2\delta\right)\right) - \theta^w_{0,ss}\lambda_\tau F\left(\underline{\eta}_{0,ss}\right) + \theta^w_{0,ss}\phi_\tau\left(1 - F\left(\underline{\eta}_{0,ss}\right)\left(1+\lambda_\tau\right)\right) \\
& - \theta^w_{0,ss}\phi_\tau\left(1 - F\left(2\delta\right)\right) \\
= \;& P\left(\text{woman marries at } ss\right) - \theta^w_{0,ss}\lambda_\tau F\left(\underline{\eta}_{0,ss}\right) + \theta^w_{0,ss}\phi_\tau\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) - \theta^w_{0,ss}\phi_\tau\left(1 - F\left(2\delta\right)\right) \\
< \;& P\left(\text{woman marries at } ss\right) - \theta^w_{0,ss}\lambda_\tau F\left(\underline{\eta}_{0,ss}\right) \\
< \;& P\left(\text{woman marries at } ss\right).
\end{aligned}
$$

A.3 Proof of Proposition 4

The average difference in age between spouses at marriage can be computed as the difference conditional on the woman marrying a younger man times the corresponding probability plus the difference conditional on marry an older man times the corresponding probability, i.e.

$$E\left[\Delta age\right] = E\left[\Delta age | younger\ man\right] P\left(younger\ man\right) + E\left[\Delta age | older\ man\right] P\left(older\ man\right).$$

Without loss of generality suppose the difference in age if a woman marries a younger man is equal to x whereas the corresponding difference in a marriage with an old man is equal to y with $x < y$. $E\left[\Delta age\right]$ can therefore be written as follows:

$$E\left[\Delta age\right] = x\theta^w_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) + y\left(1 - \theta^w_{0,\tau}\right)\left(1 - F\left(2\delta\right)\right).$$

By taking the derivative of both sides with respect to cohort size, we obtain the following equation:

$$\frac{\partial E\left[\Delta age\right]}{\partial N_0} = \frac{\partial \theta^{w}_{0,\tau}}{\partial N_0}\left[x\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) - y\left(1 - F\left(2\delta\right)\right)\right] - \frac{\partial \theta^{m}_{0,\tau+1}}{\partial N_0} f\left(\underline{\eta}_{0,\tau}\right)\theta^{w}_{0,\tau} B$$

where $B = \beta\frac{1-\beta^{T}}{1-\beta^{T+1}}\gamma\left\{E\left[\eta\left|\eta \geq 2\delta\right.\right] - 2\delta\right\}\left(1 - F\left(2\delta\right)\right)$, the factor that multiplies the meeting probability in the reservation value equation, and $f\left(\eta\right)$ the probability density function of η.

Equation (7) and $\underline{\eta}_{0,\tau} > \underline{\eta}_{0,ss}$ imply that $\frac{\partial \theta^{w}_{0,\tau+1}}{\partial N_0} > 0$. We have also shown that $\frac{\partial \theta^{w}_{0,\tau}}{\partial N_0} > 0$. Finally, $x < y$, $\underline{\eta}_{0,\tau} > 2\delta$, and hence $1 - F\left(\underline{\eta}_{0,\tau}\right) < 1 - F\left(2\delta\right)$. The result therefore follows.

B Tables and Figures

Table 1: Time Series Regression of Log Share Ever Married

	Women, by 30 White	Women, by 30 Black	Men, by 30 White	Men, by 30 Black	Men, by 40 White	Men, by 40 Black
Log Cohort Size	−0.226*	−0.957*	−0.342*	−0.658*	−0.103*	−0.316*
	(0.021)	(0.048)	(0.030)	(0.040)	(0.009)	(0.024)
R-squared	0.63	0.86	0.67	0.80	0.69	0.76

* Significant at 1%. Standard errors in parentheses.

Table 2: Cross-Sectional Regression of Log Share Ever Married by 30

	Women, White	Women, Black	Men, White	Men, Black	Men, White 1950-1960	Men, Black 1950-1960
10-Yr. Difference in	-0.354*	-0.291**	-0.742***	-0.547***	-0.987***	-0.631***
Log Cohort Size	(0.189)	(0.144)	(0.204)	(0.084)	(0.377)	(0.088)
N	147	114	147	114	49	41
R-squared	0.41	0.80	0.57	0.75	0.19	0.47

* Significant at 10%. ** Significant at 5%. *** Significant at 1%. Standard errors in parentheses.

Table 3: Summary Statistics, 1940: Low, Medium, and High-Mobilization States (White Only)

	Low	Medium	High
Percent Men Inducted into Army	0.277	0.305	0.323**
	(0.00717)	(0.00558)	(0.0117)
Share Never Married at 30	0.240	0.222	0.211
	(0.063)	(0.047)	(0.046)
Share Farmers	0.286	0.251	0.230
	(0.132)	(0.111)	(0.180)
Age	34.74	34.41	34.15
	(0.842)	(1.152)	(0.891)
Men's Employment	0.849	0.840	0.834
	(0.023)	(0.031)	(0.030)
Women's Employment	0.265	0.271	0.277
	(0.072)	(0.054)	(0.071)
Log Income	6.545	6.606	6.577
	(0.241)	(0.174)	(0.304)
Years of Education	9.838	9.671	9.156*
	(0.649)	(0.661)	(0.518)
Number of Children by Age 35	2.455	2.631	2.588
	(0.478)	(0.409)	(0.501)

* Difference in means between high- and low-mobilization groups significant at 5%. ** Significant at 1%.
Notes: Standard errors in parentheses. Averages are for white individuals only. Total number of observations equals 49. Low-mobilization states (bottom third of states) are those that report the share of men inducted as less than or equal to 0.288. High-mobilization states (top third) are those where the share of men inducted into the army is exceeds 0.312. Analysis includes 48 states (all except Hawaii and Alaska), as well as Washington, D.C. Percent inducted is the cumulative number of white men ages 18-44 inducted, divided by the number of white men ages 18-44 registered, as of Sept.1, 1945. Numbers taken from the Selective Service System's Special Monograph no. 12, Appendix F, Table 164. All other averages constructed using IPUMS Census, 1940. Share never married at 30 is calculated for white men ages 30 and 31. Observations used to construct the average number of children by age 35 are all white women ages 35 and 36. Averages for these two variables exclude Nevada due to small sample size in the state under the age restrictions. None of the results are substantively affected by including or excluding Nevada. Share farmers is the share of white men ages 18 to 50 whose residence has farm status. Age and log income are reported for white men ages 18 to 50. Employment and years of education are calculated for white men and women ages 18 to 50.

Table 4: First-Stage Regression: Log Births and Mobilization

	(1)	(2)
Mobilization Rate * 1940	0.769	0.688
	(0.478)	(0.518)
Mobilization Rate * 1941	0.558	0.482
	(0.479)	(0.518)
Mobilization Rate * 1942	1.069**	1.403***
	(0.480)	(0.518)
Mobilization Rate * 1943	0.281	1.081**
	(0.482)	(0.518)
Mobilization Rate * 1944	-0.801*	0.359
	(0.483)	(0.519)
Mobilization Rate * 1945	-1.081**	-0.205
	(0.482)	(0.518)
Mobilization Rate * 1946	-0.391	0.355
	(0.480)	(0.518)
Mobilization Rate * 1947	-0.214	0.426
	(0.479)	(0.518)
Mobilization Rate * 1948	-0.748	-0.230
	(0.478)	(0.518)

* Significant at 10%. ** Significant at 5%. *** Significant at 1%. Standard errors in parentheses.
Notes: The variable "Mobilization Rate * Year " is the interaction between the share of white men ages 18 to 44 inducted into the army (see note in Table 2) and a year dummy. Column 1 allows for year-fixed effects, while column 2 allows for time-region fixed effects. Regressions additionally include state fixed effects, as well as controls for 1940 baseline educational and income levels, interacted with a fourth-order polynomial in time.

Table 5: Second Stage Regression: Log Marriage Rates and Log Cohort Size

	(1) Men	(1) Women	(2) Men	(2) Women
Log Cohort Size	-0.100*	-0.055*	-0.061*	-0.041*
	(0.005)	(0.004)	(0.015)	(0.013)

* Significant at 1%. Standard errors in parenthesis. Column 1 allows for year-fixed effects, while column 2 allows for time-region fixed effects. Regressions additionally include state fixed effects, as well as controls for 1940 baseline educational and income levels, interacted with a fourth-order polynomial in time.

Table 6: Regressions: Log Age Difference and Log Cohort Size

	Time Series, 1930-1975	Cross-State, 1940-1950	Cross-State, 1950-1960	Cross-State, 1960-1970
Log Cohort Size	-0.592***			
	(0.043)			
10-Yr. Difference in		0.110	-0.172**	-0.130*
Log Cohort Size		(0.235)	(0.077)	(0.078)
N	46	49	49	49
R-squared	0.81	0.00	0.05	0.09

* Significant at 10%. ** Significant at 5%. *** Significant at 1%. Standard errors in parentheses.

Table 7: Estimated Parameters

Parameters	Estimates	Standard Errors
First Shape Parameter	0.020	[0.010]
Second Shape Parameter	0.071	[0.039]
Value of Being Single	0.107	[0.037]
Fraction of Men Unwilling to Marry	13.3	[0.143]

Figure 1: Number of Marriages

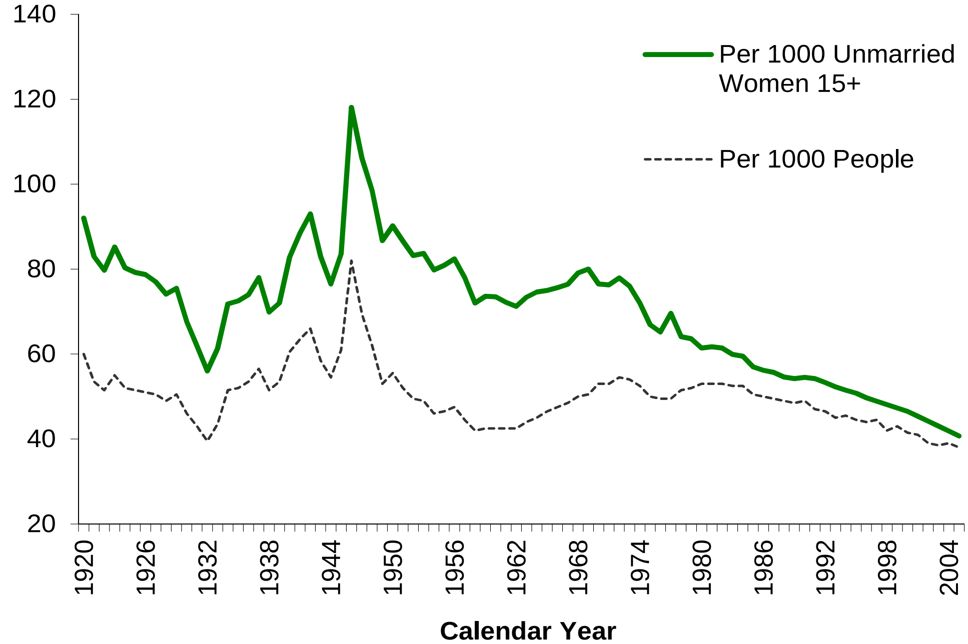


Figure 2: Share of Women Ever Married, Ages 18 to 30

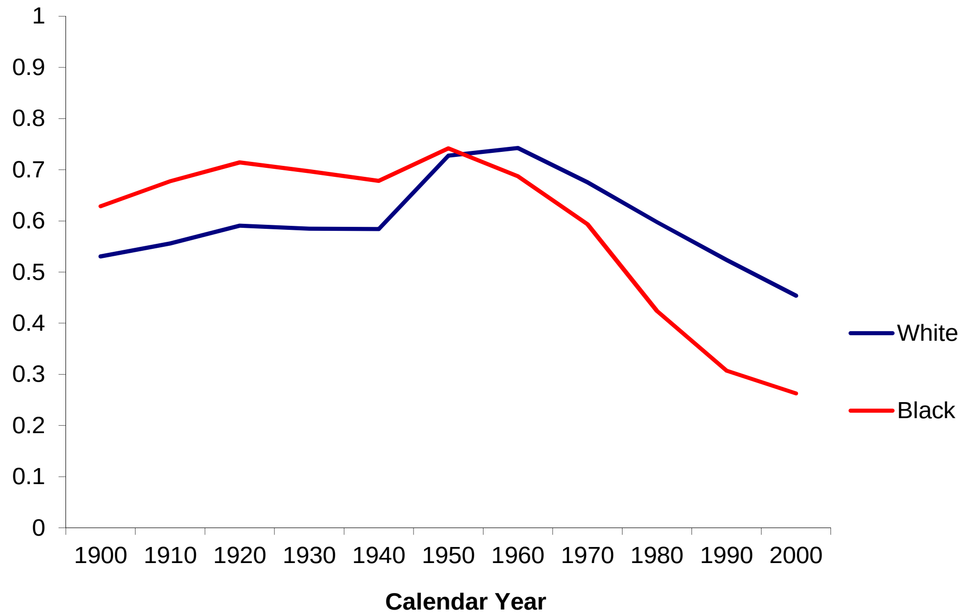


Figure 3: Different Measures of Marriage Rates

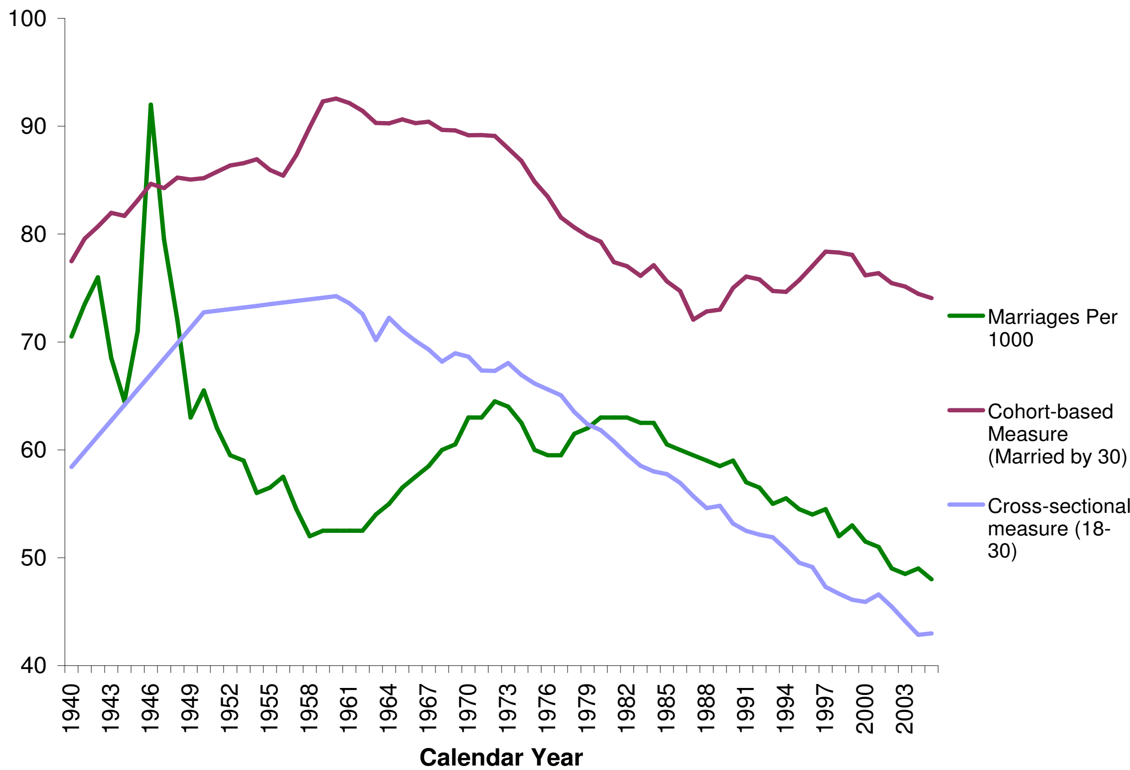
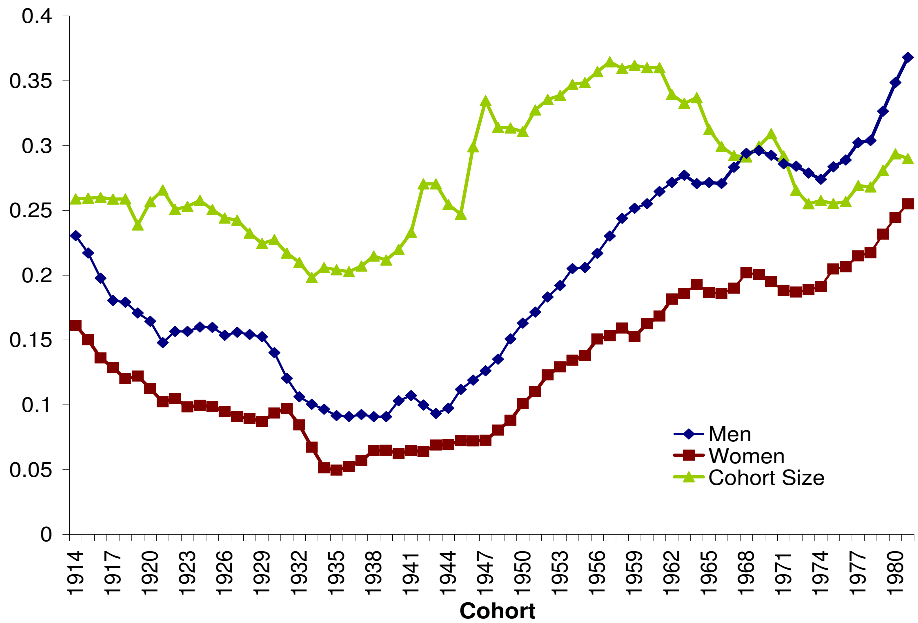
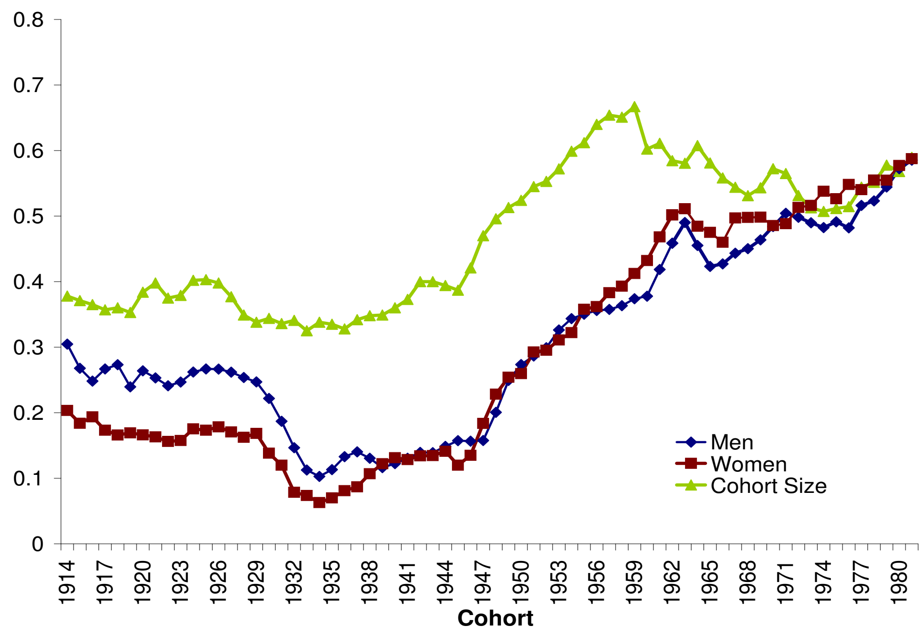


Figure 4: Share Never Married By 30

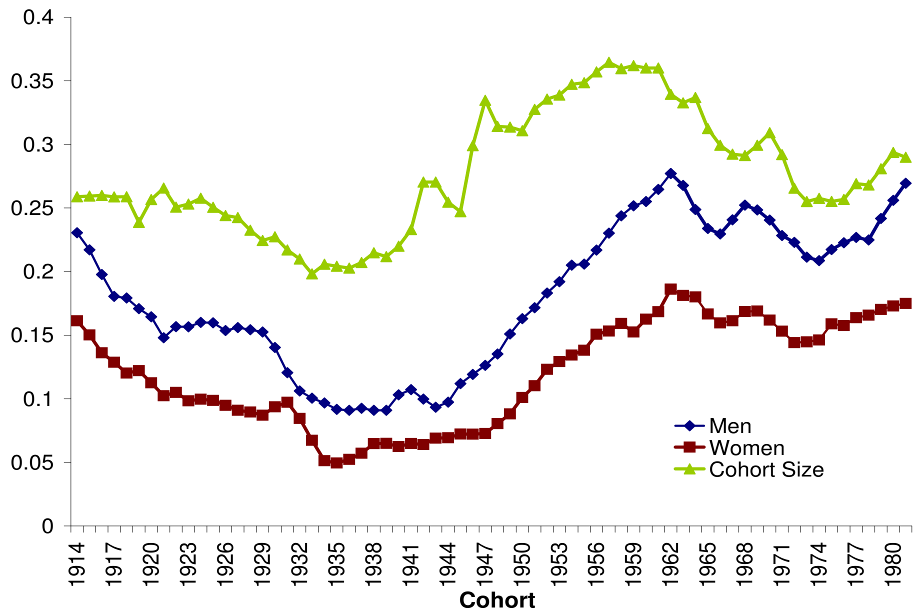


(a) White

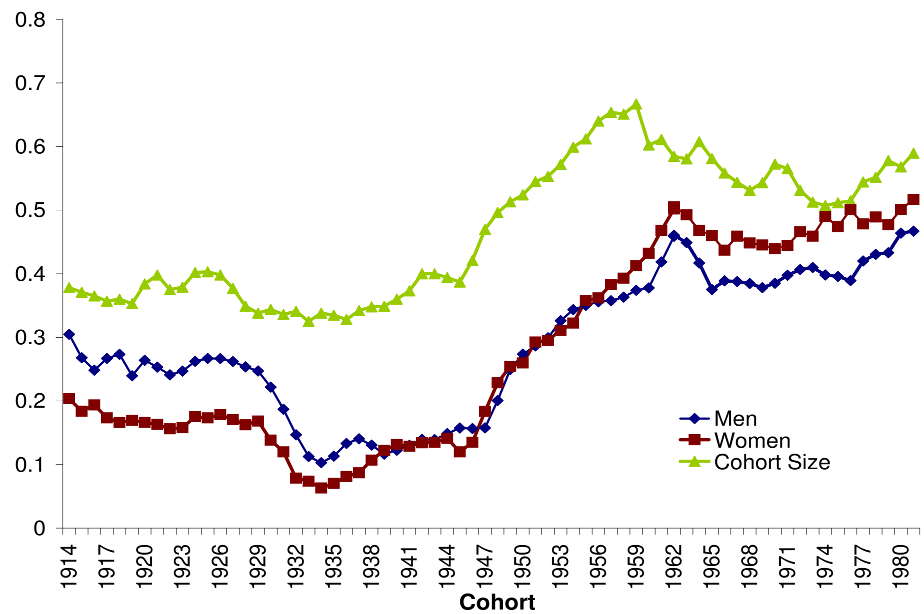


(b) Black

Figure 5: Share Never Married and Not Cohabiting By 30

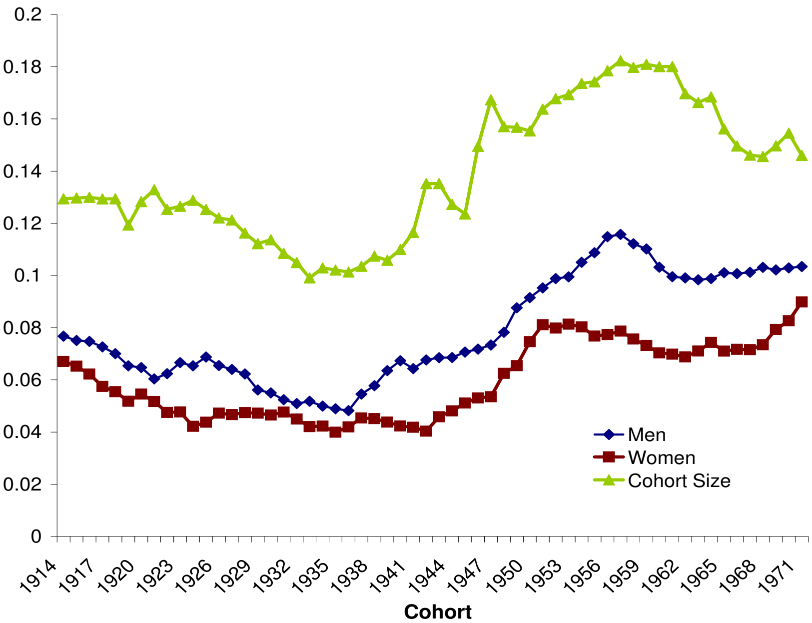


(a) White

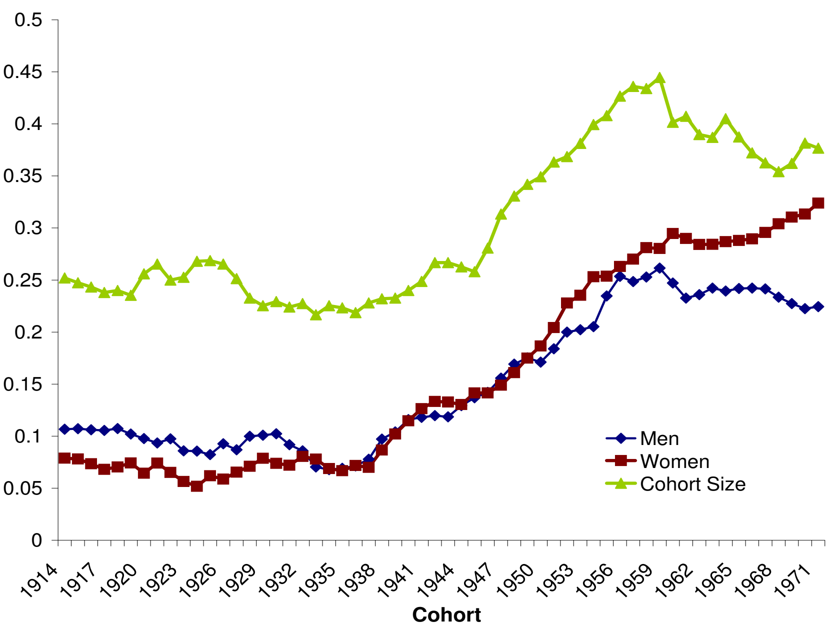


(b) Black

Figure 6: Share Never Married and Not Cohabiting By 40



(a) White



(b) Black

Figure 7: Mobilization Rates and Changes in Total Births, 1935-1940

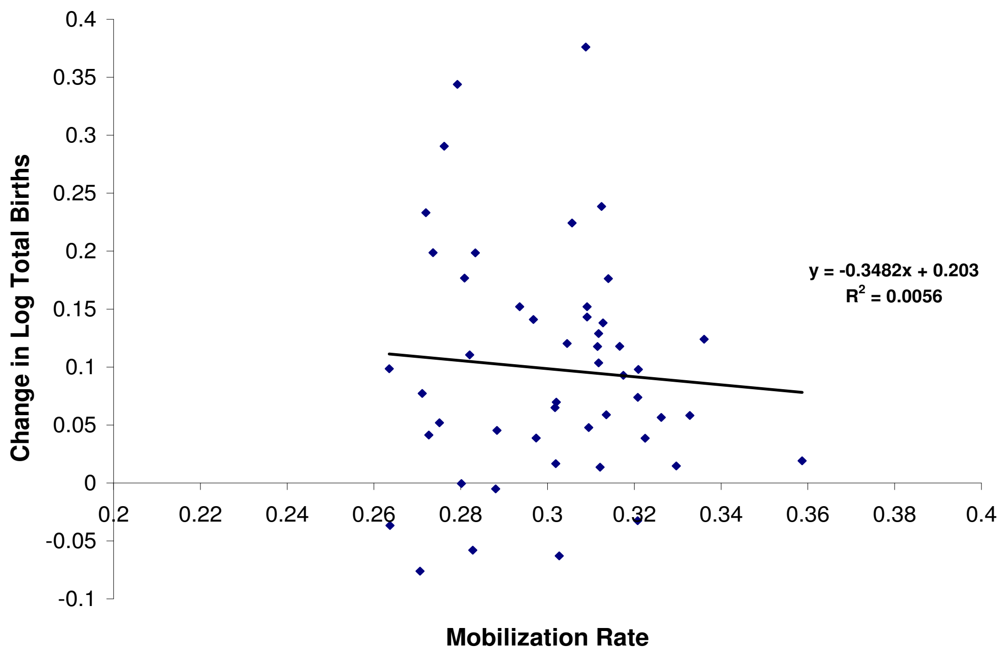


Figure 8: Mobilization Rates and Changes in Total Births, 1940-1942

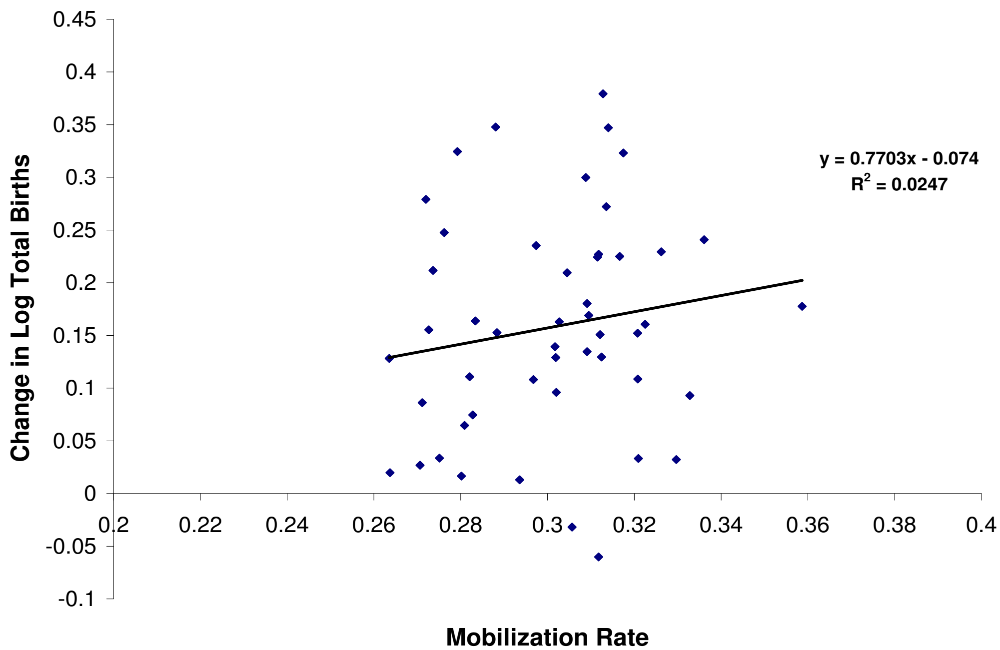


Figure 9: Mobilization Rates and Changes in Total Births, 1943-1945

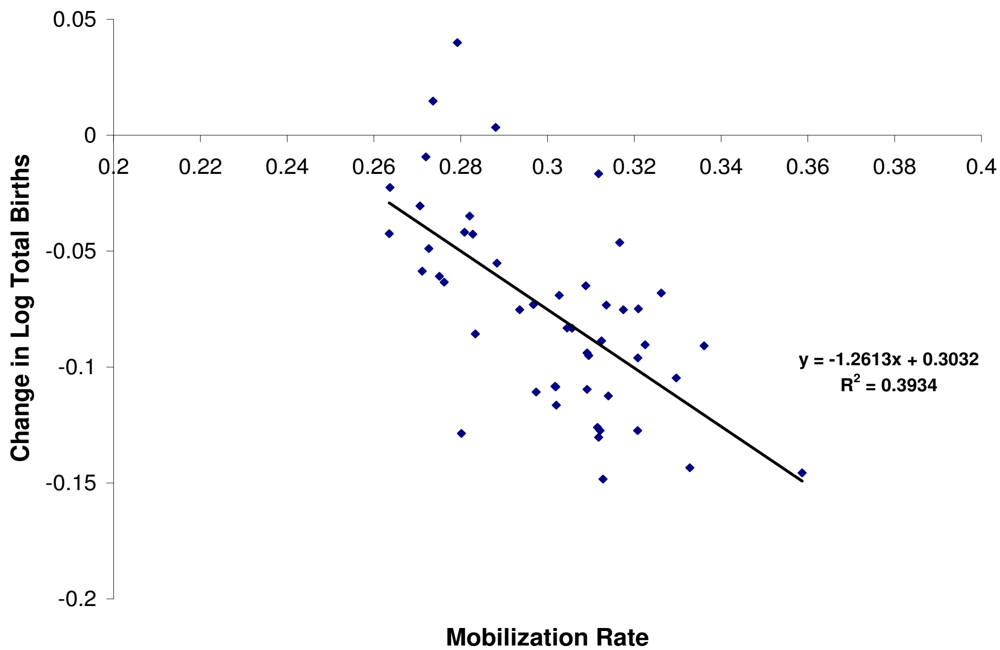


Figure 10: Mobilization Rates and Changes in Total Births, 1945-1946

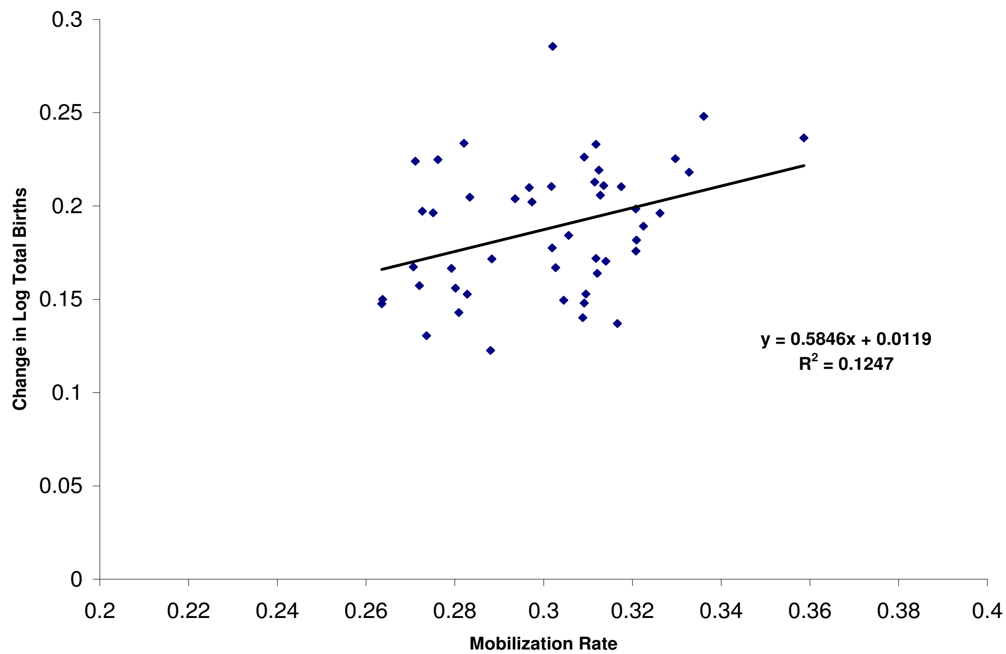


Figure 11: Age Difference Between Spouses by Cohort, Whites

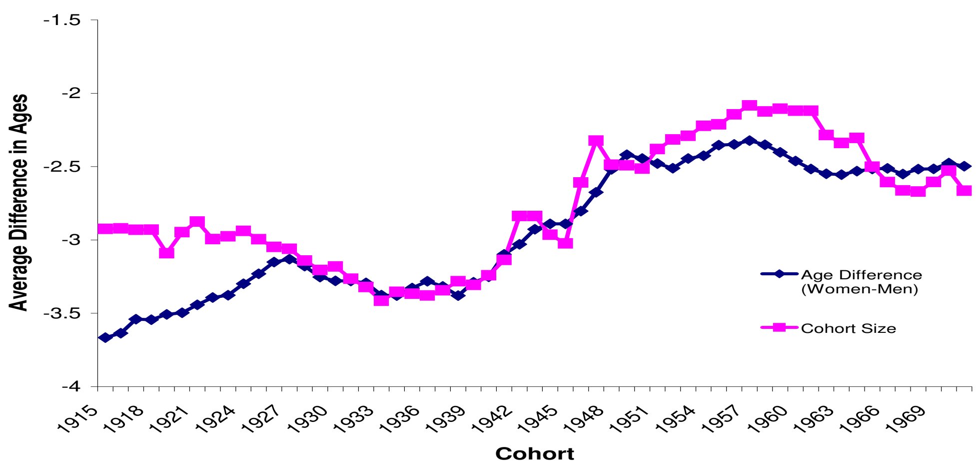


Figure 12: Observed and Simulated Never Married Rates