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Essays in Labor Economics on Marriage, Education, and Labor Supply

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in Economics

by

Mary Ann Bronson

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ABSTRACT OF THE DISSERTATION

Essays in Labor Economics on Marriage, Education, and Labor Supply

by Mary Ann Bronson

Doctor of Philosophy in Economics

University of California, Los Angeles, 2014

Professor Maurizio Mazzocco, Chair

This dissertation seeks to understand two main issues. The first issue concerns changes in the gender gaps in college attendance and choice of majors between 1960 and 2010. The second main issue concerns changes in the marriage rate in the US since the early twentieth century.

The main objective of Chapters 1 and 2 is to answer the following two questions about educational choices: Why do women today invest in a college education at much higher rates than men, whereas fifty years ago men graduated more frequently? And given their high college attendance rates today, why do women continue to select disproportionately into lower-paying majors, with almost no gender convergence along this margin since the mid-1980s? In Chapter 1, I document first that changes in returns to skill over time and gender differences in wage premiums across majors cannot explain the observed gender gaps in educational choices. I then provide reduced-form evidence that two factors help explain the observed gender gaps: first, college degrees provide insurance against very low income for women, especially in case of divorce; second, majors differ

substantially in the degree of "work-family flexibility" they offer, such as the size of wage penalties for temporary reductions in labor supply.

Based on this reduced-form evidence, in Chapter 2 I construct and estimate a dynamic structural model of marriage, educational choices, and lifetime labor supply. I use the model to analyze the contribution of changes in wages and changes in the marriage market to the observed educational investment patterns over time. I estimate that the insurance value of the college degree for women in case of divorce is equivalent to about 31% of the college wage premium. I also estimate that the share of women choosing high-return science and business majors would increase from 34% to 45% if wage penalties for labor supply reductions were equalized across occupations. Finally, I test the effects of two sets of policies on individuals' choice of major: a differential tuition policy that charges less for science and technical majors, as has been proposed in some states; and interventions intended to improve work-family flexibility. My results show that some family-friendly policies increase the share of women in science and business majors substantially, while others further widen both college gender gaps.

Chapter 3, joint work with Maurizio Mazzocco, analyzes changes in U.S. marriage rates over nearly a century. We propose an explanation for these changes in three stages. First, we show that changes in cohort size alone can account for around 50 to 70% of the variation in marriage rates since the 1930s for both black and white populations. Specifically, increases in cohort size reduce marriage rates, whereas declines in cohort size have the opposite effect. Using plausibly exogenous variation in access to oral contraceptives, and consequently the number of births, across states we provide evidence that the relationship between changes in cohort size and changes in marriage rates is causal. Next, we develop a dynamic search model of the marriage market that qualitatively generates this observed relationship, and derive a testable implication about cohort size's effect on spouses' age differences. Finally, we estimate the model and investigate its consistency with the data. We fail to reject it using the derived implication, and find that it can quantitatively explain much of the observed variation in marriage rates.

The dissertation of Mary Ann Bronson is approved.

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Sarah Reber

Maurizio Mazzocco, Committee Chair

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Chapter 3 is joint work with Maurizio Mazzocco.

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Chapter 1

Marriage, Education, and the Lifecycle Labor Decisions: Reduced-Form Evidence

1.1 Introduction

Why do women today invest in a college education at much higher rates than men, whereas fifty years ago men graduated more frequently? And given their high college attendance rates today, why do women continue to select disproportionately into lower-paying majors? The main objective of the first two chapters of this dissertation is to answer these two questions.

Historically, men made up the majority of college students, and earned more than 90% of all high-paying degrees in science and business. In the 1970s and early 1980s, men and women converged substantially both in college graduation rates as well as in their choices of college major, with more women choosing science and business degrees. It has been well-documented that women reversed the "gender gap" in graduation rates by the mid-1980s and now constitute the majority of college students, although the reason why this occurred is still an open question.[1] What has been less well-documented is that convergence between men and women in choice of major mostly ceased after the mid-1980s. In 1985, nearly 80% of education degrees and about 85% of

[1] See, for example, Goldin, Katz, and Kuziemko (2006), Chiappori, Iyigun, and Weiss (2009), Charles and Luoh (2003), Vincent-Lancrin (2006). Today women make up 58% of U.S. college students.

degrees in health support fields, but less than 30% of hard science and engineering degrees were awarded to women. The same is true today.[2]

The question of why women graduate at much higher rates than men, but with very different majors, has implications for a range of individual outcomes, as well as for macroeconomic outcomes like supply of skill to the labor market. As women outpace men in college attendance, their low participation in majors like science and engineering contributes to potentially low supply of science-related skills in the U.S. More generally, the observed patterns, like women's higher graduation rates despite lower lifetime labor supply, run counter to the predictions of a standard human capital investment model (Becker (1962), Ben-Porath (1967), Mincer and Polachek (1974)), and raise questions about the determinants of returns to different educational choices for men and women.

In Chapters 1 and 2 of this dissertation, I explain the dynamics of men's and women's educational investments from 1960 to 2010. The chapters make three main contributions. The first contribution is to document reduced-form evidence about the factors that potentially explain the above-mentioned gender gaps over time. The reduced-form evidence is the focus of the present chapter. I provide this evidence in three steps. In the first step, I show that changes in the wage premium over time and differences in major-specific premiums across men and women cannot readily account for the observed gender differences in college attendance or decisions about majors.

In the second step, I provide evidence that changes in the marriage market starting in the 1970s changed the relative returns to a college education for men and women. I use quasi-experimental variation in the timing of unilateral, no-fault divorce law reforms across states to document that the reforms increased women's college graduation rates relative to men, and made them more likely to select high-paying majors in business and science-related fields. There is a simple intuition for this finding. Women with high school education or less draw from a substantially lower wage distribution than men, and are also more likely to have custody of and financial responsibility for children. A college degree allows women access to higher paid jobs, providing insurance against

[2] National Center for Education Statistics, Digest of Education Statistics (2012), Tables 343-365.

very low income realizations for women outside a two-earner household.

In the final step, I present evidence that majors are characterized not only by different wage premiums, but also by different levels of "work-family flexibility." By flexibility I mean that some majors are associated with occupations that provide easier access to part-time or part-year work and have lower wage penalties in case of a temporary absence from the workforce or a reduction of weekly hours worked. I show that college women reduce their labor supply substantially during their childbearing years, and that these patterns differ across majors and occupations. The data patterns I document indicate that women are more likely than men both to take advantage of flexibility associated with some majors, as well as to choose more flexible majors.

The empirical patterns documented in the present chapter indicate that insurance and flexibility are important drivers of the gender gaps. However, it is difficult to quantify the impact of these factors on the gender gaps using the reduced-form analysis alone. The main reason for this is that other variables, like returns to skill, also changed over this time period, and will also affect decisions about education, labor supply, marriage, and divorce.

To address this, as a second main contribution, I develop and estimate a dynamic structural lifetime model of individual decisions about education, marriage, and labor supply, which will be the focus of in Chapter 2. The model is constructed based on the documented data patterns in the present chapter. The final contribution is to analyze the effects of different policies on educational choices of men and women. The estimated model in Chapter 2 is well-suited for this purpose because it can analyze policies' effects on decisions about labor supply, occupational choice, and household formation and dissolution. I study two sets of policies: a differential tuition policy that lowers the cost of technical majors (Alvarez (2012)), and a set of "family-friendly" policies, like paid maternity leave or part-time work entitlements, that have been proposed or enacted in various countries to improve work-family flexibility and encourage gender equality in the labor force.

The rest of the chapter is organized as follows. Section 1.2 summarizes the related literature. Sections 1.3 to 1.5 present evidence on drivers of changes in the gender gap in graduation rates and choice of majors over time. Sections 1.6 and 1.7 present evidence on the persistence in the gender gap in choice of majors over since the mid-1980s. Section 1.8 concludes.

1.2 Related Literature

Chapters 1 and 2 of this dissertation contribute to several bodies of literature. The first is the literature on gender differences in educational choices. Most studies in this literature focus on the gender gap in college graduation. A variety of explanations have been proposed. Goldin, Katz, and Kuziemko (2006) document that women historically perform better in high school than men. If this reduces the cost of going to college for women, then this can help explain the current gender gap in college attendance. However, additional factors are necessary to explain the dynamics in the gap over time. Chiappori, Iyigun, and Weiss (2009) consider the effect of schooling on the marital surplus share individuals can extract at the time of marriage when there are different shares of educated men and women in the marriage market. They show that under some conditions, women may invest more in education than men. Charles and Luoh (2003) show that the variability of earnings increases with education, but increases less for women than men. They argue that college is a less risky and therefore better investment for women. In the present paper, I find that women select disproportionately into majors with flatter earning profiles and lower wage penalties, in line with the idea that the observed variance of college earnings will be lower for women than for men. The interpretation for this pattern, however, is different, since in this paper it is an endogenous outcome about choices of major rather than an ex ante driver of overall decisions about college.

In a study that analyzes gender differences in choice of major, Wiswall and Zafar (2013) experimentally generate variation in undergraduates' subjective beliefs about future earnings in different majors to estimate a model of choice of college major. They find that earnings are a significant determinant of major choice, but residual factors, interpreted as tastes, are dominant, with women more likely to have a taste for the arts and humanities. This finding is similar to that of Beffy et al. (2011) and Gemici and Wiswall (2011). However, these studies do not account for non-financial characteristics of majors, like flexibility, which differentially affect lifetime expected returns for women and men. As a result, the differences in choices about majors that may be explained by such characteristics are instead attributed to tastes.

This dissertation also contributes to the small, but growing literature on the flexibility of professions and the relationship with human capital investments. Recent work includes Goldin and Katz (2011), Goldin and Katz (2012), and Bertrand, Goldin, and Katz (2010).[3] The paper also builds on the literature that attempts to model and estimate the dynamic, intertemporal aspects of household decisions using a collective household model (Mazzocco, Ruiz, and Yamaguchi (2009), Lundberg et al. (2003), Van der Klaaw and Wolpin (2004), Voena (2012) Gemici and Laufer (2011)).

This is the first analysis, to the best of my knowledge, to model explicitly the the effects of marriage, divorce and household labor supply on the lifetime returns to college by type of major for men and women. Understanding these broader sets of returns is important, especially for research that makes inference about potential unobserved determinants of men's and women's educational choices, like ability, effort costs, and tastes.

1.3 Trends over Time

Figures 1.1 and 1.2 provide the motivation for the main questions of the paper. Figure 1 documents the shares of men and women graduating between 1960 and 2010. Graduation rates increased substantially over this time period in line with trends in rising skill premiums, from 17% to nearly 30% for men, and from 10% to more than 35% for women. A considerable difference between men and women is that in the early 1970s, men's college attainment stalled and even fell, while women continued to increase their four-year college graduation rates over almost the entire time period, reversing the "college gender gap" that historically favored men. Today, about 58% of college students are women.

By contrast, men and women converged only partly in their choice of college major. Figure 1.2 graphs the share of undergraduate degrees in each major awarded to women starting in 1970,

[3] For early work on majors and flexibility, see Polachek (1978). See also Polachek (1981).

the earliest year that NCES data are available for most majors. As Figure 1.2 documents, a partial but significant convergence in choices of major between men and women occurred over the 1970s. Most dramatically, the share of business degrees earned by women increased from less than 10% in 1970 to more than 40% by 1985. After the early 1980s, Figure 1.2 documents that gender convergence in choice of majors almost completely ceased. Both in 1985 and 2010, nearly 80% of education degrees and about 85% of degrees in health support fields were awarded to women. By contrast, both in 1985 and 2010, fewer than 30% of hard science and engineering degrees went to women.

In the remainder of the chapter, I provide evidence on the potential drivers of these patterns. I focus first on the change in men's and women's graduation rates over time. I then examine changes in choice of college majors over the 1970s. Finally, I consider the persistence in the gender gap in majors since the mid-1980s.

1.4 Changes in Graduation Rates

In Figure 1.3, Panel A replicates graduation rates by gender, while Panel B graphs the college wage premium for men and women from 1960 to 2010. In a standard human capital investment model (Ben-Porath (1967), Becker (1962)), the wage premium is an important driver of educational choices. Differences in the wage premium for men and women over time may therefore help explain changes in their educational investments. The measure of the premium graphed in Figure 1.3 is the difference in log income for college and high school graduates between the ages of 22 and 50 working full-time, with flexible controls for age and race. Constructing the wage premium strictly for younger workers, e.g. ages 22 to 30, yields the same time series pattern.

Figure 1.3 shows that wage premiums evolved similarly for men and women between 1960 and 2010, and are unlikely to explain the gender differences in graduation rates over time.[4] In fact,

[4]Note that changes in choice of major over time affect the evolution of the wage premium in Figure 1.3. As Figure 1.2 documents, women converged partly with men in choice of majors, and selected increasingly into higher-paying majors. This drives up women's observed wage premium over the second half of the time period in Figure 1.3. Note that if one were to hold constant the share of women who choose each type of major at the 1970 level, the wage

women's premiums grew more slowly than men's premiums between the mid-1970s and 2010, while their graduation rates increased more rapidly. Interestingly, Figure 1.3 shows that for men educational choices are in line with the predictions of a standard human capital investment model. The wage premium doubled from around 30 log points in 1960 to more than 60 log points in 2012, in keeping with the large increase in college attendance rates. As the premium declined in the 1970s, fewer men invested in a college education. However, for women one does not observe such a relationship between the wage premium and college graduation. Despite falling wage premiums, women continued to increase their college enrollment substantially in the 1970s, and thus rapidly converged with men.

Why did women continue to increase their graduation rates despite falling wage premiums? The timing in the gender convergence starting in the 1970s suggests that changes in the marriage market may provide one possible explanation. In particular, 1970 marked the beginning of so-called no-fault, unilateral divorce law reforms across the U.S., which made divorce significantly easier in most states. The reforms eliminated the need to demonstrate "fault," such as abuse, adultery, or negligence in court. As has been widely documented, the reforms were followed by a rapid, immediate increase in the number of divorces (Friedberg (1998), Wolfers (2003)).

Figure 1.4 maps the share of individuals divorced since 1960, as well as the ratio of women to men enrolled in four-year-universities, with the dotted line marking the start of divorce law reforms. The similar evolution in divorce rates and the gender gap, especially the rapid increase in both series in the 1970s, suggest a possible association between these factors.[5] The intuition behind why women's return to college may increase when divorce rates rise is that a college degree can provide an important form of "insurance" against low household income for women in case of divorce. While the focus of the discussion in this section is on divorce, the same economic intuition can also apply to unmarried women. Low-skill wages for women are substantially lower than those for men. In 2000, women ages 18 to 50 with less than a college degree employed full-

premium for women would be lower than the one recorded in Figure 1.3. It would therefore make it even more difficult to explain why women increased their college attendance rates relative to men after the 1970s.

[5]Note that women's rising wage premiums relative to men in the 1960s can help account for the limited gender convergence in the female-male enrollment ratio prior to 1970, but not for its rapid increase afterwards.

time earned $27,156, compared with $36,751 for men (IPUMS USA, 2000). Moreover, women on average bear a majority of the child care costs following separation or divorce (Grail (2002)). As a result, securing the college wage premium may become more valuable to women as they anticipate spending more of their lifetime in a single-earner household, something not captured by trends in the wage premium.

Friedberg (1998) documents that divorce law reforms were introduced in different years across states. The quasi-experimental variation in timing provides a test for the explanation that expected returns to a college education increase for young women when they anticipate a higher probability of divorce. The premise of the test is that if this is true, then one should observe women increasing their relative graduation rates in those states that already passed legislation that increases the probability of divorce.

To conduct such a test, I use data from Friedberg (1998) on the timing of unilateral, no-fault divorce law reforms, and Census data on educational attainment to construct four-year college graduation rates by birth cohort in each state. I then estimate the following equation,

$$Gap_{s,c} = \alpha + \sum_{a,s,c} \beta^{a}_{s,c} Ageatlaw^{a}_{s,c} + \sum_{s} \gamma_s + \sum_{c} \lambda_c + \varepsilon_{s,c} \tag{1.1}$$

where the outcome variable "gap" is the share of women graduating from college minus the share of men graduating from college in state s and birth cohort c. The dependent variables include state and year-of-birth (cohort) fixed effects, and a set of "age-at-law" indicators that are set equal to one if cohort c in state s was of a particular age a at the time of unilateral divorce law adoption. I construct the four-year graduation rates based on individuals ages 26 to 35 in the 1960 to 2000 Censuses, starting with the 1930 birth cohorts, up to the 1974 cohort, the youngest available for the analysis in the 2000 Census. This specification is very similar to the one used in the quasi-experiment in Stevenson and Wolfers (2006), except that for the main dependent variables I use indicators for age at the time of reform, rather than indicators for the number of years since reforms occurred.[6] To

[6]Census data is only available in decennial years. As a result, I cannot observe men and women in each calendar year at the typical college graduation age, 21 or 22, or how the shares graduating change with the number of years since

reduce the number of coefficients, I assign a single indicator variable to all individuals who were not yet born when a divorce law reform occurred.

If higher anticipated divorce rates increase the share of women graduating from college relative to men, then the coefficients for the "age-at-law" indicator will be positive for cohorts who were young enough to still make a decision about their educational investment at the time of the passage of the divorce law legislation in their state. This includes those 18 or younger at the time of the reform, and potentially those up to three or four years older, who can still decide whether or not to complete their degree. The gender gap for cohorts who were old enough to have already completed their undergraduate education by the time reforms occurred should not be affected.

Figure 1.5 graphs the coefficients from this regression, with age at law on the x-axis ordered from old to young. The coefficients $\beta^{a}_{s,c}$ on the age-at-law indicators are small and insignificant for ages above 21. They start to become significant at the 10 percent level at age 20, which suggests that the response to reforms was immediate. The coefficients remain positive and significant or marginally signficant at younger ages, i.e. for those who made a decision about whether or not to go to college after divorce law reforms occurred. The test therefore does not reject the hypothesis that women's educational returns increased relative to men's following divorce law reforms.

1.5 Changes in Choice of Major in the 1970s

Another pattern documented in the paper is that the gender gap in choice of college major narrowed in the 1970s. Switching to a higher-paying major like business or sciences constitutes an additional potential source of insurance for women in case of divorce. To test whether changes in divorce laws also contributed to the increase in the relative number of women in these fields, I conduct a test similar to the previous one. For this purpose, I first classify majors into two groups, "sciences/business" and "humanities/all others," and record the share of men and the share

reforms were passed. However, I can use a comparable cohort-based measure that focuses on the age that individuals were at the time of the reform. The measure similarly captures the number of years that have passed since the reform at the time those individuals made their educational decision.

of women in each cohort who choose science/business majors. Grouping major choices into these two broadly defined categories allows me to construct a single variable, the share of individuals choosing a science and business major, which simplifies the reduced form analysis.

Because the Census does not have information on majors, for this particular test I use state-level data from the Higher Education General Information Survey (HEGIS), which provides data on undergraduate degrees earned yearly by gender and field between 1965 and 1985. I conduct a similar test as before, using the following specification:

$$Gap_{s,y} = \alpha + \sum_{n,s,y} \beta^{n}_{s,y} Yearssincelaw^{n}_{s,y} + \sum_{s} \gamma_s + \sum_{y} \lambda_y + \varepsilon_{s,y} \tag{1.2}$$

where the outcome variable is the share of graduating college women who choose a science or business major in a given year y and state s minus the share of college men who graduate with such majors. “Years-since-law” is a set of indicators corresponding to the number of years since the divorce law reform was passed in state s. Because of the small number of observations, I group the years n for the years-since-law indicator in the following way: -2 to 0, 1 to 3, 4 to 6, 7 to 9, and 10 or more. The first of these indicators allows me to test for a pre-trend, similarly as in the previous analysis using age at time of divorce law. The omitted category includes all states that are three or more years away from passing a divorce law reform.

Table 1.1 reports the coefficients on the years-since-law indicators. As expected, the coefficient corresponding to zero to two years prior to the divorce law reforms is statistically zero. The coefficient on the indicator variable corresponding to one to three years is also insignificant. This contrasts with the results obtained for the gender gap in college attendance, in which the effect of divorce law reforms was immediate. Only after four years, the effect on the gender gap in majors becomes statistically significant. The difference between the two results is likely explained by the fact that it is difficult to switch majors after already completing one or more years of study. As a result, one would not observe a response in choice of major until at least four years after the divorce law reform. The coefficient on the indicator corresponding to 7 to 9 years continues to be significant, but at 10 years, the coefficient loses significance.

The results obtained using cross-state variation in the timing of divorce law reforms provide evidence that both "gender gaps" narrowed following divorce law reforms, suggesting that both the returns to getting a college education as well as the returns to getting a higher-paying major increased for women relative to men after the reforms. How large are these effects? The size of the regression coefficients from the reduced-from analysis suggest that divorce law reforms explain around 13% of the convergence between men and women in graduation rates observed in the 1970s. The effects on the gender gap in majors are somewhat smaller. However, it is possible that the size of the coefficients understates the real effect. Firstly, individuals' mobility across states after graduation introduces substantial measurement error in the analysis. Secondly, contamination effects may play a role. For example, as more states implement reforms with time, individuals in states under the old divorce law regime may nevertheless respond to the nationwide changes, e.g. by anticipating similar reforms in their own state. Both factors would bias the coefficients downward. In Section 2.4, I estimate using the model an alternative measure of the effect of divorce laws on college attendance and compare it to the reduced-form estimates.

1.6 Persistent Differences in Choice of Major Since the Mid-1980s

The persistent difference in the choice of undergraduate major documented in Figure 1.2 raises the question why, given their high college attendance rates, women did not converge further with men along this second margin, or similarly overtake them. Before answering this question, one potential concern that needs to be addressed is that the patterns in Figure 1.2 do not account for the change in the weight of different majors over time, and may thus misrepresent the persistence of the gender gap in majors. In particular, some traditionally "female" majors like education became less popular over time, while other majors that were historically "male" and became more gender-equal, such as business, increased in popularity (NCES (2012)).

To address this, Figure 1.6 graphs separately the share of men and the share of women choosing different categories of majors over time using NCES data. For simplicity, majors are aggregated, as in the previous subsection, into two categories: science/business and humanities/other. Figure 1.6 documents two patterns. First, for both men and women the popularity of science/business majors as a share of all degrees increased in the 1970s, although the increase was larger for women, meaning that men and women converged during this period. The share of women majoring in science/business quadrupled from about 10% in 1970 to almost 40% by the the mid-1980s. Men's share increased from roughly 50% to a peak of 68% in 1986, before declining slightly. Secondly, after the mid-1980s the share of men and women choosing a science or business degree has remained roughly stable, at about 60% for men and 36% for women. This implies that convergence virtually ceased after the mid-1980s, as was also documented in Figure 1.2.[7]

To analyze whether persistent differences in choice of major are driven by gender differences in wage premiums by field, Table 1.2 compares the premiums for men and women with different undergraduate majors in 1993, 2003, and 2010, the three waves of the National Survey of College Graduates. The premiums are the coefficients from a regression of log income for full-time workers on a set of dummy variables corresponding to each undergraduate major. The omitted category is a major in the arts and humanities. The coefficient is interpreted to be the "additional" income in log points that individuals in a given major receive, relative to those in the baseline humanities major.

Table 1.2 documents two main findings. Firstly, it documents that men's and women's returns to different college majors exhibit similar patterns. The lowest-paying major for both men and women is education, which pays at least 7% less than a humanities degree in all years, although this difference is not statistically significant in 2010. For both men and women, a social sciences degree provides a very similar return to a humanities degree. Among non-science and non-business

[7]The period of interest in this paper is from 1960 to 2010, but NCES data on these measures begins in 1970. To check whether substantial changes occurred prior to 1970, I use the National Survey of College Graduates, which has data for cohorts that graduated between 1960 and 1970. In the NSCG the share of women graduating with a science/business degree between 1960 and 1970 was almost constant, 15% in 1960 and 16% in 1970. However, the NSCG sample overestimates the share graduating in 1970 relative to the NCES data (10%), which includes the entire population of graduation college students. The NSCG also overestimates the share of men with a science or business degree. It records a temporary decline in the measure over that decade, from 61% to 55%.

majors, degrees in health support fields stand out for their high return, with similar premiums for men and women that range from 14 to 19 log points (13-17%) in 1993 and 2003, and with even higher returns in 2010.[8,9]

The second finding documented in Table 1.2 is that full-time workers with science, engineering and business majors have high premiums relative to those with humanities, education, or social science degrees. Business and math/science majors are associated with an additional return of between 16 and 26 log points (15-23%) for men, and 14 to 20 log points (13-18%) for women. Degrees in engineering and technology are the highest-paying degrees. The additional premium for men is 31-36 log points (27-30%), and for women it is even higher, 38-45 log points (32-37%).

To summarize, these patterns together imply that with the notable exception of nursing/health support, a field that represents about 11% of degrees for women, women are substantially more likely than men to select majors that have on average low expected returns.

1.7 Majors and Flexibility

If women frequently choose lower-paying majors, some other characteristic of these majors should compensate them for the lower return. The popularity of degrees like education and nursing among women suggests that one such possible characteristic is the degree of "flexibility" offered in occupations associated with different majors. I define "flexibility" by high availability of part-time or part-year work and by low wage penalties in case of a temporary absence from the workforce or reduction of weekly hours worked. If women value such flexibility more than men, this may help explain observed differences in choice of major.

Before analyzing how majors differ along this margin, I provide evidence first for why college-educated women today may value such flexibility. Figure 1.7 graphs the share of college-educated

[8] Majors classified under "health" include nursing degrees and other programs related to health support occupations. Bio-med and pre-med majors, which prepare students for medical research or practice, are classified under "sciences." See Appendix A for additional details.

[9] The substantial increase in the return to a health major from 2003 to 2010 could partly be explained by the health industry's strong performance relative to other industries during the recession (Wood (2011)).

men and women employed and the share working full-time at each age between 22 and 50. The data is for the year 2000, but identical patterns hold in 1990 and 2010.[10] Figure 1.7 shows that men and women exhibit similar rates of employment and full-time work immediately after college graduation. However, college women begin to drop their employment rates and their hours worked in their late twenties, and continue to do so through their child-bearing years. Women's overall employment and full-time employment rates reach their lowest point in their mid-30s. At age 35, 60% of college-educated women work full-time, compared to more than 90% of college-educated men. Afterwards, women gradually increase their labor supply again. As Panel B shows, these large reductions in labor supply over the lifetime are primarily driven by women with young children under the age of 6 in the household.

In the rest of the section, I examine whether labor supply patterns differ substantially by the type of major chosen and its associated occupations. I analyze the following measures of flexibility in labor supply, focusing on women with young children: part-time work, employment rates, and annual hours worked, where the latter is a summary composite of the first two measures. After documenting labor supply patterns, I analyze wage penalties by major and occupation group for reductions in labor supply.

Table 1.3 documents NSCG data on college-educated women's employment rates and rates of part-time work by major in 2000. Table 1.4 documents the same measures for men. Part-time is defined as working less than 35 hours per week, as is standard in the literature. All measures are reported separately for individuals with and without children under age 6.

Table 1.3 documents three main findings. The first finding is that, across all majors, women without young children in the household work at high rates, and their part-time work rates are fairly low, although women in humanities and health are somewhat more likely to work part-time. The second finding is that women with young children under 6 reduce their labor supply substantially along both margins, and this is also systematically true across all majors. As expected, Table

[10] In the 1990 Census, 2000 Census, and the 2009-2011 ACS, I observe the same age profiles and overall employment rates. The near-identical patterns across the three decades confirm that the U-shaped employment profile represents systematic differences by age over the life-cycle for women, and is not driven by any cohort effects.

1.4 shows that this is not true for men. Finally, the third finding in Table 1.3 is that there is systematic variation across majors in the degree to which women with young children reduce their labor supply. This variation is most apparent for the recorded rates of part-time work. By a large margin, the highest rates of part-time work for women with children under 6 are observed in health and in the humanities, 43% and 31% respectively. Education majors stand out for their low part-time work rates (20%) among non-science, non-business majors; however, the part-time work measure does not capture the part-year nature of work in teaching professions.[11] Differences in employment rates across majors for women with young children are somewhat less clearcut, but suggest that women in science and engineering have fairly high employment rates, relative to most other majors. Interestingly, women with young children who are health majors have even higher employment rates than engineers with young children; however, this appears to be be driven by the high availability of part-time work in the health field.

While the rates of part-time work for science/business majors are low compared to other majors, at 21-23% they are not insignificant. The objective of the next part of the analysis is to understand which women in these fields are the most likely to work-part time or reduce their employment rates. To do this I divide women not just according to major, but also according to the occupation they work in: science/business occupations vs. all other occupations.

Figure 1.8 graphs part-time rates by major and occupation. Panel A focuses on science/business majors and graphs the share of women working part-time by age. The figure shows that women who work part-time work primarily in a non-science, non-business occupation. Among women with science/business majors who work in a science/business occupation, the part-time work rate is only around 5%. By contrast, women in non-science, non-business occupations work part-time at much higher rates, with up to 16% working part-time in their mid-30s. Panel B of Figure 1.8 shows that a similar pattern holds for humanities/other majors.

Similarly to part-time rates, the observed time taken off from work in Table 1.3 also may vary

[11]NSCG data shows that education majors with children under 6 work about the same number of annual hours as health majors, who work part-time at high rates, 1,237 and 1,253, respectively. The difference is not statistically significant.

systematically with occupation. The occupation-major analysis here is less straightforward, since there is an important inter-temporal aspect of occupation decisions, namely that occupations may change after a leave from the labor force.

To address this, I use panel data from the NLSY79 to assign individuals to one of three occupational groups. Individuals are assigned to the first group if their current or most recent occupation is in science/business and they work again in a science/business occupation within the next 6 years. Individuals are assigned to the second group if their current or most recent occupation is in science/business, but they are not employed again in a science/business occupation in a subsequent survey wave within the next 6 years. Finally, individuals are assigned to a third group if they currently work in a non-science, non-business occupation. Note that the first two groups allow me to distinguish between women in science/business who stay in or re-enter their occupation, as compared to women who were in such an occupation but leave.

Table 1.5 records the labor supply for each of these three occupation groups, separately by major. For conciseness, the table lists annual hours worked. This allows me to analyze simultaneously multiple margins of flexibility, including employment, part-time work, and part-year work. If an individual did not work in the past year, the hours worked for that individual are recorded as zero.

The evidence in Table 1.5 shows that labor supply varies substantially more across the three groups, when occupations and occupational transitions are accounted for in this way. Among science and business majors, women with young children who ultimately stay in a science/business occupation work more than 1,800 hours annually, an average of 35 hours per week across all women in this group, equivalent to working full-time. By comparison, women with young children in the other two occupation groups worked 961 and 1,272 hours. The patterns for humanities/other majors by occupational group are similar to those for science/business majors.

The evidence provided above suggests that science/business occupations are less flexible relative to other occupations when it comes to reducing labor supply, and that this is especially relevant for women with children. As might be expected based on these patterns, the share of women in science/business majors who actually work in a science/business occupation is substantially lower among women than men. Table 1.6 shows that a science/business degree is a very strong predictor of working in a science/business occupation for men. In 2000, 81% of men who were science/business majors worked in their related occupation. By contrast, this was true for only 57% of women with a science/business degree.

An important side-note to the previous tables is that they focus on women with and without children under 6, and do not take into account that there may be potential differences across majors in the overall share of women with a child under the age of 6 in the household. For example, if women with science and business are less likely to have a child under the age of 6 in the household, they may also be less likely to want "flexibility." However, Figure 1.9 suggests that this is not the case. Figure 1.9 graphs by major the share of college women who are married (Panel A) and the share who have a child under the age of 6 (Panel B) in 2000. The figure shows that there are virtually no differences in these two measures for women across the two types of major.[12]

A final measure of flexibility is the size of the wage penalty associated with part-time work and/or time taken off from the labor force. To measure how such penalties differ by major and occupation, I use NLSY79 panel data to run the following fixed-effects regression separately by major and occupation:

$$\ln w_{i,t} = \alpha_i + \beta_1 exp_{i,t} + \beta_2 exp_{i,t}^2 + \beta_3 \mathbb{I}[\text{Part-Time}]_{i,t} + \beta_4 \mathbb{I}[\text{Time Off}]_{i,t} + \varepsilon_{i,t}, \qquad (1.3)$$

The independent variables include an individual fixed effect, a polynomial in experience, an indicator variable for whether or not the individual worked part-time in the current or the previous year, and another indicator variable set equal to one if the individual left the work force for more than 9 months in the last two years.

Table 1.7 reports the coefficients on this regression by field in columns 1 and 2. In line with the previous evidence, the coefficients corresponding to part-time work and time taken off are somewhat larger for science and business majors. For women with those majors, the penalty for

[12]Interestingly, there are some differences for men. Men with science/business majors are somewhat more likely to marry earlier and have children earlier.

taking time out of the labor force is twice as large.

In the remaining four columns of Table 1.7, I record results from fixed effects regressions that additionally allow the specifications to vary with the starting occupational status, based on an individual's primary occupation between ages 26 to 30. This captures the penalties incurred for individuals who began their careers in a particular occupation. The age 26 to 30 was chosen based on the age profile for occupational transitions, which indicate that the majority of occupational transitions into science/business occupations occur by age 30.[13]

The results show that for women whose initial occupation was science and business (columns 3 and 4), the penalties for working part-time or taking time out of the labor force are substantially higher than for women in all other occupations (columns 5 and 6). In science and business occupations, the penalties for reductions in labor supply are around 16% for humanities/other majors, and between 18 and 21% for science/business majors. In non-science, non-business occupations, penalties are smaller. Penalties for part-time work are between 8 and 11%, and penalties for time taken off are around 3%. Note that the penalties estimated in column 1 are much lower than those in column 3 because a substantial portion of women with science-business majors ultimately select into a non-science, non-business occupation.

To conclude, the various measures of "flexibility" documented in this section show that majors and their associated occupations differ significantly in the degree to which they facilitate temporary reductions in labor supply, and that women incur high wage penalties in science/business occupations for time taken out of the labor force or for working part-time. The documented patterns suggest that these differences in flexibility may be a key factor explaining why, among women with science/business majors, in 2000 only about half worked in a high-paying field related to their major. Finally, the results in this section show that women are more than 1.5 times as likely to select majors that are associated with occupations with low wage penalties for reductions in labor supply, and higher availability of flexible work arrangements, like part-time work.

[13] Analysis is based on data in NLSY79. I provide evidence on this age profile pattern in Section 2.4. See also Figure 2.4.

1.8 Conclusion

Women today make up the majority of college students in the U.S. and in almost every developed country around the world (OECD, 2012). At the same time, in all of these countries today women select systematically into very different majors than men (Vincent-Lancrin (2006)). In the U.S. there has been almost no convergence between men and women in choices of major since the mid-1980s, even as women caught up to and rapidly outpaced men in graduation rates. In 1985, men were about 1.5 times as likely as women to select high-paying science and business majors, and the same is true today.

I provide evidence that two factors help explain gender differences in college attainment and choice of major for those in college: first, college degrees provide insurance against very low income for women, especially in case of divorce; second, majors and their associated occupations differ substantially in the degree of "work-family flexibility" they offer, such as availability of part-time work and the size of wage penalties for temporary reductions in labor supply. Both components of this explanation are related to the gender gap in wages. Because women draw from a lower wage distribution, they are more likely to seek insurance against very low income realizations when they are single or divorced, and more likely to specialize at least partly in non-market work when they are married. Their educational choices reflect this.

Appendix A: Classifications of Majors

The level of detail available about majors varies across datasets. The analysis categorizes majors as consistently across datasets as possible, but small differences remain. Note that in all the datasets, I assign individuals with pre-med majors preparing for advanced medical degrees with biology/pre-med majors, rather than health/nursing majors. I summarize the classifications used in the three main datasets below.

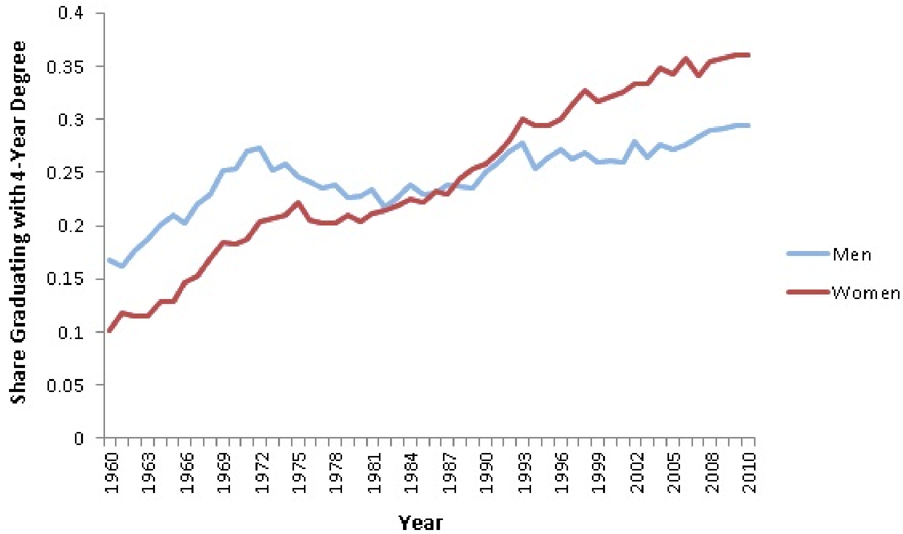
National Center for Education Statistics. The series for each major in Table 1.2 are constructed based on data from NCES Digest of Education Statistics, Tables 343-365. Majors are classified as follows. Hard science/engineering: computer and information sciences, engineering and engineering technologies, mathematics and statistics, physical sciences and science technologies, agriculture and natural resources, architecture and related. Biology: biological and biomedical sciences. Business: Business. Social Sciences: Social sciences and history, public administration and social services, communication, psychology. Humanities/Arts: English language and literature/letters, foreign languages and literatures, visual and performing arts. Health: health professions and related programs. In Table 1.6 the first three categories constitute the science/business category, and the remaining majors constitute sciences/other.

National Survey of College Graduates. The categories in Table 1.2 are constructed as follows in the year 2000 based on NSCG documentation. Sciences: Mathematical sciences, agricultural and food sciences, biological sciences, environmental life sciences, chemistry, earth science, geology and oceanography, physics and astronomy, other physical sciences. Engineering: Aerospace and related engineering, chemical engineering, civil and architectural engineering, computer and information sciences, electrical and related engineering, industrial engineering, mechanical engineering, other engineering, technology and technical. Business: Economics, management and administration, sales and marketing. Social Sciences: Political and related sciences, psychology, sociology and anthropology, social service and related, other social sciences. Health: Health and related. Education: Science and mathematics teacher education, other education. Humanities: Art and humanities, other. For health majors, I additionally use the more fine-grained classification variable available in the survey and assign all individuals who majored in medical preparatory programs to the sciences category, to distinguish them from individuals with nursing or health support majors. Those who major in health services administration are assigned to the business category. When majors are aggregated, science, engineering and business constitute the science/business category, and all remaining majors belong to the humanities/other category. In the 1993 and 2010 waves, the classification is almost identical.

NLSY79. The analysis using the NLSY79 groups majors into two categories, science/business, and all other majors. They are grouped as follows. Science/business: agriculture and natural resources, architecture and environmental design, biological sciences, business and management, computer and information sciences, engineering, mathematics, military sciences, physical sciences, and selected interdisciplinary (biological and physical sciences, engineering and other disciplines). Humanities/other: area studies, communications, education, fine and applied arts, foreign languages, health professions (except pre-med), home economics, law, letters, library science, psychology, public affairs and services, social sciences, theology, and selected interdisciplinary (general liberal arts and sciences, humanities and social sciences, recreation, outdoor recreation, counseling, other).

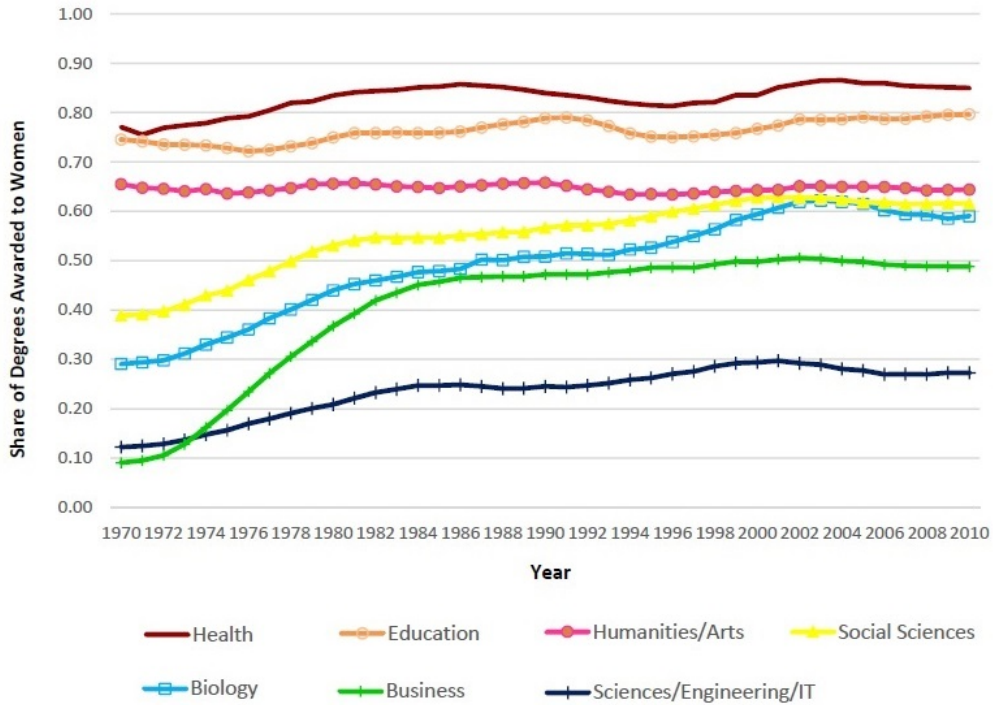
Tables and Figures

Figure 1.1: Share of Men and Women Graduating with 4-Year Degree, 1960-2010



Notes: Sample includes individuals ages 24-30. Graduation year is the year individuals were 22 years of age. Graduation rates after 2008 are constructed using NCES data. Sources: CPS (1962-2012), NCES (2012).

Figure 1.2: Share of Bachelor's Degrees Awarded to Women By Major, 1970-2010

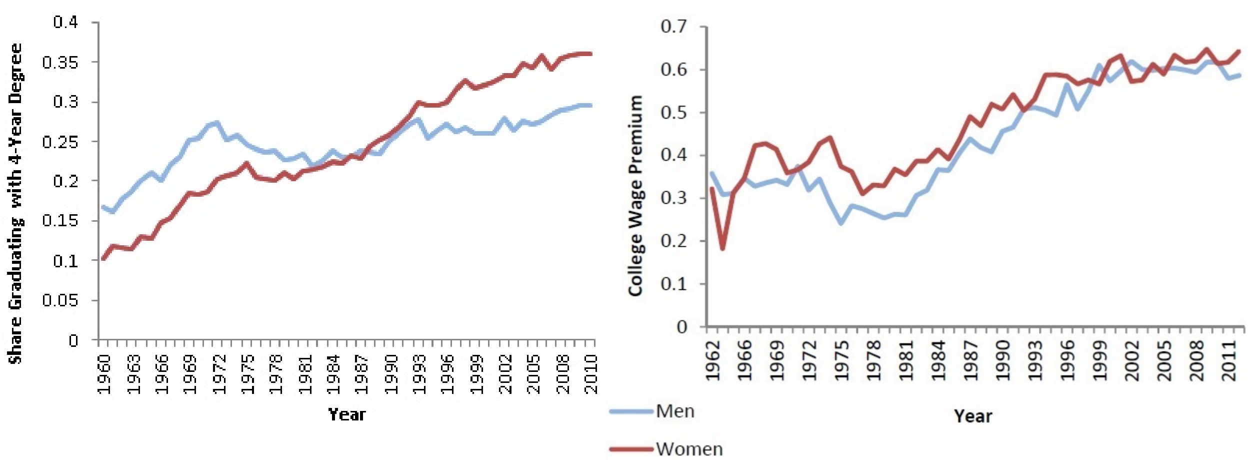


Notes: Data contains full universe of students graduating from accredited U.S. colleges. See Appendix A for details about the classification. Source: NCES (2012), Tables 343-365.

Figure 1.3: Share Graduating and College Wage Premium, 1960-2012

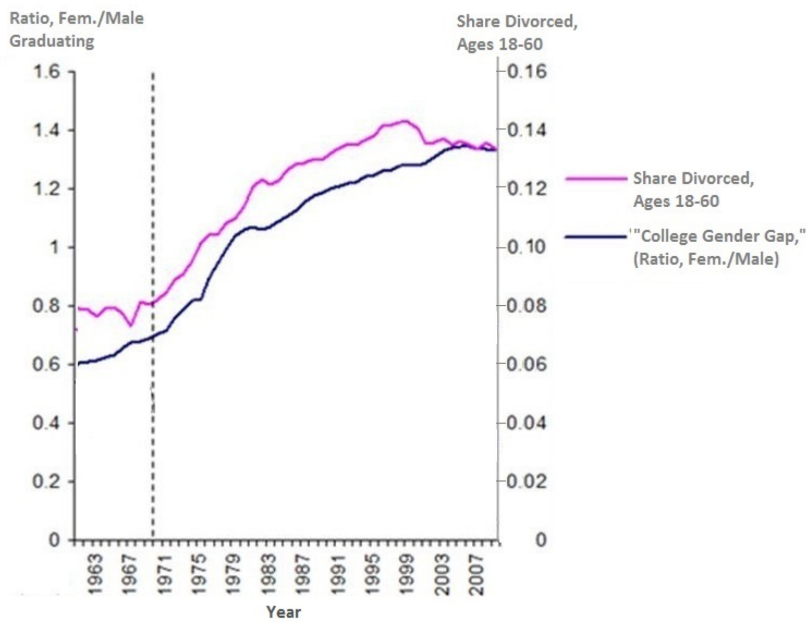
A. Share Graduating from College

B. College Wage Premium



Notes: For Panel A, see notes in Figure 1.1. In Panel B, the sample includes all individuals ages 25-50, who worked 35+ hours in the past week and 48+ weeks in the past year. The wage premium is the difference in log income between college and high school graduates. Source: CPS (1962-2012).

Figure 1.4: Share Divorced and the Ratio of Women to Men Graduating, by Year

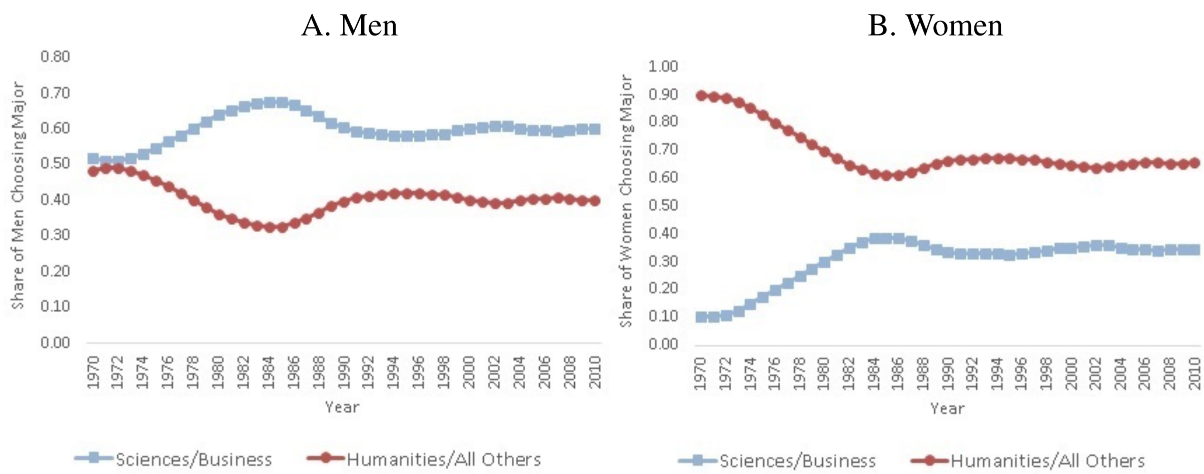


Notes: Share divorced refers to the share of all individuals between 18 and 60 divorced in each calendar year. The college gender gap is the ratio of women to men graduating that calendar year. Source: IPUMS CPS, 1962-2010; NCES, 1960-2010.

Figure 1.5: Coefficients from Regression of Gender Gap on Age at Time of Divorce Law Reform

Notes: Robust standard errors. Additional controls include state and cohort fixed effects. Dotted series represent +/- one standard error.

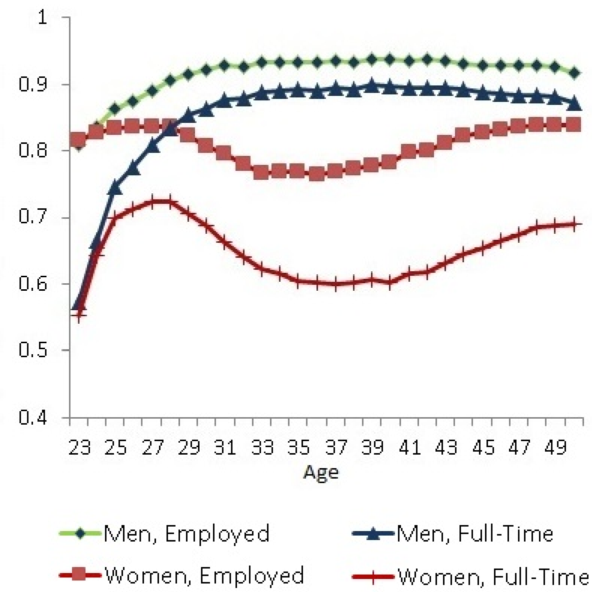
Figure 1.6: Share of Men and Women Choosing “Sciences or Business” vs. Other Major, 1970-2012



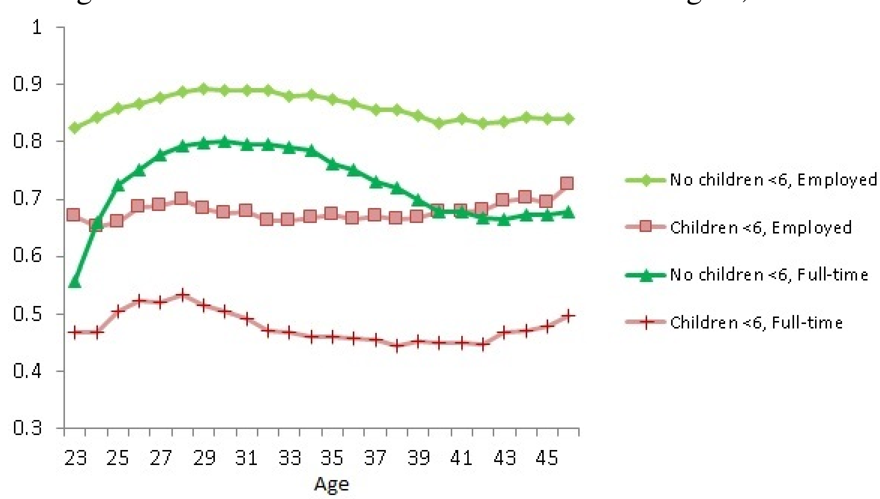
Notes: Data contains full universe of students graduating from accredited U.S. colleges. Shares sum to one in each year. See Appendix A for details about the classification of majors. Source: NCES (2012), Tables 343-365.

Figure 1.7: Share Employed and Share Working Full-Time By Age, College Graduates

A. College Men and Women, 2000

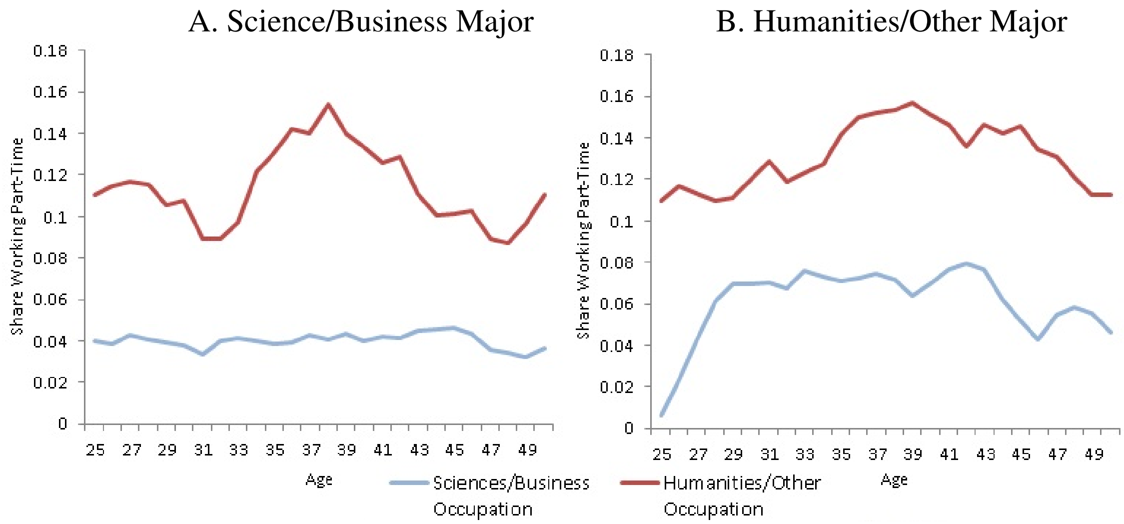


B. College Women With and Without Children Under Age 6, 2000



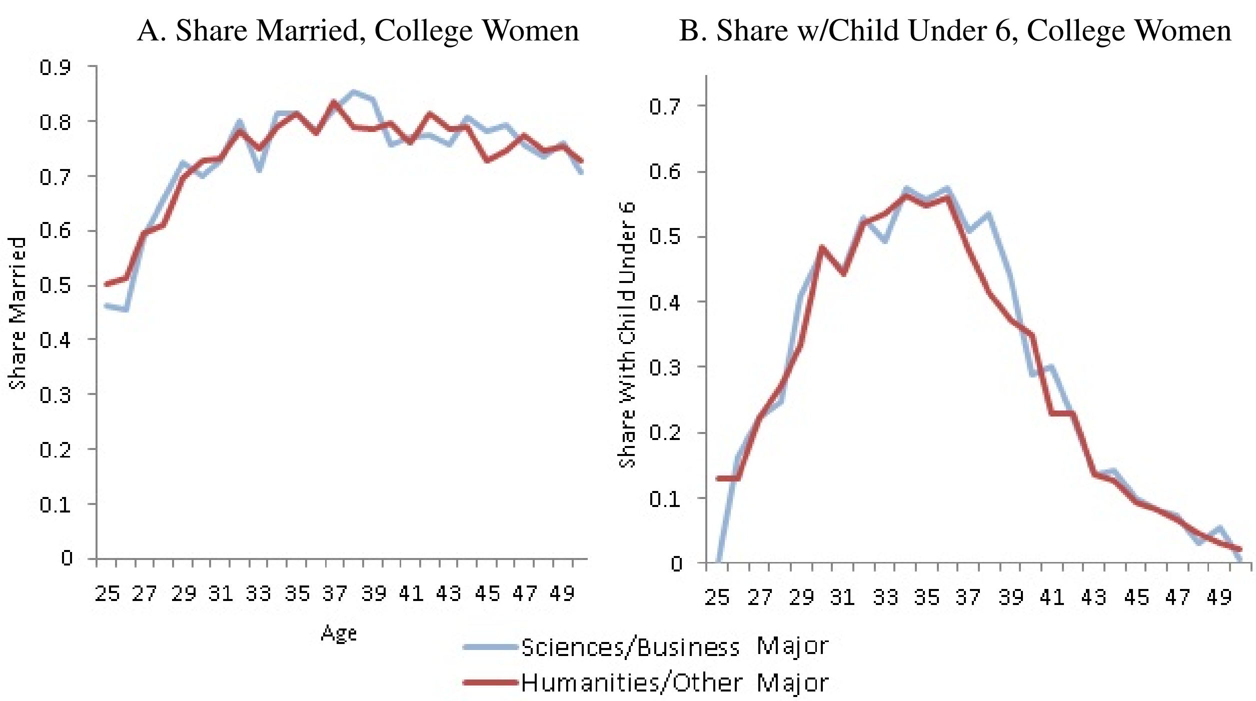
Notes: Full-time is defined as working at least 35 hours per week. Sample includes college graduates only. Source: IPUMS USA Census (2000).

Figure 1.8: Share Working Part-Time, By Occupation and Major



Notes: Part-time is defined as working less than 35 hours per week. Sample includes female college graduates who are currently employed. Source: NSCG (2000).

Figure 1.9: Share Married and Share With a Child Under Age 6, College-Educated Women by Major (2000)



Notes: Sample includes all female college graduates. Source: NSCG (2000).

Table 1.1: Effect of Divorce Law Reforms on Gender Gap in Choice of College Major

	Dep. Var.: Share Majoring in Science/Business (Women-Men)
2 to 0 years until divorce law reform	0.003
	(0.004)
1 to 3 years since divorce law reform	0.005
	(0.004)
4 to 6 years since divorce law reform	0.008**
	(0.004)
7 to 9 years since divorce law reform	0.009*
	(0.005)
More than 10 years since divorce law reform	0.002
	(0.008)

** Significant at 5%. * Significant at 1%. Source: NCES/HEGIS.

Table 1.2: Regression of Log Income on Field of Undergraduate Major, All College Graduates

	Women			Men		
	1993	2003	2010	1993	2003	2010
Sciences	0.137***	0.138***	0.196***	0.160***	0.224***	0.246***
	(0.011)	(0.022)	(0.042)	(0.009)	(0.020)	(0.039)
Engineering	0.381***	0.375***	0.452***	0.306***	0.317***	0.355***
	(0.017)	(0.037)	(0.046)	(0.009)	(0.019)	(0.040)
Business	0.164***	0.145***	0.202***	0.198***	0.234***	0.257***
	(0.010)	(0.023)	(0.053)	(0.009)	(0.021)	(0.046)
Social Sciences	0.018*	-0.053**	0.053	0.002	0.043	0.068
	(0.010)	(0.023)	(0.042)	(0.011)	(0.027)	(0.046)
Health	0.192***	0.143***	0.299***	0.161***	0.179***	0.276***
	(0.011)	(0.026)	(0.044)	(0.020)	(0.034)	(0.053)
Education	-0.084***	-0.180***	-0.097	-0.082***	-0.151***	-0.070
	(0.009)	(0.028)	(0.061)	(0.012)	(0.031)	(0.068)
Humanities[1]	-	-	-	-	-	-
N	33,111	16,432	13,209	54,992	28,075	20,117

[1] Humanities is the omitted category. * Significant at 10%. ** Significant at 5%. *** Significant at 1%. Coefficients are the outcome of a regression of log wage on a set of indicator variables corresponding to each major. Controls include indicator variables for age, race, and highest degree earned. Sample includes individuals ages 25 to 50 employed full-time, full-year. Robust standard errors in parentheses. Sources: NSCG 1993, 2003, and 2010.

Table 1.3: Women's Employment and Rate of Part-Time Work in Different Majors

	No Children Under 6		Has Child Under 6	
	Employed	Part-Time	Employed	Part-Time
Sciences	0.89	0.12	0.77	0.23
Engineering	0.89	0.10	0.74	0.21
Business	0.88	0.12	0.70	0.23
Social Sciences	0.88	0.13	0.71	0.27
Health	0.91	0.18	0.80	0.43
Humanities	0.89	0.15	0.68	0.31
Education	0.93	0.10	0.72	0.20

Notes: Part-time rates are based on individuals who are currently employed. The sample includes college graduates ages 25 to 50. Source: National Survey of College Graduates, 2003.

Table 1.4: Men's Employment and Rate of Part-Time Work in Different Majors

	No Children Under 6		Has Child Under 6	
	Employed	Part-Time	Employed	Part-Time
Sciences	0.94	0.03	0.97	0.02
Engineering	0.95	0.02	0.97	0.01
Business	0.95	0.03	0.97	0.01
Social Sciences	0.93	0.06	0.96	0.03
Health	0.96	0.04	0.98	0.02
Humanities	0.94	0.07	0.97	0.03
Education	0.94	0.03	0.99	0.01

See note in Table 1.3.

Table 1.5: Yearly Hours Worked in Different Majors and Occupational Groups

	Women		Men	
	No Children Under 6	Has Child Under 6	No Children Under 6	Has Child Under 6
Science/Business Majors:				
S/B Occupation, Now & In Future	2118.9	1816.7	2300.0	2459.5
S/B Occupation, Now Only	1850.0	961.0	2276.4	2354.4
Other Occupation	1730.7	1272.1	1950.0	2205.3
Other Majors:				
S/B Occupation, Now & In Future	2119.5	1696.9	2267.9	2450.1
S/B Occupation, Now Only	1827.8	1074.2	2177.8	2166.1
Other Occupation	1897.5	1299.5	2046.1	2266.8

Notes: S/B refers to science/business. Sample includes individuals ages 25 to 50. Individuals not employed in the current period are assigned the occupation they had in their most recent job. Source: NLSY79.

Table 1.6: Share Working in Science/Business Occupations, By Ages 30-35

	Women	Men
Science/Business Majors	0.574	0.808
Humanities/Other Majors	0.204	0.415

Source: National Survey of College Graduates, 2003.

Table 1.7: Wage Penalties for Labor Supply Reductions, by Major (Women)

	All Occupations		S/B Occupations		Other Occupations	
	S/B Major	Other	S/B Major	Other	S/B Major	Other
Part-Time	-0.126***	-0.116***	-0.178***	-0.162***	-0.083***	-0.108***
	(0.016)	(0.011)	(0.023)	(0.028)	(0.028)	(0.028)
Time Off	-0.105***	-0.043***	-0.212***	-0.159***	-0.031	-0.026*
	(0.020)	(0.014)	(0.031)	(0.039)	(0.028)	(0.028)

*** Significant at 1%. * Significant at 10%. Notes: Individual fixed effects regression. Log wage is the dependent variable. Additional controls for experience and experience squared. Robust standard errors in parentheses. Source: NLSY79.

Chapter 2

A Dynamic Model of Marriage, Education, and the Lifecycle Labor Decisions

2.1 Introduction

In this chapter, I develop and estimate a lifetime model of individual decisions that can capture the empirical features around marriage and divorce, labor supply, and education described in Chapter 1. The empirical patterns documented in Chapter 1 indicate that the insurance value of the college degree in case of divorce and differences in work-family flexibility across majors are important drivers of the gender gaps. In this chapter, I build on the evidence in Chapter 1 and use a dynamic structural model to examine in greater detail the drivers of men's and women's educational investments from 1960 to 2010.

There are several reasons why a dynamic lifetime model of individual decisions is useful for analyzing educational investment decisions. Firstly, with a model it is possible to analyze cumulative processes and interactions, such as the increase in divorce rates over time or changes in labor supply, which can be results of as well as the drivers of educational investment decisions. Using a model, it is possible to account explicitly for the fact that marriage, labor supply, and education decisions are jointly made, and that a change in the choice about one of these factors affects the optimal decision about all the others.

Secondly, it is possible to analyze the relative importance of changes in the wage structure and changes in the marriage market in generating the observed dynamics in educational choices. Finally, it is possible to perform policy simulation exercises. Because policies that seek to increase the share of science-technology majors have been recently proposed (e.g., Alvarez (2012)), performing such simulations to analyze their potential effects is particularly useful.

It follows individuals starting at age 18 over three phases of life: education, work, and retirement. In the first phase, individuals decide whether or not to go to college. If they go to college, they choose between two majors. The first major is associated with occupations that have a high return, but also a high rate of skill depreciation, meaning that individuals incur large wage penalties for any reductions in labor supply. The second major is associated with occupations that have a lower return, but also a lower rate of skill depreciation. These differences between the majors match those observed in the data. Individuals make decisions based on their expected lifetime utility from each educational choice and their unobserved effort cost of completing each major.

In the second, working phase of life individuals make decisions about time allocated to market and home production, marriage and divorce, and savings. If an individual is single, each period he or she is matched with a potential partner and decides whether to marry. If married, the partners make household decisions jointly, but there is no commitment, meaning that if in some period the partners are not both better off in the marriage than they would be if they were single, they divorce (Marcet and Marimon (1992)). There are shocks to marital match quality, as well as to wages and to fertility. After a fertility shock, the presence of a young child in the household increases the productivity of hours dedicated to child care and home production.

The final, retirement stage of life, is a simplified version of the working life stage, in which individuals make decisions only about consumption, home production, and savings. For different cohorts, decisions over the lifetime and therefore expected returns to education are affected by changes in the wage structure and by changes in the marriage market after the reform in divorce laws.

The model is estimated using the Simulated Method of Moments. The estimated model matches well the marriage and labor supply patterns of men and women over the lifecycle, and educational choices of cohorts over time. It generates a reversal in the gender gap in graduation rates, and the persistence in the gender gap in majors. In the model, the current differences in educational choices are generated through the interaction between the gender wage gap and marriage over the lifecycle. On the one hand, the "insurance" value of the degree drives up the return to college for women in case of divorce, regardless of the major they choose. As a result, they graduate at higher rates. On the other hand, conditional on being married, women are more likely to be the lower wage-earners and therefore more likely to specialize in home production and child care, since for the household this is the optimal division of labor.[1] Because of this, women are more likely than men to incur wage penalties for reductions in labor supply. As a result, they select majors that offer more flexibility. I estimate that the share of college women choosing a high return major would increase from 0.34 to 0.45 if wage penalties for reductions in labor supply were equal across occupations. Using historical counterfactuals, I also estimate that around half of the convergence in the gender gap in graduation in the 1970s and early 1980s was generated by the increase in the value of "insurance" that the college degree provides for women in case of divorce. The model implies that the insurance value of the degree for women in case of divorce is equal to around 31% of the wage premium.

In the final part of the chapter, I analyze the effects of different policies on educational choices of men and women. The estimated model is well-suited for this purpose because it can analyze policies' effects on decisions about labor supply, occupational choice, and household formation and dissolution. I study two sets of policies. Firstly, several states have recently proposed policies to encourage more students to choose science and technical majors, such as a differential tuition policy that lowers the cost of technical majors (Alvarez (2012)). I use the model to understand the impact of such a policy and how it may affect men and women differently. I find, counterintuitively, that the differential tuition policy has a large effect on women's choice of major, with more women

[1] This concept is related to Becker's work on intra-household specialization. See Becker (1985, 1991).

switching to technical degrees, and almost no effect on men.

Secondly, women's persistently lower representation in certain majors signals potential frictions in the labor market, with scope for welfare-improving policies. To this end, "family-friendly" policies, like paid maternity leave or part-time work entitlements, have been proposed or enacted in various countries to improve work-family flexibility and encourage gender equality in the labor force. I use the model to analyze the effects of such policies on occupational and educational choices. This is an important question because, as Blau and Kahn (2012) point out, one concern with "family-friendly" policies is that they have theoretically ambiguous effects on women's labor supply, occupational, and educational choice, even absent any potential discriminatory response from employers. My results show that the effects on women's occupational and educational choices differ significantly depending on the policy, with some policies substantially increasing the share of women in science and business majors, while other policies amplify *both* current gender gaps in education.

The rest of the chapter is organized as follows. In Section 2.2, I describe the model. Section 2.3 provides details about the estimation. Section 2.4 summarizes the results. In Section 2.5, I present the outcomes of policy experiments. Section 2.6 concludes.

2.2 Description of the Model

In this section, I first provide an overview of the main features of the model and the set of dynamics that the model can capture. I then provide details about each component of the model, and finally about the decisions of individuals in each period.

2.2.1 Overview of the Model

The model is a dynamic individual lifetime model, where individuals live for T periods, starting at age 18. I first outline the main idea of the model and its four most important features.

There are three phases of life: education, working life, and retirement. The first two phases, education and working life, are the main focus of the model.

In the first phase, young individuals decide whether to make an educational investment. If they do not go to college, they begin the working phase of their life. Education choices cannot be changed once the working life phase begins. If they decide to go to college, they choose between two majors. One major is associated with a high-return, high skill depreciation occupation. In this occupation, there is a high wage penalty for temporary absences from the workforce and for part-time work. The other major is associated with a lower-return, lower skill depreciation occupation. These differences in characteristics between the majors are designed to correspond to differences between science/business majors vs. humanities/other majors documented in the empirical section. This is the first key feature of the model.

In the second, working life phase individuals make decisions about marriage and divorce. If single, an individual meets a potential partner with some probability and decides whether or not to marry; if married, he or she decides whether to remain married or to divorce. Couples cooperate when making decisions but cannot commit to future allocations of resources. Divorce occurs when no reallocation of resources within the household can make both individuals better off married than single. This lack of full commitment allows for divorces to occur in the model as they do in the data and it is the second important feature of the model.

In every period of the second phase individuals also make decisions about labor supply, which is allocated to market, home production, and leisure. During the working life phase, there are three sources of uncertainty–wage shocks, marital match quality shocks, and fertility shocks. Fertility is modeled as an exogenous process, conditional on marital status and age. After a fertility shock, the presence of children has the effect of increasing the productivity of labor in home production, which allows the model to capture the large increase in hours allocated to home production and child care observed in the data after childbirth. This aspect of the model also enables me to capture potential gains to partial or full specialization in market and home production for married couples. This is the the third important feature of the model.

Finally, the model not only follows individuals from a given cohort over the lifecourse, but also simulates different cohorts over time. There are two main sources of variation across cohorts, assumed to be exogenous. One is the distribution of entry wages, conditional on sex and education. The other is the cost of divorce. In particular, there is a one-time drop in the cost of divorce in 1970 corresponding to the beginning of divorce law reforms. This captures that divorce became substantially easier after the reforms, the final important feature in the model. Young individuals in each cohort take into account the change in divorce laws when they form expectations about their own future probability of divorce.

The four outlined features make up the structure that generates the key dynamics of the model, over the lifetime as well as across cohorts. In each cohort, individuals face competing considerations. In some periods, especially when there are young children in the household, married individuals may find it optimal to specialize by having one of the spouses commit substantial time to home production, by partly or fully reducing the labor supplied to the market. It will often, but not always, be optimal for the woman to be the one to reduce her labor supply, since she draws on average from a lower wage distribution. On the other hand, working in the labor market increases human capital, and thus future wages. The spouse that reduces labor supply reduces his or her future labor market prospects. Moreover, depending on current occupation, the spouse who reduces labor supply may incur a high additional wage penalty, in addition to the wage losses from foregone experience. Individuals' decisions about labor supply and education will reflect these competing considerations. Changes in the wage structure and in marriage and divorce patterns over time will in turn affect those considerations.

With this overall idea of the model in mind, we now turn to the specific modeling choices. I first provide details about each component of the model, and then characterize the decisions of individuals in each period.

2.2.2 Preferences

Individuals derive utility from consumption c, leisure l and a household-produced good Q. Q is a privately consumed good (Q^i) if individual i is single, and it is consumed as a shared public good (Q) if the individual is married. Couples additionally derive utility from match quality θ. Preferences are separable across time and across states of the world. In each period, the utility function is assumed to be separable in (c_t^i, l_t^i) and Q_t to simplify the estimation, and to take the following form:

$$u_{single}^i = u(c_t^i, l_t^i) + A\log Q_t^i \qquad\qquad u_{married}^i = u(c_t^i, l_t^i) + A\log Q_t + \theta_t.$$

Following empirical evidence from Attanasio and Weber (1995) and Meghir and Weber (1996) that individual preferences are not separable in consumption and leisure, I assume the following functional form for the subutility $u(c_t^i, l_t^i)$:

$$u(c_t^i, l_t^i) = \frac{({c_t^i}^a {l_t^i}^{1-a})^{1-\sigma}}{1-\sigma}, \qquad \sigma > 0,\ \ 0 < a < 1$$

The last component of utility is match quality. Match quality is assumed to follow a random walk stochastic process, where

$$\theta_t = \theta_{t-1} + z_t, \quad z_t \sim N(0, \sigma_z).$$

2.2.3 Household Technology

The good Q_t is produced within the household using market good m_t, labor input d_t, and number of children n_t. To keep the computation simple, I assume a form for the household good production function that is log linear in the inputs:

$$log(Q_t) = \alpha_{1,t}\log d_t + \alpha_2 \log m_t + \alpha_3 \log(1+n_t), \tag{2.1}$$

where $d_t = d_t^i$ if the individual is single, and $d_t = d_t^i + d_t^j$, i.e. the sum of the husband's and the wife's labor allocated to home good production, if the individual is married. Note that this is equivalent to assuming that the husband's and wife's labor inputs are perfectly substitutable. Children increase the production of the household good directly, to capture that one of the additional potential returns to marriage is having a family with children.

To introduce heterogeneity in the productivity of labor in home production, $\alpha_{1,t}$ is allowed to differ over time and across households. Specifically, the parameter $\alpha_{1,t}$ can take one of five values, which will be estimated using time use data on hours allocated to home production and child care: a value for households without children; two values for households with young children, one for each of two educational levels (high school or college); and two values for households with older children, one for each educational level (high school or college). Labor productivity in home production does not depend on major.[2]

By allowing $\alpha_{1,t}$ to vary with the age and presence of children, it is possible to capture the large difference in labor allocated to home production between households without children, households with children under the age of six, and households with children over the age of six. The reason the parameter is allowed to vary additionally with education in households with children is that it allows the model to capture the systematic differences in time allocated to child care by educational attainment. For example, Guryan, Hurst, and Kearney (2008) document that college-educated women allocate more hours both to the labor market and to child care.

2.2.4 Fertility Process

Children are born according to an exogenous fertility process that depends on marital status, age, and the current number of children. A fertility shock can occur if an individual is married and of childbearing age, which in the model is set to 38 or below. The fertility hazard rates are estimated externally.

[2]Note that spouses' labor in home production is substitutable. In married households, I have to choose which spouse's education will be used to determine the household's $\alpha_{1,t}$ parameter. I assume the education of the woman determines the productivity parameter. I assign productivity based on the wife's education because women tend to supply the majority of labor in home production in the data (ATUS, 2003). This is not a restrictive assumption since spouses have the same educational attainment in the majority of couples in the data.

2.2.5 Wage Process and Human Capital

Wages in the model depend on education, current occupation, and accumulated experience. I provide information about each of these factors first, and at the end of the section I describe in detail how they enter into the wage process.

Education and Occupation

There are three educational choices: (1) high school only; (2) college with a major "L" that provides a premium in low-return, low skill-depreciation occupations; and (3) college with major "H" that provides a premium in high-return, high skill-depreciation occupations. I refer to the occupations described respectively as "$\mathscr{L}$"-type and "$\mathscr{H}$"-type. The two majors (L, H) capture the differences in the data documented in the previous section between science/business vs. humanities/other majors. College individuals then choose whether to work in a science/business occupation ($\mathscr{H}$) or all other occupations ($\mathscr{L}$).

I make a simplifying assumption that individuals with a high school education all work in the same $\mathscr{L}$-type occupation. In the NLSY79, the share of individuals without a college education who work in a $\mathscr{H}$-type (science/business) occupation is less than 8%.

Experience

Individuals who work the equivalent of at least 500 annual hours in a given period accumulate one additional period of experience. Experience in the model is occupation-specific. If an individual decides to switch occupations, he or she loses his or her accumulated experience, and must begin accumulating experience again from zero. Though it would be preferable to keep track of experience accumulated in each occupation, I make this assumption to keep the model tractable. In the NLSY79, I observe that most occupational switches occur before the age of 28, that is before individuals have had the opportunity to accumulate substantial experience, which suggests that the assumption is not highly restrictive. The share of men and women who switch between the two categories of occupations more than once after age 30 is around 12%.

Part-Time Work and Time Out of the Labor Force

Working a minimum number of hours in the model matters for accumulating experience. Additionally, the number of hours worked in a given period is important because it determines whether or not an individual will incur a wage penalty for working less than full-time, full-year. I do not model separately the decision to work part-time and the decision to take time out of the labor force. In the model, only the total number of hours worked in a given period is relevant. If the amount of labor supplied in the current period is equivalent to less than 35 hours per week, the individual incurs a wage penalty in the following period. The size of the wage penalty depends on whether the individual works in a $\mathscr{L}$- or $\mathscr{H}$-type occupation.

Wage Process

Individuals draw from wage processes specific to their sex, occupation, and education. I will describe first the wage process for individuals with a high school education, and then describe the process for individuals with a college education.

Individuals with a high school education draw a wage every period for only one possible occupation. This means that there are a total of two wage processes to be estimated for high school individuals, one for women and one for men. An individual of gender k, experience exp_t, and number of hours h_{t-1} worked in the previous period draws a wage from the following process:

$$\ln w_t = \beta_0^{k,HS} + \beta_1^{k,HS} exp_t + \beta_2^{k,HS} exp_t^2 + \beta_3^{k,HS} \mathbb{I}(h_{t-1} < \overline{h}) + \varepsilon_t^{k,HS}, \tag{2.2}$$

$$\varepsilon_t^{k,HS} \sim N(0, \sigma_{\varepsilon^{k,HS}})$$

where $\overline{h}$ is equivalent to the minimum hours worked for a full-time, full-year worker. The coefficient β_3 on the indicator function $\mathbb{I}(h_{t-1} < \overline{h})$ is the current-period wage penalty incurred for working less than full time in the previous period. If the wage drawn in a particular period falls below a value equivalent to the minimum wage, the effective wage is zero and the individual is not employed. Otherwise, the individual may work at hourly wage w_t.

The structure of the wage process for individuals with a college education is similar, except that college-educated individuals draw wages for up to two occupations, $\mathscr{H}$ and $\mathscr{L}$. An individual who enters the period having last worked in a particular occupation will draw a wage from that occupation with probability one. Additionally, with probability η the individual draws a wage from the second occupation and can choose whether or not to switch occupations. Specifically, an individual of gender k with a given major M who enters the period having last worked in occupation q_{t-1} draws the following wage for occupation $q = q_{t-1}$:

$$\ln w_t = \beta_0^{k,M,q} + \beta_1^{k,M,q} exp_t + \beta_2^{k,M,q} exp_t^2 + \beta_3^{k,M,q} \mathbb{I}(h_{t-1} < \overline{h}) + \varepsilon_t^{k,M,q}, \tag{2.3}$$

$$\varepsilon_t^{k,M,q} \sim N(0, \sigma_{\varepsilon^{k,M,q}})$$

The coefficients are indexed by k and M because the wage processes are estimated separately by sex and major.

If the individual draws a second wage for the other occupation $r \neq q_{t-1}$, the wage is characterized by

$$\ln w_t = \beta_0^{s,M,r} + \beta_3^{s,M,r} \mathbb{I}(h_{t-1} < \overline{h}) + \varepsilon_t^{s,M,r}, \tag{2.4}$$

$$\varepsilon_t^{s,M,r} \sim N(0, \sigma_{\varepsilon^{s,M,r}})$$

since the individual loses his or her accumulated experience after switching. As before, the coefficient on $\mathbb{I}(h_{t-1} < \overline{h})$ is a penalty for working less than full-time. To reflect the data, this estimated penalty will be high in $\mathscr{H}$-type occupations, and lower in $\mathscr{L}$-type occupations.

Allowing for occupational choices in addition to educational choices in the model is important for two reasons. Firstly, it allows the model to capture the fact that only a small share of men but almost half of women with a science/business major work in non-science, non-business occupations, as documented in the previous section.[3] This difference in occupational choices strongly affects the return to a science/business major for women relative to men, and the model should be able to

[3]Similarly, some men and women with humanities/other majors work science/business occupations.

capture this feature of the data. The second reason is that variation in wages across occupations is greater than the variation in wages within occupations in the data. This suggests that accounting for the actual occupational choice is important.

Finally, the modeling decision to allow individuals to draw a second wage only with some probability adds further persistence in the wage process.[4]

In the data individuals that have majors that correspond to their occupation on average earn more in that occupation than those in the occupation who have a different major. These differences observed in the data will be reflected in the estimated wage process parameters. Estimation of the wage process parameters will be discussed in detail in Section 2.3.

2.2.6 Educational Costs

Individuals who choose to go to college incur a tuition cost τ, which is deducted from their assets. Individuals also have major-specific utility costs C_L^i and C_H^i, interpreted to be the individual's ability or effort costs for completing a particular major. This is the only source of unobserved heterogeneity across individuals. Individuals draw C_L^i and C_H^i from normal distributions characterized by parameters (μ_L, σ_L) and (μ_H, σ_H). Men and women have the same distributions of effort costs. Educational decisions can only be made in the first period.

2.2.7 Cost of Divorce

If a married couple wishes to divorce, each individual incurs a one-time utility cost K_t. The cost takes two possible values. K_0 corresponds to the cost of divorce before no-fault, unilateral divorce law reforms. K_1 corresponds to the cost of divorce after such reforms. The change from K_0 to K_1 occurs in 1970, corresponding to the timing of the start of divorce law reforms. I assume that

[4]Individuals switch occupations too frequently relative to the data if they draw a wage from both occupations in every period. One way to address this frequent switching would be to introduce persistent wage shocks, one for each occupation. The drawback to this solution is that it would add two additional state variables per individual (or a total of 4 additional state variables for a married household), which would significantly increase the computational burden. Allowing individuals to draw from a second occupation with only some probability introduces the same element of wage persistence without adding any additional state variables to the analysis.

cohorts did not anticipate the change. The change in K_t can be interpreted as a change in the amount of effort required to secure a divorce, in line with historical evidence. For example, prior to reforms individuals and/or couples resorted in many cases to perjury or providing exaggerated or false testimony to provide fault-based grounds (Herbert (1988)).

2.2.8 Individual Decisions

With all the main components of the model laid out, I now describe households' decisions, starting with the working life stage. Afterwards, I describe the retirement stage, which is a simplified version of the household's problem during the working life stage. Finally, after the description of what happens over the lifecourse, I discuss the educational decision that takes place in the first period. It is best left for last to discuss this decision because it requires knowledge of the expected stream of lifetime utility from each educational choice.

During the working life phase, individuals enter every period t either single or married, and make a decision about their marital status that period. Shocks to wage, match quality and fertility are realized at the beginning of the period, before any decisions are made. An individual who has completed his or her education and enters the period as single meets a potential spouse j with probability one, and draws a match quality θ_t. This individual must now choose whether to stay single or to marry the potential partner, and to make this decision, he or she must compare the value of staying single with the value of marrying the potential partner. Similarly, an individual who enters as married makes a decision between staying married or getting divorced, and to do that he or she compares the values of those two options.

I will first describe the problem of an individual who enters the period as married. If the couple to whom the individual belongs decides to divorce, he or she will experience the value of being single, $V_t^{i,S}$, that can be computed as follows. The individual will choose the levels of own consumption c_t^i, labor supplied to the market h_t^i, labor supplied to home production d_t^i, leisure l_t^i, savings s_t^i, and the amount of the market good m_t^i devoted to home good production that maximizes his or her lifetime expected utility. If an individual receives a wage draw from more than one occupation, the individual additionally chooses which occupation to work in, q_t^i. The value of staying single for the individual is therefore equal to

$$V_t^{i,S} = \max_{c_t^i, l_t^i, m_t^i, d_t^i, h_t^i, q_t^i, s_{t+1}^i} u^i(c_t^i, l_t^i, Q_t^i) + \beta E[V_{t+1}^i(\omega_{t+1}|\omega_t)]$$

$$\text{s.t.} \quad c_t^i + m_t^i + p_k n_t^i = w_t^i h_t^i + R s_t^i - s_{t+1}^i \qquad \text{(Budget Constraint)}$$

$$Q_t^i = F^i(m_t^i, d_t^i, n_t^i) \qquad \text{(HH Good)}$$

$$w_t^{i,q=q_{t-1}} = G^i(ed^i, exp_t^i, h_{t-1}^i, \varepsilon_t^i) \qquad \text{(Wage I)}$$

$$w_t^{i,r\neq q_{t-1}} = G^i(ed^i, h_{t-1}^i, \varepsilon_t^i) \qquad \text{(Wage II)}$$

$$h_t^i + d_t^i + l_t^i = \mathscr{T} \qquad \text{(Time Constraint)}$$

where ω_t is the set of state variables in period t, and $E[V_{t+1}^i(\omega_{t+1}|\omega_t)]$ is the expected value function of the individual when he or she enters period $t+1$ as single.

Now consider the value of staying married, $V_t^{i,M}$, for the same individual that enters the period as married. The value $V_t^{i,M}$ is determined by modeling the decisions of the married household as a Pareto problem with participation constraints. In determining $V_t^{i,M}$, I follow the literature on decisions with limited commitment (e.g. Marcet and Marimon (1992, 1998), Ligon et al. 2000), and in particular its application to models of intra-household allocation (Mazzocco (2007)). This literature shows that the Pareto problem with participation constraints can be solved in two steps. In the first step, the household solves the unconstrained problem. This means that in period t the married couple chooses the vector $z_t = \{c_t^i, c_t^j, l_t^i, l_t^j, d_t^i, d_t^j, h_t^i, h_t^j, q_t^i, q_t^j, m_t, s_{t+1}\}$ to solve the following Pareto problem, with weights μ_t and $(1-\mu_t)$:

$$\max_{z_t} \ \mu_t[u^i(c_t^i, l_t^i, Q_t, \theta_t) + \beta E[V_{t+1}^i(\omega_{t+1}|\omega_t)]] + (1-\mu_t)[u^j(c_t^j, l_t^j, Q_t, \theta_t) + \beta E[V_{t+1}^j(\omega_{t+1}|\omega_t)]]$$

$$
\begin{aligned}
\text{s.t.} \quad c_t^i + c_t^j + m_t + p_k n_t &= w_t^i h_t^i + w_t^j h_t^j + R s_t - s_{t+1} && \text{(Budget Constraint)} \\
Q_t &= F^i(m_t, d_t^i, d_t^j, n_t) && \text{(HH Good)} \\
w_t^{k,q=q_{t-1}} &= G(ed^k, exp_t^k, h_{t-1}^k, \varepsilon_t), \quad k = i,j && \text{(Wage I)} \\
w_t^{k,r\neq q_{t-1}} &= G(ed^k, h_{t-1}^i, \varepsilon_t), \quad k = i,j && \text{(Wage II)} \\
h_t^k + d_t^k + l_t^k &= \mathscr{T}, \quad k = i,j && \text{(Time Constraint)}
\end{aligned}
$$

When the optimal solution z^* to the unconstrained problem is determined, one can calculate

$$V_t^{*k,M}(z^*) = u^i(c_t^{*i}, l_t^{*i}, Q_t^{*i}) + \beta E[V_{t+1}^{*i}(\omega_{t+1}|\omega_t)], \quad k = i,j$$

i.e. the value of being married for spouse k at the current Pareto weights μ_t and $(1-\mu_t)$.

In the second step, one can then check that the solution satisfies both individuals' participation constraints. Recall that if the couple divorces, each partner incurs the one-time utility cost K_t. Hence, the constraints for individuals that entered the period married take the form

$$V_t^{*k,M} \geq V_t^{k,S} - K_t, \quad k = i,j$$

If the participation constraints for both partners are satisfied at $V_t^{*k,M}(z^*)$, the allocations determined in the first stage are the final allocations and the couple stays married. In that case,

$$V_t^{k,M} = V_t^{*k,M}(z^*), \quad V_t^k = V_t^{k,M}, \quad k = i,j.$$

If both constraints are violated, the marriage generates no surplus and the couple divorces. Finally, if only one of the constraints is satisfied for the married couple, there is potential for a renegotiation. I use the result from Ligon, Thomas, and Worrall (2002) that in the optimal solution, the constrained individual's Pareto weight is increased so that the individual is exactly indifferent between staying in the marriage and leaving it. Suppose under this new weight corresponding to

$\tilde{\mu}_t$, the solution to the household's maximization problem is $\tilde{z}^*$. If the other spouse's participation constraint is still satisfied under the solution $\tilde{z}^*$, then the couple stays married. If not, then there is no value of the Pareto weight that simultaneously satisfies the participation constraints of both partners, and the individuals divorce. In that case, V_t^k is the value of being divorced, $V_t^{k,S} - K_t$, for $k = i, j$. For married couples, μ_t constitutes an additional state variable.

Individuals who enter the period as single calculate the value of being single and the value of being married to a potential partner in almost the exact same fashion. However, there are no participation constraints for individuals who enter the period as single. They simply compare their value of being single and their value of being married. If the latter is higher for both individuals, they marry. The initial Pareto weights for a couple that marries are determined using symmetric Nash bargaining.

The problem of the household in retirement is a simpler version of the household's problem in the working life. The only decisions in retirement are about consumption, leisure, savings, and the amount of time and market good allocated to home production. In the final period, all resources are used and savings are equal to zero. Because retirement decisions are not the focus of the model, I simplify the problem by not allowing individuals to divorce or marry in retirement. As a result, individuals who enter retirement married simply solve the married household's unconstrained problem, with the Pareto weights fixed throughout retirement.

Now that I have described the decisions that occur over the lifetime, I can describe the educational decision that takes place in the first period. The decision takes into account expectations about marriage, divorce, occupational and labor supply choices over the lifetime, conditional on each educational choice. The expectations at the time of the educational decision about future lifetime utility from each educational choice is expressed as $E[V_1|\text{ed}]$. Additionally, the decision about education depends on the idiosyncratic utility costs associated with going to college and choosing major L or H. An individual will choose to go to college if either

$$E[V_{1|ed=L}] - C^i_L \geq E[V_{1|ed=HS}], \quad \text{or}$$
$$E[V_{1|ed=H}] - C^i_H \geq E[V_{1|ed=HS}].$$

If neither condition holds, the individual does not go to college and begins the working stage of his or her life. Otherwise, the individual chooses the major that gives him or her the higher return, $\max\{(E[V_{1|ed=L}] - C^i_L), (E[V_{1|ed=H}] - C^i_H)\}$. This individual does not make any decisions about labor supply or marriage for two periods, equivalent to four years in the model. After completing college, the individual begins the working life stage. It is not possible to change one's education after the initial educational decision is made.

2.3 Estimation

In this section I discuss the simulation and estimation of the model. I first provide additional details about matching in the marriage market and assumptions about divorce and children, which are needed to operationalize the model. I then discuss the estimation method.

2.3.1 Implementation Details for Model Simulation

To be able to simulate the model, I must make additional assumptions which were not discussed in Section 2.2 about how individuals meet in the marriage market, about the cost of children in the household, and about how couples split wealth and child custody after divorce. I describe these assumptions and the data patterns they are founded on below.

Individuals in the model draw potential partners from their own cohort. To capture that more than 80% of individuals marry someone with the same educational attainment (IPUMS USA, 2000), I introduce an assumption about how individuals meet in the marriage market.[5] I assume

[5]In 2000, 80% of married individuals between ages 18 and 60 were married to a spouse with same educational attainment as their own, where the two categories of attainment are defined as "college" and "less than college." Note that the share did not change substantially over time. In 1970, 1980, and 1990, the share of individuals with a spouse

that the probability each period that a single individual draws a potential partner with the same educational attainment (high school or college) is equal to p_m, which is estimated in the model. Conditional on drawing a college-educated partner, the probability that the partner has a particular major simply corresponds to the share of individuals of that sex who choose that major.

Spouses in the model pool their savings after marrying. If a married household has children, the household pays a cost per child p_k equivalent to \$6000 per year for the first child and \$4500 for subsequent children, as estimated from the Consumer Expenditure Survey.[6] If a couple divorces, the individuals split their total savings evenly, to reflect available data on asset allocations after divorce. Additionally, I must choose a rule that determines how couples split custody and financial responsibilities for children after divorce. According to the Census Bureau, women represent 85% of all custodial parents. About 45% of custodial mothers receive any kind of child support, and the average amount received for these women is \$3,800 (Grail (2002)). Given the cost of children p_k, the received child support payments cover about 14.3% of the child expenditures in a divorced household with two children. As a result, I assume that divorced women are responsible for the majority of financial costs for the children, 85%, while divorced men are responsible for the remainder.

To simplify the simulation of the model, I assume that if a divorced woman with children remarries, the new household treats the children as its own. If a divorced father re-marries, he does not bring any children into the new household, and I cease keeping track of children from his previous marriage. When an individual with children from a previous marriage re-marries, both new marriage partners pay a re-marriage penalty P_{RM}, a one-time fixed utility cost, which is estimated. This assumption allows me to match the pattern that divorced individuals with children have lower re-marriage rates than those without children (NLSY79).

Since I do not keep track of the age of children, I need an assumption about when children leave the household. I assume that after age 46, a household with children transitions each period

that has the same education were 87.7%, 84.2%, and 81.7%, respectively.

[6] I use estimates of the expenditures per child from Mazzocco, Ruiz, and Yamaguchi (2009), based on data from the Consumer Expenditure Survey (CEX, 1980-1996). I adjust their estimates to be in year 2000 dollar values.

with a one-half probability to a state in which there are no more children in the household. At age 50, I assume that all individuals are without children in the household.

2.3.2 Estimation Method

In this section I discuss the estimation of the model's parameters. The parameters in the model fall into three categories, depending on whether they are estimated within the model, estimated externally, or calibrated using estimates from the literature. The first category consists of parameters that I estimate using the simulated method of moments. This category includes (1) all parameters of the home good technology production function; (2) parameters related to the marriage market, including those governing the match quality process and matching along educational lines, the cost of divorce, and the marriage market penalty for being divorced with children; (3) the probability for college-graduates of receiving a wage draw from more than one occupation; and (4), the distribution of utility costs of education by major. The second category consists of parameters that I estimate externally from the data and use in the model. This category includes all wage process parameters. Additionally, I set financial costs related to children and fertility hazards directly to those I estimate in the data. Finally, the third category consists of parameters that I calibrate using estimates from the literature. The parameters in the third category include the CRRA risk aversion parameter, the Cobb-Douglas parameter for consumption and leisure, and the discount factor.

Note that all of the parameters in the first category except one of the two cost of divorce parameters can feasibly be estimated based only on the lifetime decisions of one cohort. Estimating both cost of divorce parameters requires at least two cohorts, a cohort that made the majority of its marriage and divorce decisions before the reform, and a cohort that made the majority of its decisions after the reform. To make the estimation of the parameters in the model computationally feasible, I take advantage of this fact that most parameters estimated within the model are time-invariant and I conduct the estimation in two steps. In the first step, I estimate all the parameters of the model except K_0 using a single cohort that graduated after the introduction of divorce law reforms. In the second step, I run the model using the estimated parameters from the first step for the cohort that was exposed to the pre-reform divorce regime, and estimate the remaining parameter K_0.

The main estimation (step I) requires selecting a cohort. I choose the cohort that graduated college in 1980, for three reasons. First, it satisfies the requirement that the cohort made its decisions under the new divorce law regime. Secondly, it is possible to follow this cohort almost over the entirety of its working lifetime, which is not possible with the more recent cohorts graduating in 1990 or 2000. Finally, NLSY79 panel data is available for this cohort, which has detailed microdata about majors, occupations, and labor supply over the lifetime.

I estimate the main parameters of the model, i.e. those in the first category, using the Simulated Method of Moments (McFadden (1989)). I solve the model recursively following Keane and Wolpin (1997) and use it to generate an artificial dataset of choices about labor supply, marriage, etc. I then construct moments based on this simulated data. The estimation method chooses structural parameters that minimize a weighted average distance between a set of data moments and the corresponding moments simulated from the model.

The moments used in the estimation are as follows. The first group of moments includes a set of labor supply moments that correspond to annual hours worked by sex, education, occupation (for college-educated individuals), and family structure (single, married without children, married with children under age 6, and married with children ages 6 to 16). Additionally, I estimate the share of individuals in "$\mathscr{H}$"-type occupations by sex and major. These sample moments are estimated using the 2009-2011 ACS and the NLSY79.

The second set of moments is constructed using the American Time Use Survey and describes the average annual hours spent on child care and housework for different groups. Because the American Time Use Survey does not provide information about the major chosen, I construct these moments by sex and family structure for three groups: high school, college in a "$\mathscr{L}$"-type occupation, and college in a "$\mathscr{H}$"-type occupation.

I construct a third set of moments related to the marriage market. I estimate the overall share married and divorced, the share of married individuals that have the same educational attainment as their spouse, and the share of individuals married by 30 for the 1980 graduating cohort. I use the

CPS for these four moments rather than the NLSY because it is a larger sample and provides more precise estimates of these measures. The remaining two moments related directly to the marriage market, the difference in the hazard rate of divorce for individuals with and without children, and the hazard rate of re-marriage after divorce for individuals with children, require panel data and are constructed using the NLSY.

Finally, I use NCES data to construct the last set of moments corresponding to the share of men and women going to college, and the share of male and female college graduates choosing a science degree.

I will now discuss the intuition behind the identification of the parameters, starting with the marriage market. The share of married individuals with the same education as their partner allows me to identify p_m, the probability that an individual is matched with a partner who has the same educational attainment. The difference in the re-marriage hazard between divorced individuals with and without children identifies P_{RM}, the re-marriage penalty for divorced individuals with children. The remaining marriage market moments—the overall share married, the share ever married by 30, the share divorced, and the difference in the divorce hazard rate for couples with and without children—jointly identify the match quality process and the cost of divorce. They also contribute to the identification of the parameter determining the direct contribution of children to the home good, since the ability to have children induces single individuals in the model to marry more frequently early in life and less frequently later in life. The labor supply and home production moments are necessary to identify all the remaining parameters of the home good technology function, including the differences in the productivity of labor allocated to home production by educational attainment. The shares of men and women in "$\mathscr{H}$"-type occupations by major are necessary to identify the probability η of drawing a wage from a second occupation.

There are additional moments that could be used in the model estimation, such as marriage and divorce patterns that are specific to the educational groups, since these patterns differ for high school and college-educated individuals in the data. I leave these auxiliary moments as additional tests of the model.

The baseline wage process parameters are estimated using a fixed effect specification with an additional selection term (Wooldridge (2002)). The parameters are summarized in Table 2.1. Table 2.2 summarizes the values for the calibrated parameters. Note that some parameters, such as the discount factor β, are adjusted to take into account that each period in the model corresponds to two years.

2.4 Results

Tables 2.3-2.5 present the estimates of the main parameters, and Table 2.6 summarizes the main marriage, occupation, home production, and labor supply moments of interest in the data and those that are implied by the estimated model. Table 2.6 shows that the estimated model does a good job matching marriage patterns, as well as occupational choices and labor supply differences by gender and education. Since the allocation of hours to non-market work is of substantial interest in the model, Table 2.6 also details the hours supplied to child care and home production by individuals in different educational and occupational groups, focusing on married men and women without any children and with children under the age of 6.

Table 2.6 documents that households with children under 6 spend more than 70 hours in child care and home production in the data, and nearly as much in the model.[7] The model captures that in households with young children, women supply around twice as many hours to non-market work as men. It also captures that relative to other women, those in science/business ($\mathscr{H}$-type) occupations spend less time in child-care and home production, although the model somewhat underestimates their non-market work.

The estimates of the parameters of the home production technology in Table 2.3 provide insight about how the model matches these patterns in home production. There are two main observations in the table. The first is that after having children, estimated home labor productivity increases more for college-educated individuals than for high school-educated individuals. The reason for

[7] Sample data for time spent in home production and child care comes from the American Time Use Survey (ATUS, 2003). I follow Guryan, Hurst, and Kearney (2008) in constructing this measure.

this is that in the data, college-educated women have both higher wages and higher wage penalties for labor supply reductions. Nevertheless, when there are young children in the household, they increase their hours worked in home production and child care more than high school women. As a result, the model implies a high home labor productivity for them. College women's higher productivity in child care is one way to rationalize the empirical findings from Guryan, Hurst, and Kearney (2008) that college-educated women spend more hours on child care than high school educated women, while at the same time also spending more time working in the labor market.

A second observation is that by assumption the parameters in Table 2.3 do not differ for men and women. Nevertheless, hours spent in home production do, especially in households with children. In the model, differences in wages between spouses and the resulting household specialization generate this pattern. Because men's wages are on average higher, it is optimal in most households for the woman to supply relatively more labor to home production. This dynamic is further strengthened in the model by the fact that the household wishes to avoid wage penalties for both partners. As a result, it is optimal for only one individual in the household to reduce working hours below the full-time threshold.

Finally, the other parameters with interesting economic intuition include those that govern the cost of divorce and the utility or effort costs of schooling. Table 2.4 shows that the estimated cost of divorce decreased after reforms, in line with the historical evidence that it became easier to secure a divorce after the reforms in 1970. Table 2.5 shows that effort costs in the model are higher on average for the high-return science/business (H) major than for the lower-return humanities/other (L) major. Note that if the opposite were true, men in the model, for whom the expected return to the science/business major is always higher, would almost never choose the lower-return humanities/other major.

2.4.1 Lifecycle Patterns

Figures 2.1 through 2.3 show the lifetime labor supply, home production, and marriage patterns implied by the model for the baseline cohort graduating in 1980, along with the same patterns in the data. These lifecycle patterns provide good tests for the model's ability to match the data, because they are not matched directly. To match lifetime labor supply patterns, the model has to correctly simulate the timing in marriage and divorce decisions, children, and household specialization; the same is true for home labor production over the lifecycle.

Figure 2.1 shows that the model can capture well the market labor supplied by men and women over the lifecycle, including the substantial decrease in college women's working hours during their thirties. Both in the model and in the data, the decrease is accompanied by a significant increase in women's home production hours over same period of the lifecycle. Figure 2.2 maps simulated home production hours against actual data on home production and child care from the American Time Use Survey. The model captures well the large and rapid increase in home production hours for high school women in their late twenties and a similar increase for college women in their thirties. The model estimates a slightly earlier reduction in hours worked in home production relative to the data. It also overestimates somewhat labor supply to the market at the end of the lifecycle for both men and women. Otherwise, it captures well the lifecycle dynamics in time allocated to market and to home production, as well as difference in these patterns across sex and educational groups.

Figure 2.3 graphs lifecycle marriage and divorce patterns. In the estimation, I do not construct separate marriage and divorce moments for high school and college-educated individuals. Matching these differences across educational groups therefore constitutes an additional test for the model. Figure 2.3 shows that the model captures that high school educated individuals marry earlier, although this result is mechanical, since individuals who choose to go to college in the model enter the marriage market only after they complete their education. However, the differences in divorce patterns provide a validity test for the model, and the model correctly predicts the higher share divorced among high school-educated individuals. The reason for this is that college-educated couples in the model have higher marital surplus, both because they have higher productivity of home labor, and also because they allocate more resources to the market good component of the public home good.

Finally, Figure 2.4 and Table 2.7 focus on occupational choices and choices of major for men and women. Figure 2.4 shows that the model captures well that women enter into science/business ($\mathscr{H}$-type) occupations at a significantly lower rate than men, regardless of major. This lifecycle pattern is driven partly by the fact that both in the model and in the data many women do not enter into the low-flexibility science/business occupation in the first place, even if they have the science/business major.

Table 2.7 shows how these patterns in labor supply and occupational choices generate the expected returns to different majors for women and men. The table records the share of individuals choosing each major in the 1980 cohort and, for expositional purposes, also the model's implied expected return in terms of discounted lifetime utility for each major for women and for men. The purpose of showing the latter is to provide intuition for how the model generates the different shares of men and women going to college and choosing different majors, since individuals make their educational choices partly based on these expected overall returns in lifetime utility, and partly based on their own effort costs for each major.

Table 2.7 records the expected utility returns in two steps. First, it records the implied return to the humanities/other (L) major relative to going to high school. This return is higher for women. Next, it records the *additional* return to a science/business (H) major relative to a humanities/other (L) major. This return is substantially higher for men than for women, although it is positive for both sexes.

The reason that the model implies a high expected return to the humanities/other major for women is that even though the premium for this type of major is relatively low, it increases women's consumption and therefore their utility substantially in periods when they are relying only on their own wages, which are low compared to men's. This implies that not just the size of the premium matters for the returns to college, but also its interaction with the level of wages. Because women have a high return to a major that has low effort costs, the model implies that a lot of women go to college.

Next, the model implies a much lower additional return to science/business majors for women than for men for two reasons. Firstly, women are more likely to incur the high associated wage penalties in science/business occupations than men, since they are more likely to reduce their labor supply over the lifecycle. Secondly, as captured in Figure 2.4, women are less likely to enter science/business occupations in the first place, even if they have a science/business major. Both factors reduce the additional benefits to a science/business major. As a result, the share of women choose such a major is low. On the other hand, men in the model do not incur wage penalties and select at high rates into the science/business occupation. As a result, the share of men choosing the science/business major is high.

2.4.2 Patterns Across Cohorts

In addition to labor supply and marriage dynamics over the lifecycle, the dynamics in educational choices over time are also of central interest in the paper. Recall that there are two changes over time in the model: changes in wages and the one-time change in the cost of divorce.

Figure 2.5 shows that the model captures the main dynamics in college attendance and choice of major across cohorts with only these two sources of variation. Firstly, the model captures that men's decisions about college attendance follow closely the changes in the wage premium (Panel A). This reflects that the premium is the main driver of men's college attendance decisions over time. Secondly, the model also successfully replicates the consistent increase in women's college graduation rates and the reversal in this gender gap, even though it predicts a slightly earlier reversal, around 1980. Finally, the model also correctly replicates the persistence in the gender gap in choice of majors (Panel B).[8] In the remainder of this section, I use counterfactuals to illustrate how changes in divorce laws affect these patterns.

Change in Cost of Divorce in 1970

In this section, I conduct a counterfactual in which I vary the cost of divorce (K_0, K_1) for the 1970

[8]Note that there is an observed peak in the data in the share of men choosing science/business majors in 1980. This is also generated in the model. The reason for this peak in the model is that the college wage premium was on average very low in 1980. As a result, the men choosing to invest in college in 1980 were predominantly those with a relatively low effort cost for the higher-paying science/business major.

cohort. The analysis has two main objectives. The first objective is to understand how changes in divorce laws affected the decisions of cohorts when they were initially passed, in the early 1970s. The second objective is to compare the size of this effect with the effect implied by the reduced-form analysis in Section 1.4.

To conduct the counterfactual, I simulate the 1970 graduating cohort twice. In the first simulation, individuals in the cohort do not anticipate that there will be a change in the cost of divorce from K_0 to K_1, as in the baseline results in Figure 2.5. In the second simulation there is a change in the law, and individuals correctly anticipate it when they make their educational choice. The difference in educational decisions in the two simulations gives me a measure of the net effect that divorce law reforms have on educational choices of cohorts in the early 1970s. This measure can then be compared with the quasi-experimental reduced-form estimate of the same effect from Section 1.4.

The reason I conduct this particular counterfactual is the following. Note that it is not possible to run a cross-state difference-in-difference regression as in Section 1.4 using simulated data from the model. The reason is that for computational reasons I only consider cohorts graduating in decennial years in the model, and only for the aggregate U.S. population. The 1970 cohort simulated under two different divorce laws, however, provides both a "control group," which did not experience divorce law reforms, and a "treatment" group which did, similar to the quasi-experimental reduced-form design. I focus on the 1970 cohort because the majority of reforms across states occurred in the early 1970s, with around half of the U.S. population affected by reforms by 1974 (Friedberg (1998)). As a result, most of the identification in the cross-state experiment relies on changes in reforms that occurred in the early 1970s.

The results of this exercise imply that the change in the cost of divorce had the effect of increasing the share of women going to college in 1970 by 2.8 percentage points, from 18% to about 21%. This is a substantial increase that would have reduced the college gender gap observed in the data in 1970 by nearly one-half. The effect on graduation rates is larger than the one implied by the reduced-form coefficient, which is equal to 1.1 percentage points. This is in line with the discussion about measurement error and potential contamination effects in Section 1.5, which are likely to bias the reduced-form coefficient downwards. I discuss additional possible reasons for the difference in the measured effects in the next counterfactual, focused on a present-day change in the cost of divorce.

The model predicts that the additional return to a science/business major in 1970 is relatively small, increasing the share of women in science and business majors by about 8 percentage points. The increase in the share of women choosing this degree in 1980 is generated by the increase in women's labor supply between 1970 and 1980. This is driven in the model partly by changes in wages, and partly by their interaction with divorce law reforms, which lead women to work more to accumulate additional human capital in case of divorce. Given this higher rate of lifetime labor supply, it is optimal in the model for a larger share of women to choose science/business occupations, and as a result science/business majors. College women's labor supply does not further increase after the 1980 graduating cohort, both in the model and in the data. As a result, the model implies that the shares in science/business occupations and majors after this period also remains flat.

High Cost of Divorce in 2010

Next, I consider how a return to a regime with a high cost of divorce would influence men's and women's decisions about college today. I consider this counterfactual for two reasons. Firstly, it is interesting from a policy perspective to know how a more stringent divorce law regime might affect decisions of individuals today. I will show that such a policy would have a counterintuitive effect on educational choices. Secondly, it allows me to measure the difference in educational choices under the two regimes in exactly the same way as in the previous counterfactual exercise, and therefore to comment about how the reduced-form results would generalize to other periods.

The model implies that implementing a strict divorce law regime today would further increase the share of women graduating relative to men. The model implies that an additional 3% of women invest in a college education in 2010 under the strict divorce law regime. Figure 2.6 explains this counterintuitive result by graphing the counterfactual marriage and labor supply patterns in 2010 under a more stringent divorce law regime. First, Panel A shows that there is a negative effect on marriage rates. The model implies that under today's wages, individuals postpone marriage or do not marry at all, to avoid being trapped in a poor-quality marriage. Panel B shows that as a result, more women remain single and on average supply more labor in the early part of their lifecycle than they do under the more liberal divorce law regime. The economic intuition for why women go to college at higher rates is that the college wage premium is valuable to women because under the strict divorce law regime they marry less and spend an even larger share of their life outside of marriage.[9] The reason one does not observe this pattern in 1960 in the model despite the same high cost of divorce is driven mainly by women's low wages at the time, and thus by their low options outside of marriage.

Interestingly, the results of the exercise imply that the estimates obtained using cross-state quasi-experimental variation do not generalize to the present period. The findings suggest that when using divorce law reforms as a source of variation one should consider the effects of interactions between wages and divorce laws over time in the analysis, as reforms may yield opposite effects on behaviors like labor supply or education, depending on the time period considered. Finally, the results also suggest another potential reason why the effect of divorce law reform on education around 1970 may be higher in the model than in the reduced-form analysis; namely, part of the identification for the reduced-form coefficient relies on divorce law reforms in later periods, when the net effect of the reforms was smaller or even reversed signs.

To conclude the analysis, I use a final counterfactual to quantify the effect that insurance has on educational choices. To isolate the value of insurance, separately from other returns to college over the lifecycle, I conduct the following exercise. I consider how the lifetime utility of a college-educated women is affected if she draws from a wage distribution of a high-school educated woman after divorce, all else equal. To assign a dollar value to this utility, I calculate the annual lifetime

[9]Note that the model does not allow for alternative household structures. The rate of cohabitation would likely increase under the more rigid divorce law regime today. In as far as the literature has documented that household specialization is lower for cohabiting couples (Gemici and Laufer (2009)), many of the same labor supply patterns as in the counterfactual exercise may still be observed. This is an interesting area for further research.

stream of payments that such a woman would have to receive in order to be indifferent between having the insurance in case of divorce and not having it. The exercise implies that the size of the annual payment is about $8,200. Given that the average additional earnings of full-time college-educated women relative to high-school educated women are equal to $26,670 in 2010 (IPUMS USA ACS, 2010), this is equivalent to about 31% of their earnings premium.

Finally, the results suggest that eliminating "insurance" in case of divorce from individuals' returns would reduce the gap in graduation rates today between men and women by 36%. The model implies that the remainder of the gender gap is driven primarily by two factors. Both factors are directly related to the concept of insurance and women's options outside of marriage. First, the college degree has a high return for never married women. Second, having "insurance," or a high option outside of marriage, raises college women's decision power within the marriage. The model implies that the average Pareto weight for married college-educated women is 0.49, compared to 0.42 for high school-educated women.

2.5 Policy Simulations

In this final section I consider policies that can potentially affect individuals' choices about majors. Two policy-relevant issues related to undergraduate majors are frequently discussed in the popular and business media. One is the issue of potential skill shortages in science and engineering fields; the other is the low share of women in technical majors.[10] The two concerns are related, as women's low participation in such majors contribute to potentially low skill supply in technical fields in the U.S. The structural model developed in this paper can analyze the effectiveness of policies that address these two issues. In this section, I consider two sets of policies: one in which differential tuition is charged for different majors; and a second set in which I consider various "family-friendly" policies, which are geared at improving work-family flexibility.

[10] See, for example Koebler (2011), Carnevale, Smith, and Melton (2011), and Pollack (2013). There are differing opinions as to whether or not a shortage of STEM skills exists in the U.S. See Freeman (2006).

2.5.1 Differential Tuition for Majors

Recently some states have proposed a policy to charge lower tuition for science, technology, engineering, and math majors at state universities, most famously in Florida.[11] The direct objective of the policy is to increase the share of individuals who choose those majors. To analyze the potential effects of this type of policy on educational choices, I conduct an experiment in which I reduce the tuition cost for Type H (science/business) majors in the model by one-third. This is roughly in line with Florida's early policy suggestions.[12]

Table 2.8 summarizes the changes in educational choices under this policy. The model predicts that substantially more women choose the less expensive H (science/business) major, and it predicts almost no effect of the policy on men. The reason for this potentially surprising result is that in the baseline model, many women are on the margin between choosing the two types of majors, but men are not. To see why, recall that in the baseline model the science/business major provides a slightly higher expected return in lifetime utility for women than the humanities/other major, but that many women nevertheless choose the latter because it has a substantially lower effort cost. By contrast, the return to the science/business major for men is sufficiently high in the baseline model that almost all men who have a reasonably low effort cost for the major will choose it over the humanities/other major. In the policy experiment, when tuition for the science/business major is reduced, many of the women who were previously on the margin between the two majors are now induced to switch to the lower-cost major. Meanwhile, the men in the baseline model who choose the humanities/other major are relative outliers, who have either a very high cost for science/business majors or extremely low cost for humanities/other majors. The change in tuition does little to induce them to switch majors.

These results suggest that a differential tuition policy could induce women to switch majors, but would mostly subsidize men who would have chosen the majors of interest even without the

[11]Governor Rick Scott of Florida has been a high-profile advocate of a differential tuition system. See Alvarez (2012). Governors in several other states have praised the idea. Providing scholarships to individuals with science and technology majors is an alternative way to implement the policy and to extend it to non-state schools.

[12]However, note that the H-major category also includes business, which is not targeted by the Florida policy.

lower tuition.[13] An additional concern is that about half of women with science/business majors do not choose to work in science/business occupations, which reduces the policy's effectiveness at increasing the supply of science skills to the labor market.

2.5.2 Family-Friendly Policies

One potential way to affect women's educational choices is to address the issue of work-family flexibility directly, with policies that can make it easier for women to allocate labor between market and non-market work. In many OECD countries, "family-friendly" policies include paid parental leave, part-time work entitlements, and subsidized child care (OECD (2010)). By comparison, the U.S. has limited policies around work-family flexibility.[14] In the remainder of this section, I analyze the potential effects of family-friendly policies on household labor supply and on the occupational and educational choices of women and men.

One concern with "family-friendly" policies is that employers may respond in the long run by discriminating against women, a general equilibrium response that this model cannot evaluate. However, another concern with family-friendly policies is that they have theoretically ambiguous effects on women's labor supply, occupational, and educational choice, even without employer discrimination. For example, family-friendly policies may encourage more women to stay in the labor force or, in the case of a maternity leave policy, to return to their original positions after an absence from the work force. On the other hand, they may also have a negative impact on women's accumulated experience and thus on their labor supply and occupational choices over the lifecycle.[15] To evaluate the direct effects of such policies on women's labor supply, occupation,

[13] It is still possible that within the science/business (H) category, some men could be induced to switch majors using the policy, e.g., from business to engineering. However, given that the returns to engineering for men are significantly higher than those for business (see Table 1.2), it is likely that the same intuition applies even within the science/business category.

[14] Currently, the federal Family and Medical Leave Act of 1993 requires large employers in the U.S. with more than 50 employees to provide 12 weeks of unpaid family leave, at the end of which the worker may return to the previous position.

[15] See, for example, Blau and Kahn (2012). Cross-country comparisons in that paper suggest that relative to countries with stronger work-family flexibility policies, women in the U.S. are less likely to work overall, but more likely to work as managers or professionals.

and education choices, a household model that incorporates specialization and optimal household responses to such policies is necessary.

In this set of experiments, I consider three policies: a paid maternity leave policy, a policy that entitles all workers to a part-time work opportunity with their employer, and a subsidized child care policy.

Paid Maternity Leave

In this first experiment, I consider a maternity leave policy that provides paid extended leave for women in the model for one period, i.e. up to two years, immediately after a positive fertility shock. The amount of paid leave provided is based on the woman's earnings in the period prior to the leave. It is equivalent to her hourly pay times her hours worked that prior period, up to but no more than the full-time equivalent of 35 hours per week. Women who choose to take the paid leave may return to their previous position, meaning that in the model they do not suffer the "wage penalty" for having reduced their labor supply in the prior period. However, they will not accumulate experience for that period if they work less than the minimum number of hours required to gain experience.

The series indicated by the blue line in Figure 2.7 records women's lifecycle labor supply response to such a policy. The series indicated by the red line records labor supply in the baseline model without the policy. Because the implied effects for L and H majors are very similar, I only graph the labor supply response for the latter.

The model implies that there are three main effects on women's labor supply. First, women predictably decrease their labor supply after childbirth. Since the maternity leave is paid, the policy essentially constitutes an indirect tax on the woman's earnings. As a result, women in the model choose to take some or all of the leave, and reduce their labor supply in their thirties substantially more than they do in the baseline model. The second effect of the policy is that women's labor supply early in life increases. The reason for this is that eligibility for maternity leave in any given period depends on employment and earnings in the prior period. This incentivizes women to increase their labor supply before they have children.

The third main effect on labor supply is that women allocate less time to market work after their thirties than they do in the baseline model, especially when there are still children in the household. Under the maternity leave policy women accumulate less experience by their mid- to late-thirties, and thus have lower wages later in life. From the household's perspective, it is not optimal for the woman to work the same high hours that she would have with the additional accumulated human capital. As a result, women on average work fewer hours and spend more time in home production and leisure. As children leave the household and as women accumulate human capital towards the end of the lifeycle, they increase their labor supply again.

The policy affects men's labor supply only marginally, as Figure 2.8 shows. Men in their thirties decrease their hours somewhat relative to the baseline model. The reason for this is that the household has more income under the paid leave policy, and men as a result supply less labor than they would in the baseline model.

Finally, the model predicts that the policy increases both college gender gaps. The reason that women decrease their participation in science/business majors under the policy is related to the observation above that women take more time off from the labor force and therefore accumulate less experience. Because returns to experience are high in the science/business occupation, this further reduces women's return to the science/business major. The model also implies that *more* women go to college overall. The reason women's return to college further increases is that women on leave are compensated according to their most recent earnings, and therefore can benefit from the college wage premium even in the periods after childbirth when they are not working.

Part-Time Work Entitlement

In the second experiment, I consider a non-discriminatory part-time work policy, in which employees are entitled to work part-time without any wage penalty, if they choose. Policies aimed to provide this kind of benefit to workers have been enacted in Belgium, France, and the Netherlands among other OECD countries (OECD (2010)). To simulate the effects of the policy, I reduce the wage penalty for working part-time in the model for all occupations to zero, for both men and women.

The series indicated by the green line in Figure 2.7 shows that the policy has one main effect on women's lifecycle labor supply: women choose to supply less labor to the market over most of their lifetime. The results of the simulation imply that when women can choose their desired number of hours worked without incurring any kind of wage penalty, they supply less labor on average. This large reduction in labor supply is observed even though the policy has a strongly positive effect on women's wages in the simulation. The result implies that households highly value the ability to specialize.[16] The only exception to the observed reduction in women's labor supply over the lifecycle is that women in their early twenties work at similar rates as they do in the baseline model. This is in line with intuition. Since almost all women in the model at this age are single or married without children, they do not value flexibility.

The policy decreases the gender gap in choice of major substantially. When there is no wage penalty, more women in the model enter science/business occupations, and as a result more choose the science/business major. The model implies that the share of women choosing a science/business major increases from 34% to 45%, whereas the share of men choosing the major increases marginally from 66% to 67%. Figure 2.8 shows that the part-time policy has almost no effect on men's labor supply. As a result, the effect on men's educational choices is also small. The net effect of the policy is to reduce the gender gap in majors by about a third, from 26 percentage points to 17 percentage points. The reason the model does not imply complete gender convergence in choice of major is that women continue to supply substantially less labor over their lifetime than men. Therefore, the additional financial return to choosing this major is still lower for women.

Subsidized Child Care

In the final experiment, I consider the effect of a child care subsidy, as has been implemented in a number of countries (OECD (2010)). In the baseline model, the household pays for childcare

[16] This is in line with empirical evidence in Goldin and Katz (2012) on the pharmacy occupation, which has almost no wage penalty today for part-time work. Goldin and Katz document high part-time rates among female pharmacists despite the high compensation associated with the occupation.

costs if there is a child under the age of six in the household and both spouses work. In that case, the household pays an hourly childcare cost for the number of hours worked by the spouse with the lower labor supply. In this particular policy experiment, I provide a full subsidy, i.e. I reduce the cost of childcare to zero.[17]

The series represented by the dotted red line in Figure 2.7 shows that subsidized child care has a small effect on women's labor supply patterns. The policy increases the labor supply of college women in the model mostly in their early childbearing years. In the baseline model, the childcare cost is usually an implicit tax on the woman's wage, since she is typically the one to provide the lower labor supply in the household. Among college-educated women in the model, those who respond to the policy are primarily women in science/business occupations with a low wage draw that period. The effective increase in their wage due to the subsidy is enough for some of these women to increase labor supply to avoid a wage penalty or to accumulate experience, which has a high future return. However, because these estimated effects are small, the implied effects on the returns to education are also minor.

Table 2.9 summarizes and compares the effects of the three family-friendly policies on educational choices. As a whole, the policy results suggest that a program that successfully increases the availability of part-time work would be the most effective in increasing women's participation in both science/business occupations and majors.

2.6 Conclusion

To analyze the drivers of educational decisions identified in Chapter 1 in greater detail, I construct and estimate in this chapter a dynamic structural model of lifecycle marriage, labor supply, occupational and educational choices. Using the model, I show how changes in marriage patterns following divorce laws and changes in skill premiums over time affect the educational decisions of different cohorts. I estimate that the insurance value of a degree for women today is equiva-

[17] I set the child care cost to $5.60 per hour, based on data from the Census Bureau on weekly child care expenditures for mothers of children 5 or under who use child care (Laughlin (2013)).

lent to about 31% of their wage premium. I also estimate that the difference in the shares of men and women choosing science/business majors would decrease by about a third if penalties to labor supply reductions were equalized across occupations.

More than 75% of college-educated women today have children over the course of their life.[18] This means that for the vast majority of college-educated women today the question is not "children or career," but rather "children and career, or children and no career." The findings in this paper suggest that many college-educated women select lower-paying occupations and majors to have flexibility to allocate more time to child care and home production, especially when their children are young. For this reason, I use the structural model to test several policies that potentially improve work-family flexibility for women across occupations. I find very large differences in the effects of these policies, with some even further widening both gender gaps in education. The findings call for more research in the future on flexible policies that make it easier for women to participate in science and technical occupations, especially as many developed economies grow increasingly more concerned about the low number of graduates in science and technical fields.

[18] In the 2009-2011 ACS, among women with a 4-year college degree, at age forty 75% have a child currently in the household. This is likely to be a lower bound for the share of college-educated women who ever had a child.

Tables and Figures

Figure 2.1: Weekly Hours Worked in Labor Market, Data and Model

A. Women | B. Men

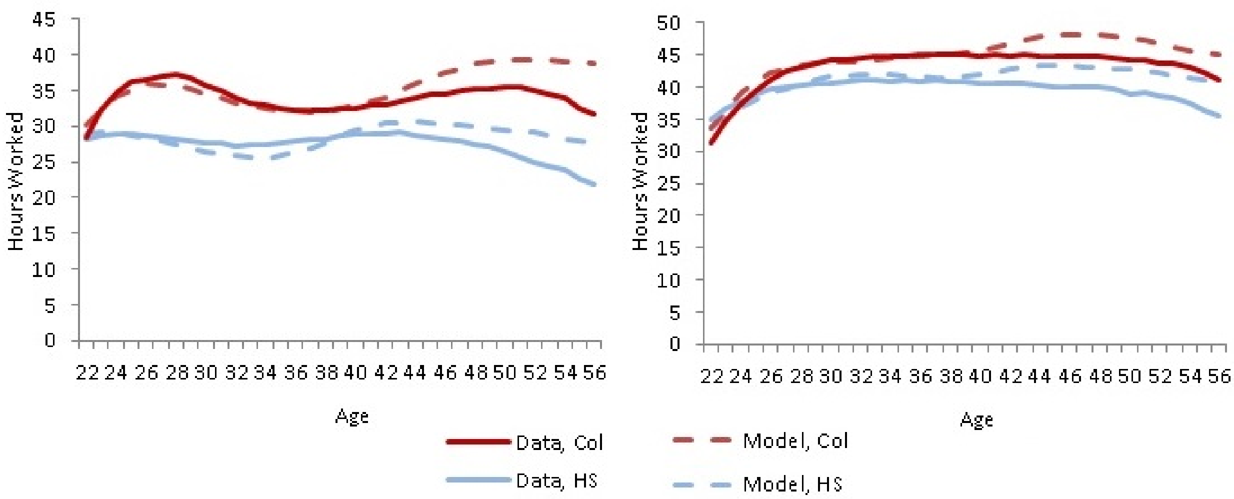


Figure 2.2: Weekly Hours Worked in Home Production, Data and Model

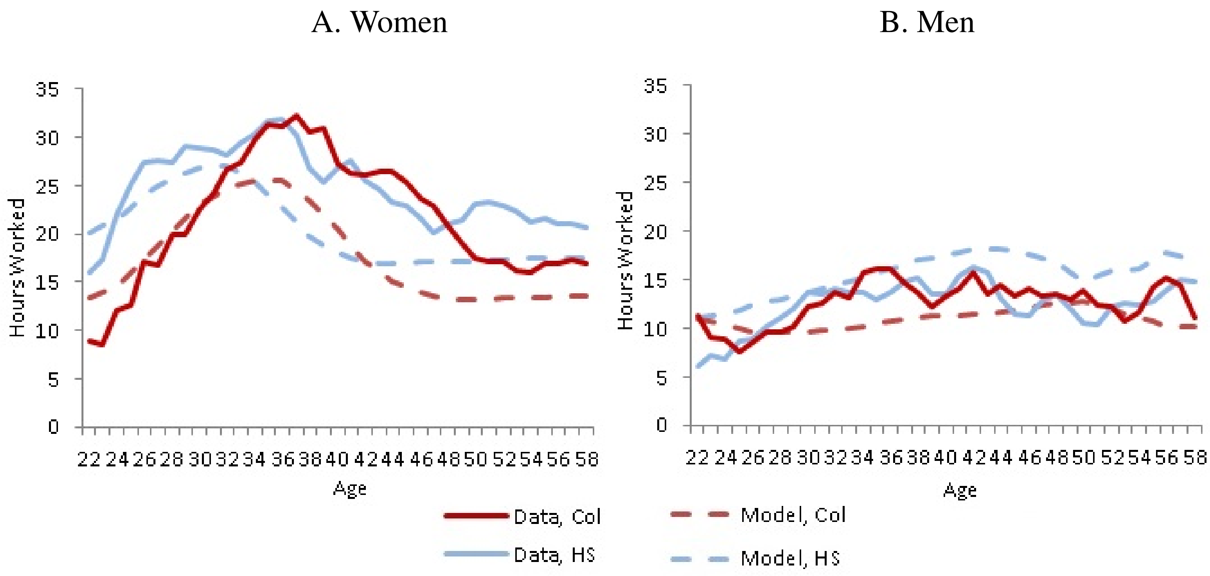


Figure 2.3: Share Married and Divorced (Women), Data and Model

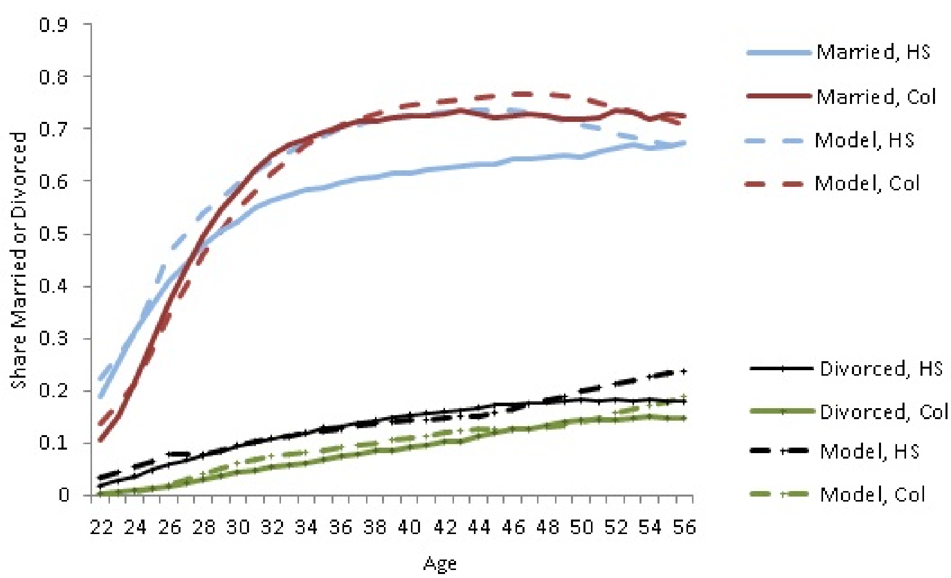


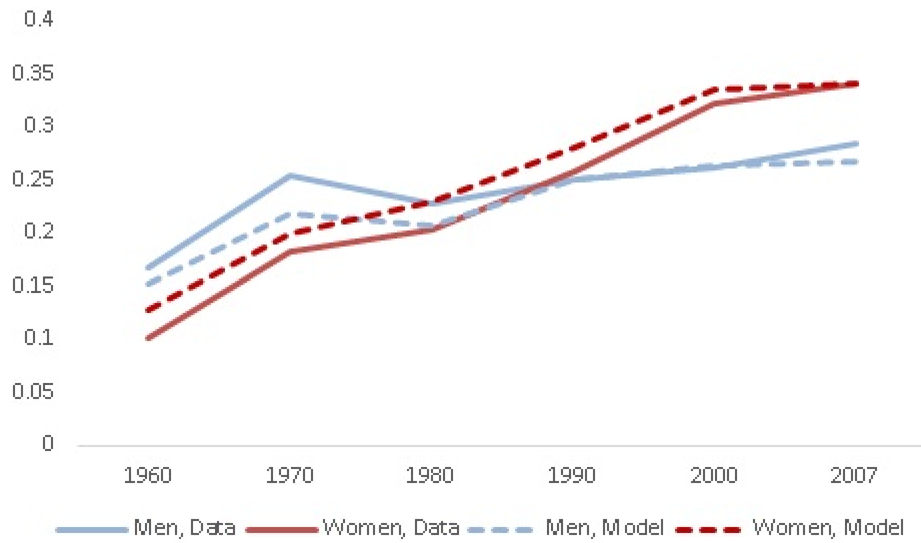
Figure 2.4: Share in $\mathscr{H}$-Type Occupation by Major, Data and Model

A. Women

B. Men

Figure 2.5: Men's and Women's Educational Choices Over Time, Data and Model

A. Share Graduating from College, Data and Model



B. Share Graduating Choosing a Type *H* Major, Data and Model

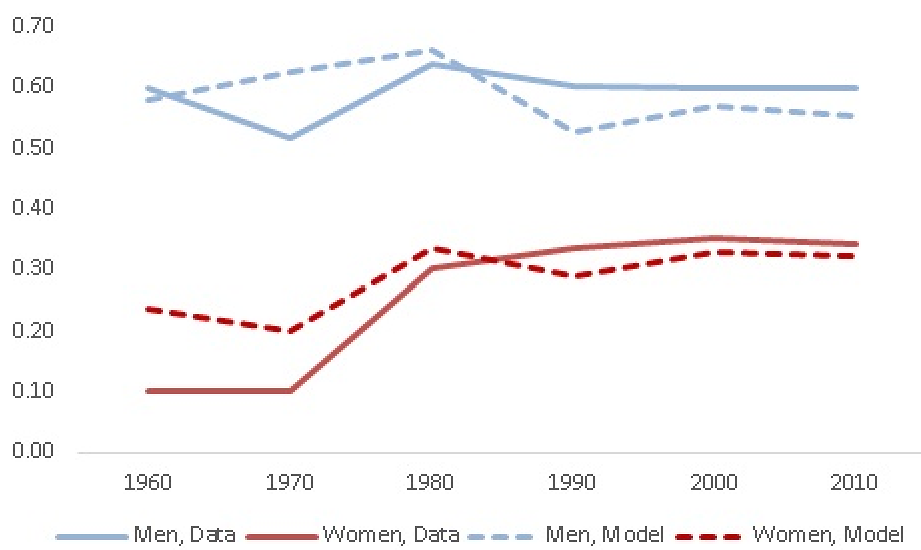


Figure 2.6: Share Married and College Women's Labor Supply Under Alternative Divorce Law Regimes, 2010 Graduating Cohort

Figure 2.7: Simulated Labor Supply Under Family-Friendly Policies, Women

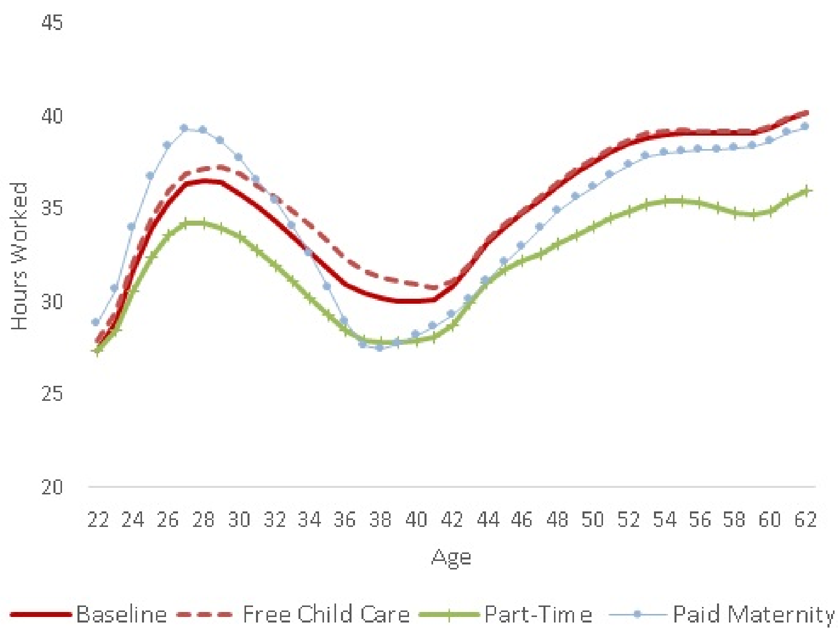


Figure 2.8: Simulated Labor Supply Under Family-Friendly Policies, Men

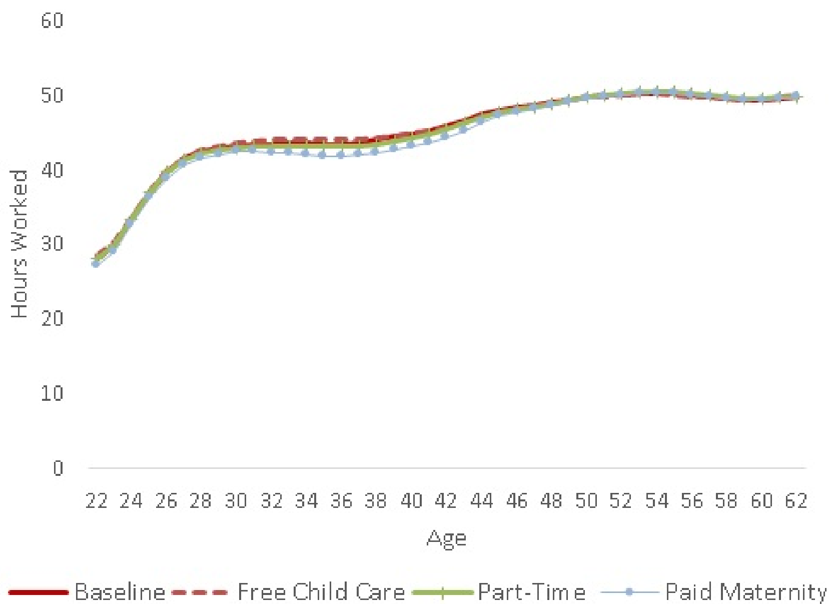


Table 2.1: Wage Process Parameters for Baseline Cohort

	High School Occ. $\mathscr{L}$	Major L Occ. $\mathscr{L}$	Major L Occ. $\mathscr{H}$	Major H Occ. $\mathscr{L}$	Major H Occ. $\mathscr{H}$
Entry-Level Wage Gap	0.21	0.11	0.14	0.12	0.19
	(0.04)	(0.03)	(0.03)	(0.04)	(0.05)
PT/Time-Off Penalty, Men	0.08	0.09	0.16	0.09	0.16
	(0.006)	(0.017)	(0.033)	(0.023)	(0.020)
PT/Time-Off Penalty, Women	0.09	0.09	0.17	0.10	0.19
	(0.006)	(0.11)	(0.026)	(0.06)	(0.022)
Experience, Men	0.05	0.08	0.09	0.09	0.11
	(0.004)	(0.005)	(0.005)	(0.005)	(0.007)
Experience, Women	0.04	0.08	0.09	0.07	0.10
	(0.004)	(0.003)	(0.004)	(0.004)	(0.006)
Experience^2, Men	-0.002	-0.003	-0.003	-0.003	-0.003
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Experience^2, Women	-0.002	-0.003	-0.003	-0.003	-0.003
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Standard errors in parentheses. Source: NLSY79.

Table 2.2: Calibrated Parameters

Description:	Parameter:	Value
Risk aversion parameter	σ	2.50
Cobb-Douglas parameter on consumption/leisure	a	0.40
Discount factor	β	0.92

Table 2.3: Estimates of Home Good Technology Parameters

Description:	Parameter:	Estimate
Home labor productivity, by education and child status	α_1	
No Children:		0.178
		(0.28)
High School, Children <6:		0.713
		(0.29)
College, Children <6:		0.967
		(0.15)
High School, Children 6+ Only:		0.591
		(0.33)
College, Children 6+ Only:		0.821
		(0.36)
Productivity of market good	α_2	0.344
		(0.27)
Children's contribution to home good	α_3	2.511
		(0.89)

Standard errors in parantheses.

Table 2.4: Estimates of Marriage Market Parameters

Description:	Parameter:	Estimate
Mean of initial match draw	μ_θ	-1.45
		(1.20)
Variance of initial match draw	σ_θ	1.53
		(0.80)
Mean of match quality shocks	μ_z	0.24
		(0.34)
Variance of match quality shocks	σ_z	1.36
		(0.49)
Probability of drawing a partner with the same education	p_m	0.81
		(0.21)
Cost of divorce before reform	K_0	10.84
		(1.28)
Cost of divorce after reform	K_1	1.08
		(0.17)
Re-marriage penalty for individuals with children	P_{RM}	2.91
		(1.91)

Standard errors in parantheses.

Table 2.5: Estimates of Utility Cost of College and Occupational Wage Draw Parameters

Description:	Parameter:	Estimate
Mean of utility cost, Type L major	μ_L	12.33
		(2.62)
Variance of utility cost, Type L major	σ_L	2.39
		(2.01)
Mean of utility cost, Type *H* major	μ_H	22.43
		(2.77)
Variance of utility cost, Type *H* major	σ_H	6.67
		(2.48)
Probability of wage draw from second occupation	μ_L	0.11
		(0.15)

Standard errors in parantheses.

Table 2.6: Summary Statistics, Data and Model Simulation

Moment:	Data	Model
Share Married, Ages 22 to 60	0.61	0.65
Share Divorced, Ages 22 to 60	0.12	0.11
Weekly Hours Worked in the Labor Market:		
Men, HS	40.74	41.19
Men, College	42.99	43.97
Women, HS	28.43	29.04
Women, College	34.57	35.16
Share in $\mathscr{H}$-type occupations:		
Men, Major *L*	0.42	0.43
Men, Major *H*	0.68	0.66
Women, Major *L*	0.24	0.26
Women, Major *H*	0.51	0.51
Weekly Hours Spent in Home Production & Child Care, Married Men:		
No Children, HS	8.54	7.30
No Children, $\mathscr{L}$	10.74	7.93
No Children, $\mathscr{H}$	6.71	5.40
Children <6, HS	24.03	19.71
Children <6, $\mathscr{L}$	24.89	24.05
Children <6, $\mathscr{H}$	25.01	20.74
Weekly Hours Spent in Home Production & Child Care, Married Women:		
No Children, HS	25.11	20.98
No Children, $\mathscr{L}$	22.17	17.18
No Children, $\mathscr{H}$	19.47	14.51
Children <6, HS	44.36	47.31
Children <6, $\mathscr{L}$	48.14	49.30
Children <6, $\mathscr{H}$	45.03	36.63

Table 2.7: College Returns and Decisions for 1980 Graduating Cohort, Model and Data

	Women	Men
Discounted Lifetime Utility Returns, Model:		
Return to Major *L*, vs. HS	4.06	2.31
Additional return to Major *H*, vs. Major *L*	0.62	3.06
College Choices, Model:		
Share graduating	0.23	0.21
Share choosing Major *H*	0.33	0.66
College Choices, Data:		
Share graduating	0.20	0.22
Share choosing Major *H*	0.30	0.64

Table 2.8: Effect of Tuition Policy on Share Choosing Major *H*

	Baseline	Policy
Women	0.33	0.41
Men	0.66	0.67

Table 2.9: Main Effects of Family-Friendly Policies on Educational Choices

Policy	Main Effects
Paid Maternity Leave	Both college gender gaps increase: · Women major less frequently in science/business · Women increase their college attendance further
Part-Time Entitlement	Gender gap in majors narrows: · Large increase in share of women choosing science/business majors · Small increase in women's college attendance
Subsidized Child Care	Gender gap in majors narrows marginally · Marginal increase in share of women choosing science/business majors · No effect on attendance

Chapter 3

Cohort Size and the Marriage Market: Explaining Nearly a Century of Changes in U.S. Marriage Rates

3.1 Introduction

What causes variation in marriage rates over time?[1] For both economists and policy-makers, this is a question of significant interest, as a large body of evidence suggests that marriage rates have important implications for other economic variables. Such variables include fertility rates, children's welfare, children's education, labor force participation, hours of work, income inequality, the fraction of individuals on government aid, population growth, and workers' productivity.[2] In spite of this, according to the existing literature, no prevailing explanation can account for the variation in marriage formation over time and across geographies. Many existing theories about changes in the marriage rate, which will be reviewed in the next section, have either been empirically rejected, or have explanatory power that is limited to specific periods or specific groups of individuals.

[1]This chapter is joint work with Maurizio Mazzocco.

[2]See for example Killingsworth and Heckman (1986), Moffitt (2000), Angrist and Lavy (1996), Gruber (2004), McLanahan and Percheski (2008).

The main contribution of this paper is to provide an explanation for changes in U.S. marriage rates that holds empirically over nearly a century. We show that one variable, changes in cohort size, explains the majority of the variation in U.S. marriage rates since the early twentieth century, both over time and across states. The paper consists of 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. We provide reduced-form evidence in two steps. First, using both time-series variation and cross-sectional state-level variation in cohort size and marriage rates we find that there is a strong and negative relationship between marriage rates and cohort size. On average, a 10-percent increase in cohort size is associated with a 2 to 3 percent decrease in the share ever married by 30 and around a 0.5 to 1 percent decrease in the share ever married by 40. These are sizeable effects that account for around 50 to 70 percent of the variation in marriage rates since the early twentieth century. Our results indicate that changes in cohort size account well for medium- and long-term changes in marriage rates, but cannot explain the short-run year-on-year variation. They also suggest that part of the effect of cohort size on marriage rates operates through an increase in the age at first marriage.

In the second 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 based on Bailey (2010), we employ the interaction between the 1957 introduction of *Enovid*, later known as the birth-control pill, and cross-state variation in anti-obscenity laws, which limited the use of contraception in some states until the mid-1960s, to generate exogenous variation in the number of births and therefore cohort size. Our results indicate that, in states that limited contraceptives, cohort size increased relative to states that did not have such limits, and that this change generated a decline in marriage rates. The exogenous variation, therefore, gives results that are consistent with the time-series and cross-state variation.

In the second part of the paper, we propose a model that can potentially generate the relationship between changes in cohort size and changes in marriage rates observed in the data. 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 justified by the fact that men's fertile lifespan is longer than women's, and the common insight that one benefit of marriage is that it is an effective arrangement for raising children. 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 intuition underlying this result is straightforward. In the model, the marriage market is populated by young women and by young and old men. A positive shock to cohort size with some degree of persistence will therefore have two effects. First, old men become scarce relative to young women. As a consequence, young women will be less likely to meet and marry them. This direct effect will reduce the marriage rate. Second, young men will become more selective. The reason for this is that they will have a higher probability of meeting a young woman in case they choose to wait until they are old. As a consequence, if they meet a potential spouse whose quality is not sufficiently high, they will decide to wait. This indirect effect will further reduce the marriage rate. The model can therefore qualitatively explain the negative relationship between cohort size and marriage rates. We then derive an implication that can be used to test the model. We show that in our 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 we do not reject it. 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 a substantial portion of 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.

Our findings have an important policy implication. In recent years, politicians and policy makers have started to discuss and implement policies that attempt to improve the well-being of low income families by increasing the fraction of married individuals. For instance, during the Bush Administration, proposals pending approval planned to allocate up to 1.5 billion dollars to implement and evaluate policies aimed at promoting marriage.[3] Our results suggest that these policies may prove to be largely ineffective since a significant part of the changes in marriage rates is generated by forces that are mostly outside the control of policy makers. This is likely to be particularly true in the short and medium term which is the time horizon available to politicians and policy makers.

We conclude this section with one remark. The variable cohort size is not exogenous. It is affected by variables like new fertility technologies, technological progress, child care availability, supply of housing, and changes in labor supply decisions of women. This fact does not diminish the importance of our results for social scientists and policy makers. They indicate that to understand the dynamics of the marriage market one has to study the evolution of cohort size and of the variables that have an effect on it.

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 estimate the search model. Section 7 concludes.

3.2 Existing Explanations

In this section, we describe some of the existing explanations for the variation in U.S. marriage rates for which some empirical evidence is provided.

One set of explanations for historical changes in marriage rates focuses on changes in income. Cherlin (1981) among others notes the correspondence between rising incomes after World War II and the associated marriage boom during this period. A related common insight is that low income during the Great Depression was the main factor behind the reduction in marriage rates during

[3] Seefeld and Smock (2004) provide a nice discussion of the recent interest of policy makers in marriage as a policy tool.

this period. A positive relationship between income and marriage rates, however, has not been successfully tested over different periods. Hill (2011) rejects the hypothesis that there is a positive correlation between income and marriage rates after 1960. Wolfers (2010) looks at the relationship between marriage rates and recessions for the past 150 years and rejects any pattern between marriage and periods of economic decline. These results suggest that income may generate part of the variation in marriage rates. But they also suggest that variation in income cannot be the general explanation that underlies the changes in marriage rates observed in the past century. In addition, there is a potential reverse causality problem with this theory that is not addressed: men have higher labor hours and earn more following marriage.[4] It is therefore difficult to determine whether an increase in income causes marriage rates to rise or whether an increase in the marriage rate generates higher income levels.

An alternative set of explanations focuses on technological innovations. Akerlof, Yellen, and Katz (1996) consider the adoption of new fertility technologies such as the pill and abortion in the sixties and seventies. The authors argue that these technologies increased the number of out-of-wedlock births and reduced the number of marriages. A potential direct effect on marriages could be that the adoption of the new technologies mechanically reduced conceptions and therefore 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. They argue that a decline in the cost of abortion and the increased availability of contraception decrease the incentives of women to obtain a promise of marriage if premarital sexual activity results in pregnancy with a consequent rise in out-of-wedlock births and decline in marriage rates. This explanation may provide important insight into changes during and around the 1970s, but it would not be able to explain intervals of increasing marriage rates after the 1970s, and it cannot address historical variation prior to this period.

Greenwood and co-authors also focus on technological innovations and suggest that the decline in the price of household appliances explains patterns for several household outcomes, including

[4] A number of studies have documented this relationship. See, for example, Mazzocco, Ruiz, and Yamaguchi (2007).

marriage. Greenwood and Guner (2009) argue that labor-saving technological progress in the household sector made it easier for singles to maintain their own home, increasing the value of being single and reducing the marriage rate. Their theory addresses in particular the period of rapidly falling marriage rates starting around the mid-sixties. However, the theory would have trouble explaining much of the remaining historical variation in marriage rates. In as far as technological progress has constantly improved household appliances in the past 100 years, we might predict that marriage rates either decline systematically throughout this period or decline and then remain constant. However, this contradicts what we observe in the data.

A different well-known set of theories considers men's marriageability and policies such as welfare aid, both of which may affect women's desire to marry. These theories have been widely empirically tested, with mixed results. Ellwood and Crane (1990) review papers that have tested the hypothesis proposed by Wilson (1987) that limited labor market opportunities reduce the number of marriageable men, and thus the marriage rate. While some papers provide evidence in support of this theory, others reject this hypothesis (e.g., Plotnick (1990) and Lerman (1989)). The degree to which men's employment opportunities affect marriage rates is largely still an open question. A related issue affecting especially black men's marriageability in the last three decades is the rise in incarceration. Charles and Luoh (2010) study the relationship between incarceration and marriage rates and find that higher incarceration rates decrease the fraction of married individuals both in black and white populations. Quantitatively, however, such an explanation is primarily relevant for black populations and only for the period after 1980, when drug-related policies significantly increased incarceration rates. Ellwood and Crane (1990) have also evaluated the link between welfare aid and marriage. They conclude that there is very little empirical support for the proposition that welfare benefits played a major role in marriage trends for black or white women. Finally, rising income and employment opportunities for women may also affect their desire to marry. However, we do not know of papers that formally test this hypothesis over time or across geographies. One reason may be that potential reverse causality complicates such an empirical analysis: women who face poorer marriage prospects may both invest more in human capital and work more.

Finally, there are two explanations that have commonality with the one we propose. The first explanation is Easterlin's hypothesis. Easterlin (1987) argues that the relative size of a cohort can explain many of the variables that determine the economic and social outcomes of a birth cohort: earnings and unemployment rates, college enrollment rates, divorce, fertility, crime, suicide rates, and marriage. Easterlin's explanation is that when income is above the aspiration level for a given cohort, the individuals in that 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 outcomes. 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 (Pampel and Peters (1995)).

The second explanation that is related to ours is the sex-ratio hypothesis. According to this hypothesis, when the marriage market is characterized by a high sex ratio, measured as the number of available women divided by the number of available men, the marriage rate should decline. Becker (1973) is one of the first attempts to theoretically characterize a possible relationship between the sex-ratio and marriage rates using a matching model. In a well-known series of case studies, Guttentag and Secord (1983) argue that declines in the number of women relative to men lead to higher marriage rates, among other social changes. Some papers have tried to test this relationship. For instance, Schoen (1983) considers changes in sex ratios due to two factors: changes in the rate of population growth and differences between women and men in preferences for the age of their spouse. He then evaluates the effect of these changes on marriage rates in the U.S. and finds little effect. Angrist (2002) looks at variation in immigration rates from different European countries to the U.S. at the beginning of the twentieth century. He exploits the fact that the majority of migrants were men and that marriages were often formed between individuals belonging to the same ethnicity, and finds that ethnicities that experienced lower sex ratios display higher marriage rates. Finally, Abramitzky, Delavande, and Vasconcelos (2011) consider variation in the sex-ratio due to World War I casualties in France and find that a higher sex ratio is associated with a lower marriage rate for women and with a larger marriage rate for men. In the search model we develop, variations in marriage rates are determined by changes in the sex ratio, which are endogeneosly generated by changes in cohort size. The mechanism we propose is therefore partially related to papers that use changes in the sex ratio to explain changes in marriage rates.[5]

3.3 Data

In this section we describe the data used in the analysis. Throughout the paper we rely on the following datasets: National Vital Statistics (1909-1980), Census total population counts (1910-1980), Survey of Epidemiology and End Results (SEER) population estimates (1969-2000), IPUMS CPS (1962-2011) and IPUMS USA (1940-2000). In the Data Appendix B we provide a detailed description of how the datasets are used to construct the main variables employed in the empirical analysis. In this section, we give a brief summary of that description.

In the empirical analysis we employ two main variables: cohort size and the share ever married by a given age for a given cohort. We construct two different measures of cohort size: cohort size at birth and cohort size at marriageable age. The first variable is used with longitudinal variation, whereas the second one is employed with cross-state variation. In this paper we are interested in the evolution of the variable cohort size at the ages in which individuals choose whether and whom to marry. With longitudinal variation, however, we use cohort size at birth as the main independent variable for two reasons. First, as shown in Figure 3.1, when cohort size is computed for the U.S. population there is little difference between cohort size at birth and cohort size at marriageable age, since migration from and to the U.S. was limited during the time period we consider. Second, we can construct the variable cohort size at birth for cohorts born in 1909 and after. The variable cohort size at marriageable age can only be constructed for cohorts born in 1940 and after. By

[5]Several papers have studied the relationship between the sex ratio and other economic variables. One example is the paper by Neal (2004) where the author studies the relationship between sex ratios and rates of single motherhood. A second example is the paper by Grossbard-Shechtman (1984) where the author analyzes the link between the sex ratio and labor force participation of women. Another example is the paper by Bitler and Schmidt (2011) which shows that there is a causal relationship between the sex ratio and birth rates using differences across states and over time in mobilization rates during the Vietnam war.

using cohort size at birth we can therefore consider a larger number of cohorts without significant effect on the analysis. When we use cross-state variation, we have to use the variable cohort size at marriageable age because of large migration flows across states during the period of investigation.

The variable share ever married at age 30, 35, and 40 is constructed using either the decennial Censuses or the SEER population estimates. Appendix B describes the exact procedure used to construct this variable.

3.4 Empirical Results

This section is divided into five parts. We first describe the two measures generally used to study the evolution of marriage rates and compare them to our alternative measure, which we believe is better suited to the examination of marriage choices over time. We then provide empirical evidence on the relationship between cohort size and marriage rates using longitudinal variation. In the third subsection, we describe findings obtained using cross-state variation. We then discuss endogeneity issues that may affect the longitudinal and cross-state variation. Finally, we provide evidence that changes in cohort size generate changes in marriage rates by using variation in early access to oral contraceptives across states as a plausible source of exogenous variation in cohort size.

3.4.1 Different Measures of Changes in the Marriage Rate

When analyzing changes in marriage rates over time, existing studies typically employ one of the following two variables: the number of new marriages per population; and the share of individuals currently or ever married within some age range, e.g. between the ages of 18 and 30. We use a different measure, the share of individuals in a given cohort ever married by a given age. In this section we compare the behavior of these three measures. As a first contribution, we show that, when used to study the evolution of the marriage market over time, the commonly used measures are problematic because they confound different effects and can potentially lead to incorrect inference about overall marriage behaviors.

Figure 3.2 graphs the three measures described above: the number of marriages per 1,000 individuals; the share ever married in a cross-section of women between the ages of 18 to 30; and the share of women in a given cohort ever married by 30. We refer to these measures as the population-based measure, the cross-sectional measure, and the cohort-based measure, respectively. The first two measures are based on the calendar-year, whereas our measure is based on the cohort's year of birth. To make them comparable, we add 25 years to the cohort's birth year and plot our measure together with the other two. It is immediately evident from Figure 3.2 that the three measures give very different pictures of how marriage rates in the U.S. have evolved over time. It is worthwhile to discuss why they behave differently and the main problems with each measure.

The first measure, the number of marriages per 1,000 individuals, is based on the number of new marriages that take place in a given year. It is available from the U.S. Vital Statistics. This measure can be useful to detect years that experience an unusual growth or decline in marriage rates. For example, it captures well the rapid rise in the number of marriages in the immediate post-war years. However, this measure is not as well suited to study the evolution of the share ever married for two main reasons. First, using the population-based measure it is impossible to distinguish between first, second, or later marriages. This distinction is important if the researcher is interested in evaluating the fraction of people who choose to marry, since the second and later marriages should not be included in a measure designed to describe the evolution of the share ever married. The second and more serious problem for evaluating marriage trends is that this measure conflates changes in the numerator, new marriages, with changes in the denominator, population. As a result, if the population undergoes any substantive growth or decline due, for instance, to changes in fertility or migration patterns, one may draw the wrong inference. For example, starting from 1946 this variable drops steeply for about fifteen years. One might infer that marriage rates were falling steadily in the forties and fifties, but both the cross-sectional measure and the cohort-based measure show that marriage rates were flat or rising during this time. Part of the explanation for the large decline in the population-based measure is that during those years the U.S. experienced a sharp increase in population with the baby boom. Similarly, during the sixties and the first half of the seventies, the population-based variable displays rapid growth which can been interpreted as a big increase in marriage rates. The cross-sectional and cohort-based measure, however, show that this interpretation is misleading. During this period, marriage rates experienced a slight decline. A potential explanation for the rise in the population-based measure is that the U.S. population declined during the baby bust that characterized the U.S. in the sixties.[6]

The second measure illustrated in Figure 3.2 is the share ever married in a cross-section of women between the ages of 18 to 30, which was constructed using the CPS and Census. It follows closely the cohort-based measure for most of the period. It is only during the second half of the eighties and the nineties that the two variables diverge. The cross-sectional measure would suggest a sustained drop in marriage rates during this period, whereas the cohort-based measure documents a mild increase in the share ever married. The reason for the divergence is that the cross-sectional measure is strongly affected by changes in the age at first marriage: when examining the share of people ever married within an age range, one cannot determine whether individuals are simply delaying marriage or whether they choose not to marry. Starting from the second half of the eighties, the age at first marriage experienced a significant increase, which explains why the cross-sectional measure declines during this period whereas the cohort-based measure increased.

It is important to note that the cohort-based measure may also be affected by changes in the age at first marriage if the age cut-off is too low. It is therefore important that one chooses properly the age cut-off: if one believes that an age cut-off of 30 is too low, an age cut-off of 35 or 40 should be used. In the next subsections, we will show that the trends the cohort-based measure captures look very similar at different age cut-offs. In Figure 3.2, we choose an age cut-off of 30 mostly for expositional purposes, so that we can include as many recent cohorts as possible.

Because of the advantages the cohort-based measure has over the other two variables, throughout the rest of the paper we will only use that measure to study the evolution of marriage rates.

[6]A second potential explanation for the increase in the population-based measure in the seventies is that the large baby boom cohorts were coming of age. Even though these individuals were marrying at lower rates, their large overall number will generate and increase in the number of registered marriages, the numerator.

3.4.2 Change in Marriage Rates Over Time

In this subsection we provide evidence on the relationship between changes in cohort size and changes in marriage rates using longitudinal variation. The evidence is presented in two steps. We first provide evidence on the general nature of this relationship. We then try to understand whether cohort size can explain the short-run, medium-run, or long-run changes in marriage rates.

In Figure 3.3 we plot cohort size and the share never married by age 30, separately for women and men, for all cohorts born between 1914 and 1981. The first panel describes these variables for the white population whereas the second panel plots them for the black population. We plot the share *never* married because visually it is easier to detect a positive correlation between the two variables. Figure 3.3 contains one main finding. To describe it, we initially focus on cohorts born before 1960. For those cohorts, there is a strong positive correlation between cohort size and the share never married. The decline in size for cohorts born in the 1920s and 1930s is associated with a similar drop in the share never married. This decline corresponds to the well-documented "marriage boom" that starts in the mid-1940s and lasts through the early 1960s, the period in which the cohorts born in the twenties and thirties were active in the marriage market. The sharp increase in the size of cohorts born between 1946 and 1959, which correspond to the post-war baby boom generations, is associated with a share never married that nearly tripled during this period. Births and the share never married for the black population follow similar patterns.

It is left to explain why we lose the positive correlation between our two main variables for cohorts born in the 1960s and 1970s. Note that those cohorts were active in the marriage market starting from the 1980s, which is the period in which cohabitation began to become a popular form of household formation and potentially a close substitute for marriage. To understand whether cohabitation can resolve the inconsistency between the early and later cohorts, in Figure 3.4 we plot the variables reported in the previous figure, with the exception that now cohabiting individuals are treated as married individuals instead of being treated as never-married individuals. Remarkably, once cohabiting households are accounted for, the relationship between cohort size and household formation resembles again that of the earlier cohorts. Falling cohort sizes in the 1960s and early

1970s correspond to a decline in the share never married and not currently cohabiting by 30. Increasing cohort sizes in the second part of the 1970s are associated with a rise in the share never married and not cohabiting.

Figure 3.4 contains a second noteworthy finding. The strong positive correlation between cohort size and share never married characterizes both the white and the black populations. We emphasize this similarity between the white and black marriage markets because it challenges the common perception that the two markets are governed by different rules and exhibit different marriage behaviors. Figure 3.4 suggests that the two marriage markets are similar in at least one respect: they respond to changes in cohort size in a similar way.

In Figures 3.3 and 3.4, we use an age cutoff of 30. The results may therefore be affected by changes in age at first marriage. To address this concern, in Figure 3.5 we plot cohort size and shares never married and not cohabiting by age 40. Using this new cutoff age, we find patterns that are similar to the ones observed in the first two figures: there is a positive and strong correlation between cohort size and share never married and not cohabiting.

To show these patterns more formally, in Table 3.1 we record the average response of marriage rates to changes in cohort size for age cutoffs of 30, 35, and 40. For ease of exposition, for the rest of the paper we will consider the effect of cohort size on the share ever married instead of the share never married. Specifically, in the table each coefficient is the outcome of a separate regression of the log share ever married or currently cohabiting on log cohort size. There are three results that are worth discussing. First, elasticities recorded in Table 3.1 are highest at 30, and gradually decrease with age, for both sexes and both races. This finding suggests that changes in cohort size are associated with two effects: (i) a change in the eventual share ever married or cohabiting; (ii) a change in the age at first marriage, where an increase in cohort size is associated with a higher age at first marriage. This finding also indicates that the coefficient that is better able to isolate the relationship between variation in cohort size and variation in marriage rates from changes in age at first marriage is the one that uses 40 as a cutoff age. The second result is that the effect of cohort size is quantitatively large and is highest for black women. An increase of 10% in cohort size reduces the share ever married or cohabiting by 40 by a percentage that ranges from 0.66% for white women to 4.53% for black women. In percentage points, this amounts to a decline that is between 0.6 and 3.8 points in the share of individuals ever married or cohabiting by 40, a large effect. The last finding we wish to emphasize is that the cohort size variable explains a large fraction of the variation observed in marriage rates. For instance, when we use 40 as the cutoff age, the R-squared is between 0.58 and 0.85.[7]

In the rest of the paper, we will focus on the white population only for one main reason. White women and men have similar cohort size at the time of marriage, whereas black men have a significantly lower cohort size than black women because of higher mortality and incarceration rates. As a consequence, the investigation of the marriage market for blacks requires a different type of analysis which we undertake in a separate paper.

In the remaining part of the section we will try to understand whether cohort size can explain the short-run, medium-run, or long run changes in marriage rates. Observe that the regressions in log-levels documented in Table 3.1 capture the short-run, the medium-run, as well as the long-run effect of changes in cohort size on changes in marriage rates. To see this, note that the regressions in levels measure the correlation between changes in our two main variables independently of when the changes took place. The same weight is assigned to the change in these two variables between the cohort born in 1935 and the cohort born in 1958 as to the change between the cohort born in 1935 and the cohort born in 1936. To measure the short-run, medium-run, and long-run effects of changes in cohort size on changes in marriage rates, we regress n-year differences in marriage rates on n-year differences in cohort size where n is set equal to 1, 2, 3, 4, 5, 7, and 10. To capture the effect of adjacent cohorts, for $n > 1$ we use differences in cumulative cohort size as our independent variable, where cumulative size for cohort born in period t for the n-year difference is constructed by adding up cohort size from $t-n+1$ to t.[8] We interpret the coefficient estimates on the 1-year

[7] Note that we are working with non-stationary time series, and must therefore verify that the series are cointegrated to eliminate worries about spurious regression. A Johansen cointegration test rejects the null hypothesis that the series are not cointegrated at the one-percent level. Therefore, our OLS results are consistent and estimate a meaningful (non-spurious) relationship.

[8] We have also estimated the effect of cohort size using simple n-year differences. The estimates display similar patterns, but the coefficients are generally smaller and are less precisely estimated.

and 2-year differences as the short-run effect of cohort size on marriage rates, the coefficients on the 3-year, 4-year, and 5-year differences as the medium-term effect, and the coefficients on the 7-year and 10-year differences as the long-term effect. This choice is somewhat arbitrary, but it helps us focus the discussion.

The results are presented in Table 3.2. The first two columns report the short-run effect. Clearly, in the short run cohort size has at best a weak effect on marriage rates. The estimates for the 1-year differences indicate that a 1-year change in cohort size is not sufficient to trigger a change in marriage rates. With the exception of the coefficient for women by age 30 and men by age 40, all the coefficients for the 1-year difference are statistically equal to zero. The estimated coefficient for men by age 40 is the only one in all of the estimations we have performed that is positive and statistically significant. This result is generated by the two spikes in births that occurred in 1942 just at the start of World War II and in 1946-1947 after World War II ended. If one drops the observations that characterize the period around World War II, the coefficient becomes zero. The effects are slightly larger when we employ 2-year differences. Now the coefficient for women by age 35 is also negative and statistically significant. But even for 2-year differences, the effect of cohort size is very weak. The effect of cohort size on marriage rates is much stronger when we study the medium-term effect. With 3-year differences all the coefficients become large in size, negative, and statistically significant. The only exceptions are the coefficients by age 40 which are negative but statistically not significant. When we increase the differences to four and five years, the coefficients become larger in size and are now all statistically significant. The long-term effect of cohort size is even stronger. The coefficient estimates for the 7-year and 10-year differences suggest that, for an age cutoff of 40, an increase in cohort size of 10% generates a drop in marriage rates of 0.6-0.8%. This is a significant decline since it implies that a standard deviation increase in cohort size decreases the share ever married by 40 by a third to a half of a standard deviation. These findings indicate that cohort size can explain the medium and long term variation in marriage rates, but not the short term variation. Changes have to cumulate for longer than one or two years to generate significant fluctuations in marriage rates.

To summarize, our results indicate that in the time-series data there is a strong and negative relationship between marriage rates and cohort size. The results also indicate that changes in cohort size account for a large fraction of the medium-run and long-run time series variation in marriage rates. In the following subsections, we further explore the empirical link between cohort size and marriage rates. Because cohabitation has become a close substitute for marriage since the early eighties, in the rest of the paper we will continue to use the same adjusted measure of household formation: the share ever married or cohabiting by a given age. Unless we specifically note otherwise, when we use the shorthand "ever married" we refer to those ever married or currently cohabiting.

3.4.3 Change in Marriage Rates Across States

In this subsection we provide additional evidence on the relationship between cohort size and household formation rates by using variation across states. The idea is that if changes in cohort size generate variation in marriage rates, we should observe such an effect not just across time but also across geography. We should observe that states with larger increases in cohort sizes experience larger drops in marriage rates.

To use cross-state variation, there is one issue we need to address that is not present in the longitudinal analysis. Changes in the size of a cohort at the time individuals are of marriageable age are endogenous because they are partially driven by migration decisions. These decisions are generally related to differences across states in economic and social conditions which also affect marriage rates. To exacerbate the problem, migration can be sex-biased. It can therefore skew sex ratios and affect marriage rates directly. To address this potential source of endogeneity, we use total births in a given year and state, which are arguably unaffected by the endogeneity issues discussed above, as an instrument for the size of the marriage market.

To perform the analysis at the state level, we rely on the decennial Census over the entire period of interest since it is the only dataset with sufficiently large sample sizes for all states. Appendix B provides details about how we construct the three main variables needed for the analysis: cohort size at birth, cohort size at marriageable age, and share ever married for each state and cohort. Using the Census data we can only use the decennial cohorts born between 1910 and 1970, since the share ever married can only be computed for them. We therefore perform the empirical analysis by first constructing ten-year differences for the log of share ever married, the log of cohort size at birth, and the log of the size of the marriage market for each cohort and state. We then use these variables in our regressions. Because of the small number of observations, we pool all cross-sections and regress the differences in log share ever married on the differences in log cohort size where log cohort size is instrumented using log cohort size at birth. We add year fixed effects to the regression to control for general time trends.

Table 3.3 presents the results of the cross-state regressions separately by gender. The first column presents the estimates obtained using a standard OLS regression of 10-years differences in log share ever married on 10-year differences on the size of the marriage market. Similarly to the findings obtained using longitudinal variation, the estimated coefficient is negative and statistically significant for the two age cutoffs we consider, for both men and women. As argued above, the OLS estimates may be biased because of migration decisions. In particular, if migration is partially driven by the desire to find a spouse, the estimates will be downward biased. A similar result applies if migration is partially driven by job-related reasons. In this case, on average one would expect individuals to move to states with higher earning opportunities. If individuals with higher income are also more likely to marry, this would generate a positive correlation between cohort size growth and marriage rates, biasing the results away from the strongly negative relationship we predict.

We can now describe the estimates obtained when we control for endogeneity by instrumenting the size of the marriage market using cohort size at birth. At the bottom of the table we present the first stage results. Two findings are worth discussing. First, the first stage suggests that cohort size at birth explains a large fraction of the marriage market size. Second, the coefficient on log cohort size at birth is estimated to be 0.44, which is well below 1. This indicates that cross-state migration plays an important role in changes in cohort size at marriage age. In column two we present our IV estimates. They are all negative, statistically significant and, as expected, larger in size than the OLS estimates. For the share ever married by 30, the coefficient is estimated to be -0.104 for women and -0.123 for men. When we use the age cutoff of 40, the coefficient drops to -0.058 for women and -0.047 for men. The size of the IV estimates are therefore similar to the ones obtained using the longitudinal data. They also similarly decline when we increase the age cutoff. In the third column, we add time-region fixed effects to the IV specification to make sure that our findings are not driven by systematic differences in trends across regions. The coefficient estimates are similar to the ones presented in column 2, but we lose significance for men when we use an age cutoff of 40 because of an increase in the standard errors.

3.4.4 Potential Endogeneity Concerns

The findings obtained using cross-state regressions strongly corroborate the negative relationship between changes in cohort size and changes in marriage rates observed when longitudinal variation was used. However, there are reasons that prevent a causal interpretation of the relationship. While using number of births as our independent variable allows us to reasonably avoid reverse causality problems as well as important endogeneity concerns due to migration, one may nevertheless worry about omitted variables: state-level characteristics that drive changes in birth rates for a particular cohort as well as changes in marriage decisions of individuals that belong to that cohort 20 to 30 years later. Such omitted variables would have to be highly persistent shocks that affect growth in births in a given year as well as growth in subsequent marriage rates about 20 to 30 years later.

It is not easy to think of variables that fit this description. Potential examples include 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. A positive trend in men's earnings in some states fits this description. If children are a normal good, states with such positive trends may see both increased births in 1950 relative to 1940 as well as a greater number of marriageable men in 1980 relative to 1970. Alternatively, improving fertility technologies may have had a differential effect in states that are strongly religious or have stronger preferences for forming a family compared to states that do not. In states with weaker preferences for family, one might expect depressed birth rates as well as lower marriage rates in the future.

Note that in these and most credible cases we would typically expect an increase in both births and subsequent marriage rates or a decline in both variables. This bias would work against our favor and would result in a positive coefficient on cohort size, which is not what we find. Nevertheless, without exogenous variation in cohort size we cannot entirely eliminate the possibility that some biases could work in our favor.

3.4.5 Instrumental Variables Strategy

To address the potential endogeneity issues outlined in the previous subsection, we construct an instrument which is based on an idea first proposed by Bailey (2010). The idea is to use the interaction between the introduction of *Enovid* in 1957, later known as the the birth control pill, and cross-state variation in anti-obscenity laws, which limited the use of contraception, to generate exogenous variation in number of births and therefore cohort size.

In 1873, the U.S. Congress enacted the Comstock Act which had two main goals. The direct objective was to ban the interstate mailing, shipping, and importation of products and printed materials that were considered to be "obscenities". Since the Act considered anything employed for the prevention of conception an obscenity, it outlawed any interstate transaction involving contraceptives. The indirect objective was to "incite every State Legislature to enact similar laws" as stated by U.S. Representative John Merriman during an interview with the New York Times on March 15, 1873. The Comstock Act was highly successful in achieving this goal. By 1900, 42 states had approved anti-obscenity laws and by 1943 the number of states had increased to 48.

There is considerable variation across states in the type of anti-obscenity statutes that were enacted. As a consequence, these laws had different effects on the introduction of the pill in different states. As suggested by Bailey (2010), the states can be grouped into four categories depending on the type of law they enacted. The first group includes states that explicitly banned the sale, advertisement, and distribution of information of any product for the prevention of conception. This category includes seventeen states. The second group consists of all states that had the same ban on sales, advertisement, and distribution of information as the first group of states, but added an exception for physicians and pharmacists who were allowed to sell, advertise, and distribute information on products and materials related to birth-control methods. Seven states belong to this category. The third category includes states that explicitly only banned the advertisement and distribution of information of products and materials for the prevention of contraception, but did not outlaw their sale. Six states enacted this type of statute. The final group is composed of states that approved a law that banned the sale, advertisement, and dissemination of information of obscene products and materials, without explicitly classifying the prevention of conception as obscene. This category includes eighteen states. In our analysis, we refer to states in the first two groups as having a sales ban. We control explicitly for whether or not a state had a physician exception.

An important question is whether some states enacted stricter anti-obscenity laws because they had more conservative constituencies or because of other observable cross-state differences. Bailey (2010) provides evidence that this is not the case. For instance, among the states that adopted sales bans of contraceptives one can find both typically conservative and typically liberal states. California and Washington, two of the states that repealed anti-abortion laws before the Roe v. Wade decision, enacted the strictest version of the bans whereas Alabama, a generally conservative state, adopted a statute that did not explicitly categorize the prevention of conception as obscene.

These anti-obscenity state laws lasted until the sixties when they were repealed or struck down by the individual states or by the 1965 U.S. Supreme Court's decision in *Griswold v. Connecticut*. Specifically, two states repealed their anti-obscenity statutes in 1961, one state in 1962, four states in 1963, and Connecticut in 1965 after the U.S. Supreme Court's decision. *Griswold v. Connecticut* expedited the repeal of anti-obscenity statutes in all the remaining states between 1965 and 1971. In the empirical part, we follow Bailey (2010) and use the period between 1957 and 1965, the year of the U.S. Supreme Court decision, as the period in which the introduction of *Enovid* interacted with the anti-obscenity laws generated what is arguably exogenous variation in cohort size.

We will start by giving some descriptive and graphical evidence on the effect of the source of variation described above on our variables of interest. To do this, we divide the states in two groups: states that enacted sales bans of anti-conception methods and the remaining states. Before presenting the evidence, it is important to emphasize a difference between this paper and Bailey (2010). Bailey is interested in the relationship between the anti-obscenity laws and birth rates of married women after the pill was introduced. She finds that states with the sales ban experienced a marital birth rate that was 8% higher than the remaining states during the period 1957-1965. In this paper we are interested in the link between the anti-obscenity laws and the following three variables: cohort size at birth, cohort size at marriage age, and marriage rates. Differences in the growth in cohort size at birth between the two groups of states provide the first evidence of the effect of the Comstock laws on cohort size between 1957 and 1965. We find that in states with the sales ban, the growth in cohort size at birth was about three percent higher than in states with no ban.

We will now show graphically the relationship between the anti-obscenity laws and cohort size, starting with cohort size at birth followed by cohort size at marriage age. In Figure 3.6 we report the difference in growth of cohort size at birth between states with the sales ban and the remaining states from 1950 to 1970. We follow Bailey (2010) and present the graphical results separately for the four Census regions. There are two features that are worth discussing. First, after the introduction of the pill, in all regions, states with the ban experienced larger growth in cohort size at birth. The Figure suggests that in the South the states with the ban were reducing the gap in number of births with the states without ban before the introduction of the pill. But it also suggests that the process was expedited by the introduction of the pill. The second noteworthy feature is that in all regions, when states started to outlaw the sales bans on contraceptives, the growth in cohort size at birth in states with the ban started to converge to the growth in states with no ban. The convergence continues until 1965 when the *Griswold v. Connecticut* decision took place, at which point the two groups of states have similar rates of growth in cohort size. These two features taken together generate the hump shape in the difference in growth of number of births that characterizes all census regions.

In Figure 3.7 we replace cohort size at birth with cohort size at age 25, which represents a measure of cohort size at marriageable age. The figure displays patterns that are similar to the ones observed for cohort size at birth. Between 1957 and 1965, in all regions the difference in growth in cohort size at 25 between the two groups has the familiar hump shape that was observed in the previous graph. This indicates that cohort size at the time of marriage was affected in the expected way by the interaction of the Comstock laws with the introduction of the pill: it increased the growth in size of cohorts born in states with the ban in the period considered until the anti-obscenity laws started to be repealed. The growth in cohort size in these states then started to converge to the growth experienced by states with no ban.

Figure 3.7 describes graphically the first stage of an IV regression where the introduction of the pill interacted with the sales bans is used as an instrument. We will now formally use that variation in a standard IV setting. To do that, we construct two dummy variables. The first one, ban_s, is equal to one for all cohorts from state s that adopted a sales ban on contraceptives and zero otherwise. The second dummy variable, $ban * pill_{c,s}$, takes a value of one if a cohort c was born between 1957 and 1965 in a state s that enacted a contraceptive ban and zero otherwise. Then in the first stage we regress the n-year difference in log cohort size at the time of marriage on the two dummy variables and a set of controls, i.e.

$$\log \frac{y_{c,s}}{y_{c-n,s}} = \alpha + \beta_1 ban_s + \beta_2 ban * pill_{c,s} + \sum_{c,r} \pi_{c,r} + X' \gamma + \varepsilon_{c,s}, \tag{3.1}$$

where $y_{c,s}$ is the size of cohort c in state s, $y_{c-n,s}$ is the same variable for cohort $c-n$, $\pi_{c,r}$ are cohort-region fixed effects, and X' is a set of control variables that includes an indicator equal to 1 if the state had a physician exception, the physician indicator interacted with $pill_{c,s}$, and an indicator equal to 1 if the state enacted an advertising ban on contraception.

The results of the first stage are presented in Tables 3.4, where we report the effect of the anti-obscenity laws on 1-year, 3-year, 5-year, and 7-year differences in log cohort size. Consistent with the graphical evidence provided in Figure 3.7, we find that after the introduction of the pill and before the repeal of the Comstock laws, the sales ban on contraceptives had a positive and statistically significant effect on cohort size at marriage age in all cases. Consistently with the graphical evidence, the effect increases when we go from a 1-year difference to a 5-year difference. For the 1-year difference, a sales ban increases cohort size by 0.012%, whereas the 5-year differences raises cohort size by 0.041%. The coefficient on the 7-year difference is similar in size. The F-tests to evaluate the strength of the instruments are between 10.11 and 19.22 in our four specifications. An additional result is that the coefficient on ban_s is always small and statistically insignificant suggesting that the sales ban had no effect on cohort size before the introduction of the pill. This finding is consistent with Bailey's results which indicate that the Comstock laws had no effect on other forms of contraception.

In the second stage we use a specification similar to the one employed with the cross-state variation, i.e.

$$\log \frac{mar_{c,s}}{mar_{c-n,s}} = \beta_0 + \beta_1 \log \frac{size_{c,s}}{size_{c-n,s}} + \sum_{c,r} \pi_{c,r} + X'\gamma + \varepsilon_{c,s},$$

except that now we instrument cohort size growth with ban_s and $ban * pill_{c,s}$. Before presenting the results, it is important to remark that to construct the share ever married we must use the decennial Censuses since the CPS does not have enough state-level observations. The decennial Censuses have two limitations. First, in principle the share ever married can be computed for each cohort born in a particular state if one observes in the Census data a recall variable measuring the age at first marriage. Unfortunately, after 1980 this variable is not available in the Censuses. The second limitation is that in each decennial Census we only observe a particular cohort at a particular age. We therefore cannot directly compute the share ever married for each cohort. Instead, we rely on the the following strategy. In each Census, we first consider all individuals between the ages of 25 and 45. We then compute the share ever married for each cohort born in a particular state. Notice that we cannot use this variable directly in our regressions because it is affected by the age at which we observe a particular cohort in a particular Census. To address this issue, we regress the computed share ever married on age, state, cohort, and cohort-region dummies. We then remove the effect of age by subtracting the estimated coefficient on the age dummy multiplied by the dummy itself. Finally, we use the constructed variable in our regressions.

The second stage results are reported in Table 3.5 for men and women separately. The coefficient estimates have the expected negative sign, are statistically different from zero, and large in magnitude. They indicate that during the period considered a 1% increase in cohort size at marriageable age generated a reduction in marriage rates between 0.24% and 0.44%. The point estimates in the IV regressions are somewhat larger in size than the corresponding estimates using the longitudinal or cross-sectional variation. Because we have to construct the age-adjusted share ever married using individuals observed between the ages of 25 and 45, the coefficients are not directly comparable to the results obtained using the longitudinal and cross-sectional variation for share ever married by 30 or 40. Note, however, that in general we would expect the IV coefficients to be negative and larger in magnitude given the discussion in section 3.4.4 on potential endogeneity concerns, since most plausible omitted variables would bias the coefficients positively toward zero. We conclude that the IV findings are consistent with the results obtained using longitudinal and cross-state variation and they suggest that there is a causal relationship between changes in cohort size and changes in marriage rates.

We conclude this section with a discussion of a potential threat to our IV strategy. It is possible that the negative relationship between cohort size and share ever married we find in our IV regressions is generated by some type of selection process governing who becomes a mother in states without the ban after the introduction of the pill. The most serious hypothesis that could confound the interpretation of our results is that, after the introduction of the pill, mothers in states without the ban give birth to fewer children who are positively selected along some dimension. If those children are more likely to marry, as the literature suggests, our IV regressions will estimate a negative relationship between our two main variables. To evaluate this hypothesis, we follow Ananat and Hungerman (2012) and test whether children born in states where the pill was banned are more or less likely to have low birth weight. We employ the same specification used in the first stage of the IV estimation except that we use levels instead of differences with two new dependent variables: the share of children born with extremely low birth weight, which is defined as a birth weight below 1500 grams, and the share of children with low birth weight, which is a birth weight below 2500 grams. The estimation results are reported in Table 3.6. Using both dependent variables, the estimated coefficient on the interaction between the ban and the introduction of the pill is small and statistically insignificant. We find therefore no evidence that the initial access to the pill had an effect on the fraction of children born with low weight. Our result is different from the one obtained in Ananat and Hungerman (2012), where the authors find that initially access to the pill increased the share of children born with low weight. But the sample used is also different. Here, we consider the sample of married women, whereas Ananat and Hungerman (2012) study the behavior of single women younger than 21. The different results can therefore be rationalized by a more widespread early use of the pill by married women relative to young single women.[9]

Ananat and Hungerman (2012) also find weak evidence that early access to the pill had the effect of increasing the share of children born in poor families. To measure that effect, we would need micro data. Unfortunately, for our sample period the only micro data containing the required variables is the Census, which is only available for the decennial years and therefore cannot be used here. Observe, however, that this negative selection would be a threat to our IV estimates only if children born in low income families are more likely to marry. In this case, the negative selection would generate the negative relationship between cohort size and marriage rates we observe in the data. But the literature on household formation appears to rule out this alternative.[10]

The results discussed in this section indicate that there is a causal negative relationship between cohort size and marriage rates. The rest of the paper proposes and tests a potential mechanism that can explain this relationship.

[9] A second difference between our paper and Ananat and Hungerman (2012)'s paper is that we use the interaction between bans on contraception and the introduction of the pill as our main source of exogenous variation, whereas Ananat and Hungerman (2012) use restrictions on access to the pill for minors.

[10] For instance, the handbook chapter by Black and Devereux (2010) indicates that there is a positive intergenerational correlation in income and education. Moreover, Stevenson and Wolfers (2007) find no difference in the share ever married by education and therefore income. These two results suggest that children of low income parents are not more likely to marry.

3.5 A Dynamic Search Model of the Marriage Market

In this section we develop a dynamic search model of the marriage market. We decided to focus on this type of model for two reasons. First, it is one of the simplest models that can match the negative causal relationship between cohort size and marriage rates observed in the data. Second, the search model allows us to easily capture the dynamic nature of the marriage market which is an essential part of the mechanism that enables us to explain the link between cohort size and marriage rates. This explains why we did not consider 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), which is not as well suited to incorporate the dynamic aspects of the marriage market.

In the rest of the section, we first outline the main features of the model. We then show that the model can match the negative causal relationship between cohort size and marriage rates observed in the data. Finally, we derive an implication that will be used in the next section to test the model.

3.5.1 Characterization of the Model

The model characterizes an economy populated by $T+1$ overlapping generations of men and women. In each period $t = \{0, ..., T\}$, 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 of maximizing their lifetime utility.

We now introduce the main assumption of the model. We assume that women meet men 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. Two ideas form the basis for this assumption. First, women's fertile lifespan is shorter than men's since women are fertile only in their first part of their adult life, whereas men are fertile for most of their adult life. Second, one important benefit of marriage is that it is an effective arrangement for having and raising children. These two ideas imply that, with age, the value of getting married for a woman declines faster than the value for a man. Our assumption that this value for a woman is zero in the second part of her adult life is a special case of an economy in which the value of marriage for women and men follow this pattern. Our main 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 with a declining value of marriage and men to marry for more than two periods makes the model more complicated without changing the qualitative nature of the results.

The within-period utility of being single will be denoted by δ, whereas the within-period utility of being married for the couple as a whole will be denoted by η. 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. 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 old man chooses to be single in the second period, his lifetime utility takes the following form:

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

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

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

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 an old man and a woman, the man will die one period earlier. As a consequence, their 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, therefore,

$$w_{1,t}^m(\eta) = v_{1,t}^m + \gamma \left[v_{0,1,t} - v_{1,t}^m - v_{0,t}^w \right] = v_{1,t}^m + \gamma_m \left[\frac{1-\beta^T}{1-\beta} \eta + \beta^T \delta - v_{1,t}^m - v_{0,t}^w \right], \tag{3.2}$$

where the parameter $\gamma \in [0,1]$ allows for possible asymmetries in the way the marriage surplus is divided and $v_{1,t}^m$ and $v_{0,t}^w$ are the value of being single in this and future periods that were computed above. 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 can now solve the model starting with the decisions of a man of age 1 in period t. With probability $\theta_{1,t}^m$, he meets a woman 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. We can therefore determine the match quality η above which the couple will choose to marry. 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 between a woman and a man of age 1 is

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

We can now derive the expected value function for an old 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:

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

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

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

It is composed of three parts. The first part describes the value for the old man of meeting a woman with a match quality η sufficiently high that the couple will choose to marry multiplied by the corresponding probability. The second part characterizes the value of meeting a woman with a match quality η that is below the reservation value $\underline{\eta}_{1,t}$ times the probability of this event. Finally, the last part captures the value of not meeting a woman in the current period multiplied by the probability. By replacing $\underline{\eta}_{1,t} = 2\delta$, by dividing both sides of the equation by $\frac{1-\beta^T}{1-\beta}$, and by simplifying some of the terms, we obtain the following equation for the value function:

$$v_{1,t}^{m}\frac{1-\beta}{1-\beta^T} = \delta + \gamma\{E\left[\eta\,|\eta \geq 2\delta\right] - 2\delta\}\left(1 - F\left(2\delta\right)\right)\theta_{1,t}^{m}. \tag{3.3}$$

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

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

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:

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

We can now substitute for the expected value function of an old man using equation (3.3) and simplify some of the terms to obtain the following equation for the reservation value of a young man:

$$\underline{\eta}_{0,t} = 2\delta + \beta \frac{1-\beta^T}{1-\beta^{T+1}} \gamma \{E[\eta \,|\eta \geq 2\delta] - 2\delta\} (1 - F(2\delta))\, \theta^m_{1,t+1}. \tag{3.4}$$

Using $\underline{\eta}_{0,t}$, one can derive the expected value function for a woman and a young man. They are presented in Appendix A.

Before concluding this subsection, we want to remark that it would be straightforward to add congestion effects to our search model. Their advantage is that, if properly modeled, they mechanically amplify the effects generated by our standard search model. Their main limitation is that they are impossible to test without data on dating behavior which is not available for our sample period. For this reason, we have decided not to incorporate them in our model.

3.5.2 Steady State

In this subsection, we use the reservation values discussed above to solve for the steady state equilibrium in the marriage market. 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 young man meets a woman $\theta^m_{0,t}$ and the corresponding probability for an old man $\theta^m_{1,t}$. Let N^i_t be the number of individuals of gender i and age t who are present in the marriage market. Then $\theta^m_{0,t}$ and $\theta^m_{1,t}$ can be derived by noting that

$$\theta^m_{0,t} = \theta^m_{1,t} = \frac{N^w_{0,t}}{N^m_{0,t} + N^m_{1,t}}. \tag{3.5}$$

The probability $\theta^m_{0,t}$ is the correct measure in our model of what is called sex-ratio in the literature on household formation. The number of individuals of age 0 is exogenously given by the cohort size of a generation. However, the number of old men in the marriage market $N^m_{1,t}$ is endogenously determined by the decisions of young men. As a consequence, to derive $\theta^m_{0,t}$ and $\theta^m_{1,t}$ we need to solve for $N^m_{1,t}$. This variable can be computed as the number of young men who did not meet a woman at $t-1$ plus the number of young men who met a woman at $t-1$ but draw a match quality η lower than the reservation value, i.e.

$$N^m_{1,t} = N^m_{0,t-1}\left(1-\theta^m_{0,t-1}\right) + N^m_{0,t-1}\theta^m_{0,t-1}F\left(\underline{\eta}_{0,t-1}\right) = N^m_{0,t-1}\left(1-\theta^m_{0,t-1}\left(1-F\left(\underline{\eta}_{0,t-1}\right)\right)\right). \quad (3.6)$$

We can now replace for $\theta^m_{0,t-1}$ using (3.5) 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 old men in the marriage market $N^m_{1,t}$ are constant over time. We therefore have that

$$N^m_1 = N^m_0\left(\frac{N^m_0+N^m_1-N^w_0\left(1-F\left(\underline{\eta}_0\right)\right)}{N^m_0+N^m_1}\right).$$

We can now solve for N^m_1 and obtain

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

Generally, men and women have identical cohort size, i.e. $N^m_{0,t} = N^w_{0,t} = N_{0,t}$.[11] In this case the

[11]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 incarcerated 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 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 young men in steady state, we can substitute for θ_1^m in the equation that determines the reservation value (3.4). We can then derive, for the case in which $N_0^m \neq N_0^w$, the following equation for the steady state reservation value:

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

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

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

Note that $F\left(\underline{\eta}\right)$ is monotonically increasing in $\underline{\eta}$. As a consequence, there is a unique solution for $\underline{\eta}_{ss}$. Moreover, if men and women have identical cohort sizes, the steady state reservation value is independent of N_0^m and N_0^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 $N_0^m = N_0^w$.*

3.5.3 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 size 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 incarceration and mortality rates are higher for men of marriageable age than for women.

Suppose the economy is in steady state when it is hit by an unexpected shock in period $t = \tau$ that changes permanently the cohort size from N_0 to $N_0 + \Delta$. We consider the case of a permanent shock because in the data changes in cohort size tend to be persistent and even reinforcing. According to equation (3.5), 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 old men in steady state, i.e. $N_{1,\tau}^m = N_{0,\tau-1} F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}} = N_0 F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}$. Substituting for $N_{1,\tau}^m$ in the probabilities $\theta_{j,t}^m$, we have that in period τ

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

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. In our economy there are always more men than women in the marriage market. As a consequence, the probability that a woman meets a young man, $\theta_t^w = \frac{N_{0,t}}{N_{0,t}+N_{1,t}^m}$, is equivalent to the probability that a man meets a woman. Therefore, the previous result also implies that a positive cohort shock increases the probability that a woman meets a young men.

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 man compares the value of getting married at τ with the value of waiting until next period. The value of waiting depends on the probability he will meet a woman in period $\tau+1$. This probability depends on the number of old men at $\tau+1$, which can be written as follows:

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

Using equation (3.6), we can substitute for $N_{1,\tau+1}$ to obtain the following expression:

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

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

$$\underline{\eta}_{0,\tau} = 2\delta+\beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\{E[\eta\,|\eta\geq 2\delta]-2\delta\}(1-F(2\delta))\frac{1}{1+\left(1-\theta_{0,\tau}^m\left(1-F\left(\underline{\eta}_{0,\tau}\right)\right)\right)}. \tag{3.7}$$

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

$$\underline{\eta}_{0,ss} = 2\delta+\beta\frac{1-\beta^T}{1-\beta^{T+1}}\gamma\{E[\eta\,|\eta\geq 2\delta]-2\delta\}(1-F(2\delta))\frac{1}{1+\left(1-\theta_{0,ss}^m\left(1-F\left(\underline{\eta}_{0,ss}\right)\right)\right)}. \tag{3.8}$$

Earlier in this section we have shown that, with a positive shock to cohort size, $\theta_{0,\tau}^m > \theta_{0,ss}^m$. 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 and permanent shock to cohort size in period τ increases the reservation value $\eta_{0,\tau}$. A negative shock has the opposite effect.*

Proof. In the appendix. ■

Using Proposition 2 we can now show that a positive cohort shock reduces the fraction of women in a given cohort who marry and that a negative shock has the opposite effect. The following Proposition establishes this result.

Proposition 3 *A positive and permanent 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, old men become a scarce resource. This change has two effects. First, the fraction of women who marry mechanically declines because women are now less likely to meet old men, who have lower reservation values and, hence, higher probability of marriage. Second, young men become more selective because they will have a larger group of women to choose from when they will be old. As a consequence, the fraction of women who marry decreases. 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 can explain the negative relationship observed in the data between cohort size and marriage rates.

We will now derive an implication of the model that will be used in the next section as a test. In the following Proposition we shows that an increase in cohort size has the effect of decreasing the average difference in age at the time of marriage.

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 intuition behind Proposition 4 is based on the following two effects. First, women are less likely to meet old men because there are relatively fewer of them. Since the reservation value of old men does not depend on cohort size, a consequence of this is that the average age difference between spouses will decline. Second, women are more likely to meet young men because of the increase in cohort size. For a constant reservation value for young men, this effect would decrease the average age difference at marriage. However, young men become more selective in their decision to marry. As a result, fewer of the young men that meet a woman end up getting married. The combined effect of young men on the average age difference is therefore unclear. Proposition 4 shows that the effect generated by old men and the effect produced by a greater probability of meeting a young men dominate the effect of a higher reservation value. The increase in cohort size therefore reduces the average age difference at marriage.

Proposition 4 enables us to directly test the mechanism used by our model to generate the negative relationship between cohort size and marriage rates observed in the data. In our model the reduction in age difference between spouses subsequent to an increase in cohort size is generated by the change in the probability that a man meets a woman, $\theta^m_{0,t}$, or equivalently by a change in the sex ratio, which is at the core of our search model. It is difficult to think of an alternative well-specified and testable model that generates the same implication. The testable implication contained in Proposition 4 is therefore informative about the ability of our model to characterize the marriage market.

3.6 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 it can quantitatively match the variation in marriage rates observed in the data. In both parts we will only consider white individuals because, as argued in the previous section, the assumption $N_0^m = N_0^w = N_0$ is not satisfied in the data for the black population.

3.6.1 A Test of the Search Model

Proposition 4 establishes an implication of our search model. If the model is correct, an increase in cohort size should reduce the average age difference between spouses, 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.

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.

To implement the test, we construct the variable age difference between spouses as follows. First, we consider all married women of a given age. We then compute the difference between their age and the age of their spouse. Finally, we calculate the average for each cohort. The Data Appendix B provides details on the data sets and ages of married women used to construct this variable. It is worth remarking that we cannot perform the test using the introduction of the pill as an instrument. The reason for this is that the age difference cannot be constructed consistently for all cohorts born between 1945 to 1970 using the Censuses, the only data in which we have enough observations at the state level. To compute that variable for the required cohorts, we would have to use the recall variable age at first marriage, which is observed, jointly with a recall variable age of the first spouse, which is not observed.

In Figure 3.8 we report graphical evidence on the relationship between age differences at marriage and cohort size using longitudinal variation for whites. With the exception of the first twelve cohorts, Figure 3.8 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 declines. When the cohort size drops, the age difference between spouses becomes more negative and therefore increases. In our model, the change in age difference at marriage is generated by changes in the sex ratio interacted with the decision about whether to marry. A decline in age difference should therefore be accompanied by an increase in the sex ratio. To determine whether that is the case, in Figure 3.8 we also report the evolution of the sex ratio. This variable is constructed by dividing the number of young women by the number of young men plus the number of old men, where young women are all women between the ages of 18 and 29, young men are all men in the same age range, and old men are all men between the ages of 30 and 40. Given the way our variable is constructed, the first available observation is for 1932. The sex ratio evolves over time in a way that is consistent with our model. An increase in cohort size is accompanied by a rise in sex ratio and by a decline in the age difference between spouses.

The first column of Table 3.7 reports the coefficient estimate and R-squared obtained by regressing the logarithm of the average age difference between spouses on the logarithm of cohort size for whites using time-series variation. To avoid mixing observations from differenct data sets, we only consider cohorts born in 1930 or after which are all observed in the CPS. 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. It indicates 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.

In the second column of the same table we report the coefficient estimate obtained using cross-state variation. We follow a similar approach to the one used in the regressions estimating the effect of cohort size on the share ever married and instrument the changes in cohort size at adulthood with changes in cohort size at birth. The estimated coefficients for the decennial cohorts born between 1910 and 1970 are consistent with the search model. Our estimated coefficient is negative and statistically significant indicating that an increase in cohort size has a negative effect on the age difference between spouses. We therefore cannot reject the model developed in the previous section.

3.6.2 Estimation of the Search Model

In this subsection, we estimate the dynamic search model developed in this paper with the objective of evaluating whether it can quantitatively explain the changes in marriage rates observed in the data. This exercise enables us to additionally test 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 one of the most flexible distributions. Evidence of this is that many popular distributions like the uniform, the exponential, and the gamma are 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 3.5, 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 $[-0.2, 0.2]$.[12]

The third set of assumptions we make are related to the lifespan of an individual. In practice, individuals from a given cohort will participate in the marriage market 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 multiple cohorts

[12] Changing the interval does not change the outcome of the estimation.

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, five 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, consecutive increases in cohort size will have a larger effect than a single increase, because the newcomers compete in a marriage market that is progressively more crowded. A similar argument applies to declines in cohort size.[13]

As additional assumptions, we set the annual discount factor equal to 0.98 and consider a symmetric Nash-bargaining by setting γ equal to 0.5. Finally, we augment the model to allow for a fraction of men that are unwilling to marry independently of the value of match quality. We will denote this fraction with $1-\phi$. This parameter only affects the probability that a young or old man meets a woman. Specifically, 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 young or old man does not change since 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 $1-\phi$. These parameters are 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 three steps. For a given set of parameters that characterize the model, we first simulate the individual decisions. We then compute a function of the differences

[13] We have decided not to include this feature in the theory part because it makes the model more complicated without changing the insight it provides.

between some of the statistical moments that characterize the data and the corresponding moments obtained from the simulated data. Finally, the estimated parameters are obtained by minimizing this function.

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 use cohorts born from 1909 to 1929 to initialize the model. We therefore have 51 moments that will be matched using 4 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 that the value of being single δ is linked to the fraction of individuals never married in a given cohort. Everything else equal, a higher value of δ increases the share of individuals who choose to stay single in each cohort. The parameter ϕ enters equation (3.4) 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 young men and therefore on marriage rates if ϕ is larger. In the model, changes in meeting probabilities are mainly generated by variations in cohort size. The parameter ϕ can, therefore, be identified by measuring how the fraction of never married individuals 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.[14]

[14] We have proven that the parameter α_1 can be identified using as a moment the probability that an old man marries plus the corresponding probability for a woman; α_2 can be identified using as a moment the probability that a young 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

The estimated parameters are reported in Table 3.8. 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 old man marry only if $\eta > 2\delta$. Given our estimate of δ, this means that couples 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 old men marry only if they draw a relatively high value for match quality. In addition, observe that young men have a higher reservation value. Consequently, women paired with young men will also marry only if their match quality is relatively high. Finally, our model rationalizes the data by estimating that 13.3% of the men in the marriage market in a given period 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 3.9 the marriage rates observed in the data jointly with the marriage rates simulated using the estimated model, both as a function of cohort. Figure 3.9 shows that the model can quantitatively replicate the medium and long run variation in the share of never married across cohorts observed in the data. It is noteworthy that the search model can generate the big change in the share never married observed in the data between the early thirties and the late seventies, an increase from about about 5 percent in the thirties to about 17 percent seventies. This outcome is particularly remarkable since our model is very parsimonious. We only have four parameters to match 51 moments.

We conclude this section with an important remark. In this paper, we have provided evidence that positive changes in cohort size cause a decline in marriage rates and that negative changes have the opposite effect. We have also proposed a possible mechanism which cannot be rejected using our tests. However, as it is standard when a particular model is tested, we are not able to directly

a young man does not marry and of the probability that an old 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.

reject that changes in cohort size cause marriage rates to vary through different mechanisms. One possibility is that an increase in cohort size reduces income of that cohort and the reduction in income generates the decline in marriage rates. To evaluate this alternative, we can only rely on the evidence provided in other papers and described in section 3.2, which suggests that changes in income cannot explain changes in marriage rates for some of the periods we investigate in this paper. A thorough investigation of the effect of changes in cohort size on other economic and social variables is important, but it is left for future research.

3.7 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 cross-state variation in the adoption of the pill we provide evidence in support of the following two results. First, cohort size can explain on its own more than 50% of the variation in U.S. marriage rates. Second, an increase in cohort size reduces marriage rates and a decrease has the opposite effect. We cannot rule out the hypothesis that cohort size affects marriage rates through a third variable - for instance, income. We leave this for future research. However, if one believes that the cross-state variation in the introduction of the pill is exogenous, in the paper we have provided evidence that there is a causal relationship between cohort size and marriage rates.

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 a positive change in cohort size reduces the age difference between spouses, and a negative change increases it. Finally, we test the model in two different ways. We first test whether the derived implication can be rejected. In the data, an increase 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 that we document. The estimated model can match the long-run variation in marriage rates observed in the data.

Our results have important implications for policy analysis. In recent years, politicians and policy makers have begun to consider and implement policies that increase the fraction of married individuals with the intent of reducing the poverty rate. Several examples of such policies exist. For instance, Temporary Assistance to Needy Families (TANF) allows states to use a fraction of its funds to implement policies designed at increasing the share of married individuals. In West Virginia, households receiving TANF receive an additional 100 dollars a month if they are headed by a legally married household. States have also undertaken other policies to promote marriage. Former Utah governor Leavitt declared a “Marriage Awareness” week and created a commission whose goal was to develop strategies to promote marriage. “First Things First” in Tennessee and “Healthy Marriages Grand Rapids" in Michigan are other examples of programs designed to promote marriage. Our findings indicate that these types of policies may at best only have short term effects, since in the medium and long run marriage rates are for the most part outside the control of policy makers and politicians.

Appendix A: Proofs

Proof of Proposition 2

Consider a positive change to cohort size. According to equation (3.8), in steady state the reservation value of a young man is the solution to the following equation:

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

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:

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

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 (3.7) implies that $\underline{\eta}_{0,\tau} > \underline{\eta}_{0,ss}$.

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 young man times the probability she marries him plus the probability she meets an old man times the probability she marries him, i.e.

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

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

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

Proof of Proposition 4

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

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

Without loss of generality suppose the difference in age if a woman marries a young man is equal to y_1 whereas the corresponding difference in a marriage with an old man is equal to y_2 with $y_1 < y_2 = y_1 + z$. Let $P = \theta^w_{0,\tau}\left(1-F\left(\underline{\eta}_{0,\tau}\right)\right) + \left(1-\theta^w_{0,\tau}\right)\left(1-F\left(2\delta\right)\right)$ be the probability that a woman marries and let $x = \underline{\eta}_{0,\tau}$. Then, $E[\Delta age]$ can be written as follows:

$$
\begin{aligned}
E[\Delta age] &= \frac{y_1 \theta^w_{0,\tau}(1 - F(x)) + y_2\left(1 - \theta^w_{0,\tau}\right)(1 - F(2\delta))}{P} \\
&= y_1 + \frac{z\frac{N_1}{N_0 + N_1}(1 - F(2\delta))}{P} \\
&= y_1 + \frac{z}{\frac{N_0(1 - F(x))}{N_1(1 - F(2\delta))} + 1}.
\end{aligned}
$$

Notice that y_1, z, and $1 - F(2\delta)$ are constant with respect to changes in cohort size N_0. In addition, N_1 is also constant with respect to N_0 since it measures the number of old men in the marriage market the period of the shock which is not affected by contemporaneous cohort size.

To prove the proposition, we have therefore to prove that an increase in N_0 increases $N_0(1 - F(x))$. We will prove this result by contradiction. Consider a small change in cohort size Δ and suppose it reduces $N_0(1 - F(x))$. We will show that, if this is true, the number of old men in the period that follows the shock to cohort size is so large that the reservation value of young men in the period of the shock is smaller than the reservation value in steady state. This result contradicts the previous finding that an increase in N_0 increases x. Hence, an increase in N_0 must reduce $N_0(1 - F(x))$.

Denote with θ' and x' the probability that a man meets a woman and the reservation value of a young men both in the period in which the increase in N_0 occurs. Furthermore, denote with N_1' the number of old men the period that follows the increase in N_0. The change in N_1' can be computed as the number of young men who in this period do not meet a woman plus the number of young men who meet a woman and decide not to marry. Hence,

$$\begin{aligned} N_1' - N_1 &= (N_0+\Delta)\left(1-\theta'\right) + (N_0+\Delta)\,\theta' F\left(x'\right) - N_0(1-\theta) - N_0\theta F(x) \\ &= -(N_0+\Delta)\,\theta'\left(1-F\left(x'\right)\right) + N_0\theta\,(1-F(x)) + N_0 + \Delta - N_0 \\ &= -(N_0+\Delta)\,\theta'\left(1-F\left(x'\right)\right) + N_0\theta'(1-F(x)) + N_0\left(\theta-\theta'\right)(1-F(x)) + \Delta \\ &= -\theta'\left((N_0+\Delta)\left(1-F\left(x'\right)\right) - N_0(1-F(x))\right) + N_0\left(\theta-\theta'\right)(1-F(x)) + \Delta \\ &\geq N_0\left(\theta-\theta'\right)(1-F(x)) + \Delta, \end{aligned} \tag{3.9}$$

where the inequality follows from the assumption that $(N_0+\Delta)(1-F(x')) - N_0(1-F(x)) < 0$.

We will now show that $N_1' - N_1 \geq N_0(\theta-\theta')(1-F(x)) + \Delta$ implies that θ'', the probability that a man meets a woman the period after the shock takes place, is lower than θ. As a consequence, the reservation value in the period of the shock is lower than before the shock. Observe that

$$\theta = \frac{N_0}{N_0+N_1} = \frac{1}{1+\dfrac{N_1}{N_0}}$$

and

$$\theta'' = \frac{N_0+\Delta}{N_0+\Delta+N_1'} = \frac{1}{1+\dfrac{N_1'}{N_0+\Delta}}.$$

Hence, $\theta > \theta''$ if and only if $\dfrac{N_1'}{N_0+\Delta} > \dfrac{N_1}{N_0} = \sqrt{F(x)}$. Equation (3.9) implies that

$$\frac{N_1'}{N_0+\Delta} \geq \frac{N_1 + N_0(\theta-\theta')(1-F(x)) + \Delta}{N_0+\Delta}.$$

Simple algebra implies that $\dfrac{N_1 + N_0(\theta-\theta')(1-F(x)) + \Delta}{N_0+\Delta} > \dfrac{N_1}{N_0}$ if

$$\frac{N_0(\theta-\theta')(1-F(x)) + \Delta}{\Delta} > \frac{N_1}{N_0} = \sqrt{F(x)}.$$

Replacing for θ and θ', the left hand side of the inequality can be written as follows:

$$
\begin{aligned}
& \frac{N_0\left(1-F\left(x\right)\right)}{\Delta}\left(\frac{1}{1+\sqrt{F\left(x\right)}}-\frac{N_0+\Delta}{N_0+\Delta+N_0\sqrt{F\left(x\right)}}\right)+1 \\
= & \frac{N_0\left(1-F\left(x\right)\right)}{\Delta}\left(\frac{N_0+\Delta+N_0\sqrt{F\left(x\right)}-\left(N_0+\Delta\right)\left(1+\sqrt{F\left(x\right)}\right)}{\left(1+\sqrt{F\left(x\right)}\right)\left(N_0+\Delta+N_0\sqrt{F\left(x\right)}\right)}\right)+1 \\
= & \frac{N_0\left(1-F\left(x\right)\right)}{\Delta}\frac{-\Delta\sqrt{F\left(x\right)}}{\left(1+\sqrt{F\left(x\right)}\right)\left(N_0+\Delta+N_0\sqrt{F\left(x\right)}\right)}+1 \\
= & \frac{-N_0\left(1-F\left(x\right)\right)\sqrt{F\left(x\right)}+\left(1+\sqrt{F\left(x\right)}\right)\left(N_0+\Delta+N_0\sqrt{F\left(x\right)}\right)}{\left(1+\sqrt{F\left(x\right)}\right)\left(N_0+\Delta+N_0\sqrt{F\left(x\right)}\right)} \\
= & \frac{N_0+\Delta+N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+N_0F\left(x\right)+N_0F\left(x\right)\sqrt{F\left(x\right)}}{N_0+\Delta+2N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+N_0F\left(x\right)}
\end{aligned}
$$

Hence, we have the desired inequality if

$$
\frac{N_0+\Delta+N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+N_0F\left(x\right)+N_0F\left(x\right)\sqrt{F\left(x\right)}}{N_0+\Delta+2N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+N_0F\left(x\right)}>\sqrt{F\left(x\right)},
$$

or equivalently,

$$
\begin{aligned}
& N_0+\Delta+N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+N_0F\left(x\right)+N_0F\left(x\right)\sqrt{F\left(x\right)} \\
> & N_0\sqrt{F\left(x\right)}+\Delta\sqrt{F\left(x\right)}+2N_0F\left(x\right)+\Delta F\left(x\right)+N_0F\left(x\right)\sqrt{F\left(x\right)}.
\end{aligned}
$$

Some of the terms cancel out producing the following inequality:

$$
N_0+\Delta>N_0F\left(x\right)+\Delta F\left(x\right),
$$

which is equivalent to

$$
1>F\left(x\right),
$$

which is always satisfied. As a consequence, $\theta>\theta''$ which implies that the reservation value of young men in steady state is greater than their reservation value in the period of the shock, which contradict our result that the reservation value increases with an increase in cohort size. As a consequence, $N_0(1-F(x))$ must increase with cohort size. Hence, the expected value of the age difference at marriage declines with a positive shock to cohort size.

Expected Value Functions

For completeness, in this appendix we derive the expected values for young men and women. The expected value of a young man takes the following form:

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

The first term represents the value of meeting a woman with a match quality η higher than the reservation value times the probability of this event. The second term describes the value of meeting a women characterized by an η lower than the reservation value multiplied by the corresponding probability. The third term measures the value of not meeting a woman when young times the probability.

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:

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

F The first term measures the value of meeting a young man with an η higher than the reservation value times the corresponding probability. The second term is the value of meeting a young men whom it is optimal not to marry times the probability of this event. The third and fourth terms describe the same values for old men.

Appendix B: Data

Table 3.9 provides a summary of the datasets employed in the construction of the main variables of interest. In the rest of the appendix, we give additional details about how we construct the variables cohort size at birth, cohort size at marriageable age, share ever married, and age differences of spouses.

In the paper we use two different measures of cohort size: cohort size at birth and cohort size at marriageable age. Cohort size at birth is used in three ways: as the main independent variable when we employ longitudinal variation; as an instrument for cohort size at marriage age in the cross-state regressions; and as one of the variables used to determine the effect of the introduction of the pill in states with different anti-obscenity laws. With longitudinal variation we use cohort size at birth as the main independent variable instead of cohort size at marriageable age for two reasons. First, as shown in Figure 3.1, when cohort size is computed for the U.S. population there is little difference between cohort size at birth and cohort size at in adulthood, since migration from and to the U.S. was limited. Second, we can construct the variable cohort size at birth for cohorts born in 1909 and after. The variable cohort size at marriageable age can only be constructed for cohorts born after the 1940s. By using cohort size at birth we can therefore consider a larger number of cohorts without significant effect on the analysis. As indicated in Table 3.9, in the longitudinal analysis cohort size at birth is constructed using the U.S. Vital Statistics which provide information on this variable by race from 1909 to 1980. For the cross-state regressions, there are two data sets that can be used to measure cohort size at birth: the U.S. Vital Statistics which record births by race and by state from 1940; and the decennial Censuses which provide information on population counts from the beginning of the twentieth century to 2010. In the cross-state regressions, we work with the decennial cohorts 1910-1970. For consistency, rather than combining two different datasets, we use the Censuses over the entire period of interest. One limitation of the decennial Censuses is that population counts are published for 5-year age groups. From each decennial Census, we therefore record the number individuals between the ages 0 to 4 and we use it to construct the cohort size at birth. Our results do not change if we use data from the U.S. Vital Statistics for the 1940 to 1970 cohorts. In the regressions that use the introduction of the pill as an instrumental variable, we consider cohorts born between 1945 and 1970. We can therefore use the U.S. Vital Statistics to compute cohort size at birth for all of them.

Cohort size at marriageable age is used as the main independent variable in the cross-state regressions and in the regressions that use the introduction of the pill as an instrument. There are two datasets that can be used to measure this variable: the decennial Censuses and the SEER population estimates. SEER records cohort sizes at different ages starting from 1969. Hence, using this dataset we can construct cohort size at marriage age only for some of the decennial cohorts born between 1910 and 1970, which are the ones we consider in the cross-state regressions. For consistency, we therefore construct cohort size at marriage age by recording the number of individuals between the ages 20 and 24 in the decennial Censuses 1930-1990. We also experimented with the 5-year age group 30-34 with similar results. In the regressions that use the pill and the anti-obscenity laws as instruments, we use cohorts born between 1945 and 1970 which are all observed in SEER at age 25 or older. We therefore measure cohort size at marriage using the information in SEER at age 25.

The variable share ever marriage is constructed using a different procedure depending on whether we use longitudinal or cross-state variation. With longitudinal variation, we employ a combination of the CPS, which covers the period 1962-2011, and of the decennial Censuses. In the CPS, we observe the age and the marital status of each respondent. We can therefore easily compute the share ever married by age 30, 35, or 40 for each cohort born after a particular year. For instance, for the variable share ever married by age 30, we can use the CPS for all cohorts born on or after 1932; for the variable share ever married by age 40, we can use the CPS for all cohorts born on or after 1922. For cohorts born before those years, we use the 1960, 1970, and 1980 Censuses, which contain information on the marital status and the age at first marriage, a recall variable. Using these two variables, we construct the share ever married by age 30 and 35 for different cohorts by considering all individuals who in a given Census are between the ages of 30 and 45. We use a maximum cutoff age of 45 to avoid potential measurement errors due to differential mortality rates of married and non-married individuals. For the share ever married by age 40, we use the same procedure with a maximum cutoff age of 50. With cross-state variation, for all cohorts we only use information from the Censuses, as sample sizes in the CPS are too small to provide reliable estimates at the state level. In the longitudinal as well as in the cross-state variation, we cannot construct the share ever married for cohorts born before 1914 because the 1960 Census is the first one that records the age at first marriage.

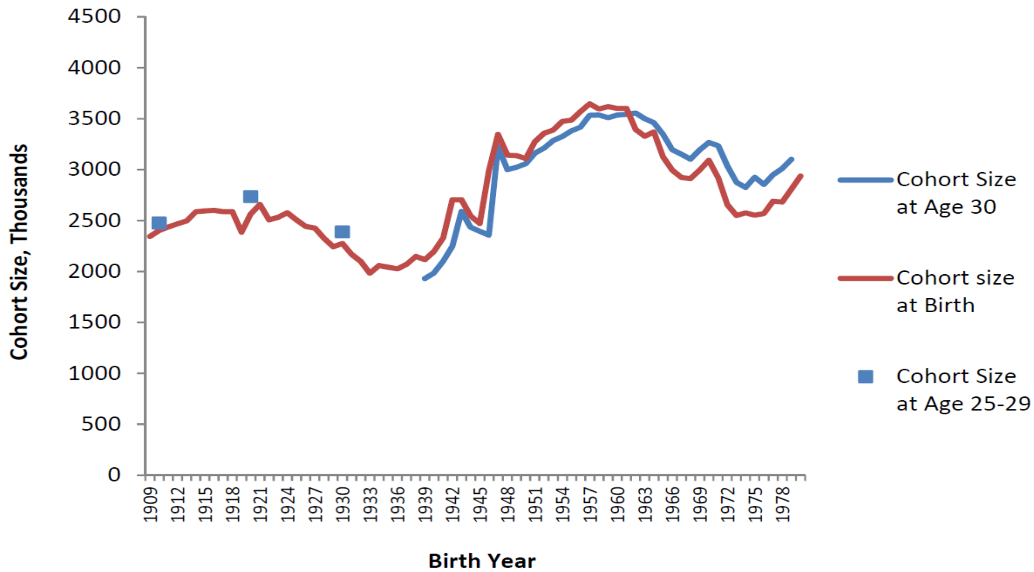
We use the variable "Relationship to household head" in the Census and CPS to record households in which a cohabiting partner is present. The Census began recording unmarried partners only in 1990, and the CPS only in 1995. As a result, in the longitudinal analysis we may miss cohabitations for cohorts born before 1965 when we use 30 as the age cutoff, or 1955 when we use 40 as the cutoff. In the cross-sectional analysis, we may similarly miss cohabitations for cohorts born before 1960 or 1950, depending on the age cutoff. In the data we observe that cohabitation for early cohorts is limited. For the 1965 cohort, the share of individuals cohabiting at age 30 was 2.8% For the 1955 cohort, the share cohabiting at age 40 was 0.76%. We examined data in the

National Survey of Families and Households (NSFH) to test whether we miss a substantial number of cohabitations for the cohorts for which we do not have the cohabitation variable, especially at the lower age cutoffs. The first wave of the NSFH (1987-1988) is nationally representative and provides retrospective data on marriage and cohabitation. We use the dataset to examine cohabitation patterns at age 30 for cohorts born 1957 or earlier. We found that cohabitation at age 30 is almost non-existent for pre-baby boom cohorts. From 1945 to 1957, the average share of individuals cohabiting is 0.5%. We conclude that we only marginally underestimate the share ever married or cohabiting at age 30 for the early baby boom cohorts.

The variable age difference between spouses is used to test the search model using both longitudinal and cross-state variation. With longitudinal variation, we use cohorts born between 1930 and 1975 so that we can use the CPS to construct this variable. Specifically, for each cohort we consider all women between the ages of 30 and 35 who are married. We then compute the difference between their age and the age of their spouse. Finally, we calculate the average for each cohort. When we employ cross-state variation, the average age difference is computed using the 1940-2000 Censuses, since the CPS does not have enough observations at the state level. In this case, for each decennial cohort born between 1910 and 2000 we consider all women of age 30 to 35 who are married, compute the age difference with their spouse, and calculate the average at the state level.

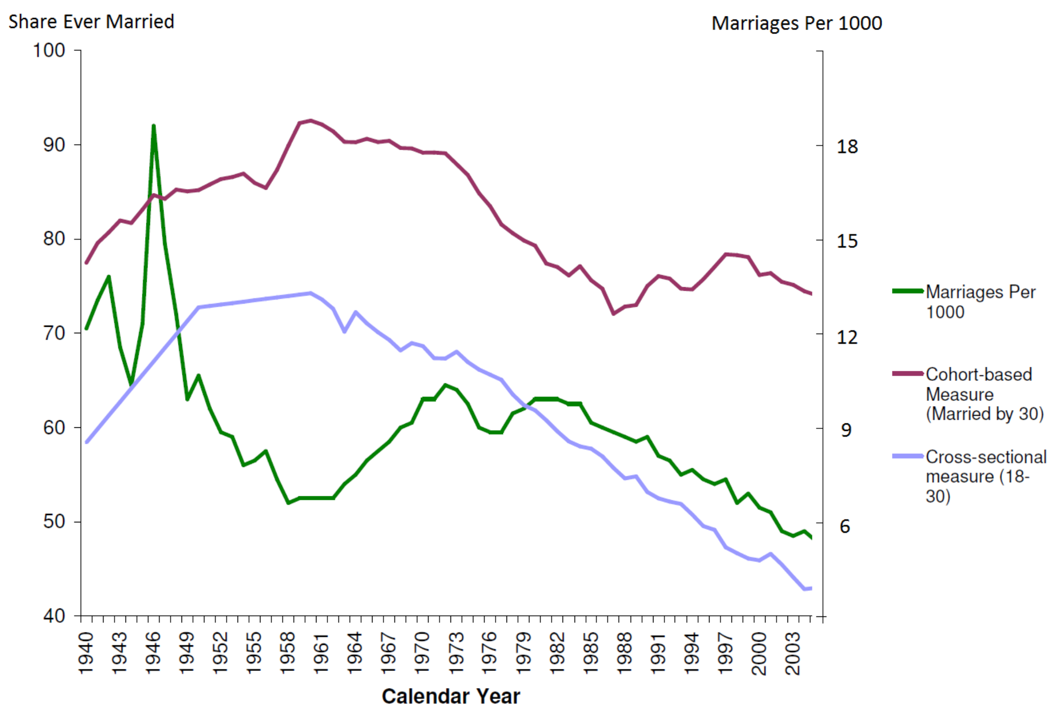
Tables and Figures

Figure 3.1: Cohort Size at Birth and at Age 30



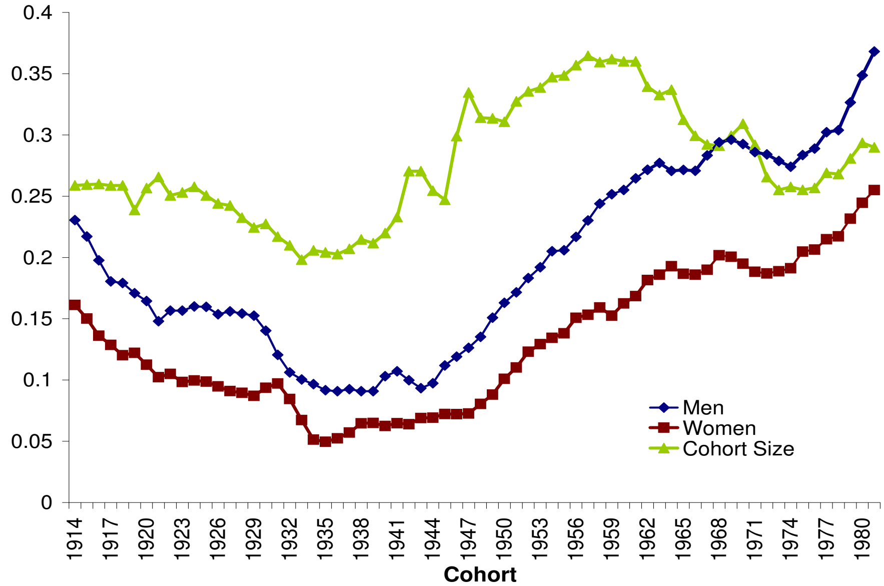
White individuals only. Sources: Vital Statistics of the U.S.; NIH SEER population estimates (1969-2010); U.S. Census 1940-1960.

Figure 3.2: Different Measures of Marriage Rates

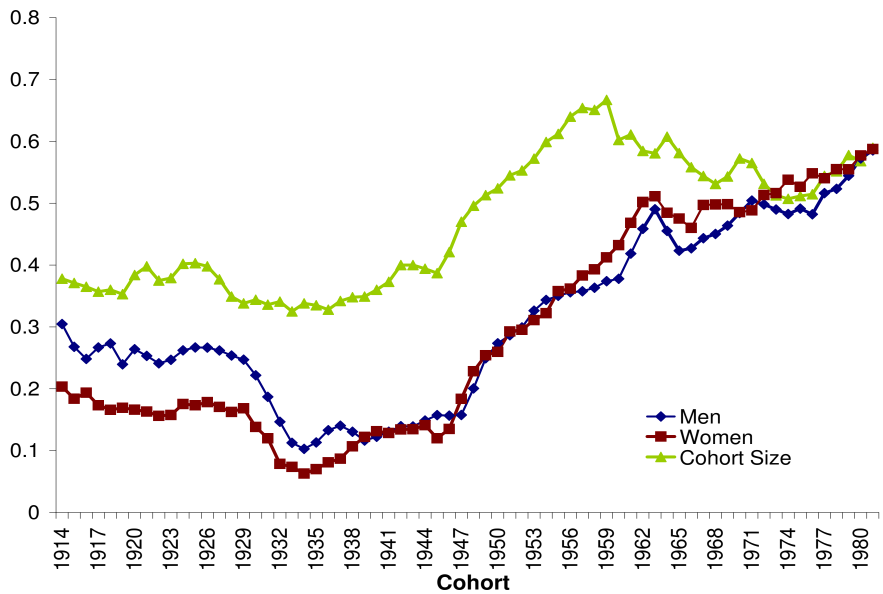


Sources: Vital Statistics of the United States; IPUMS CPS, 1962-2005; IPUMS USA, 1940-1960. The left vertical axis marks the percentage of individuals ever married for the cohort-based and cross-sectional measures; the right axis corresponds to marriages per thousand. To facilitate direct comparison, cohabitations are not accounted for in any of the three measures.

Figure 3.3: Share Never Married By 30



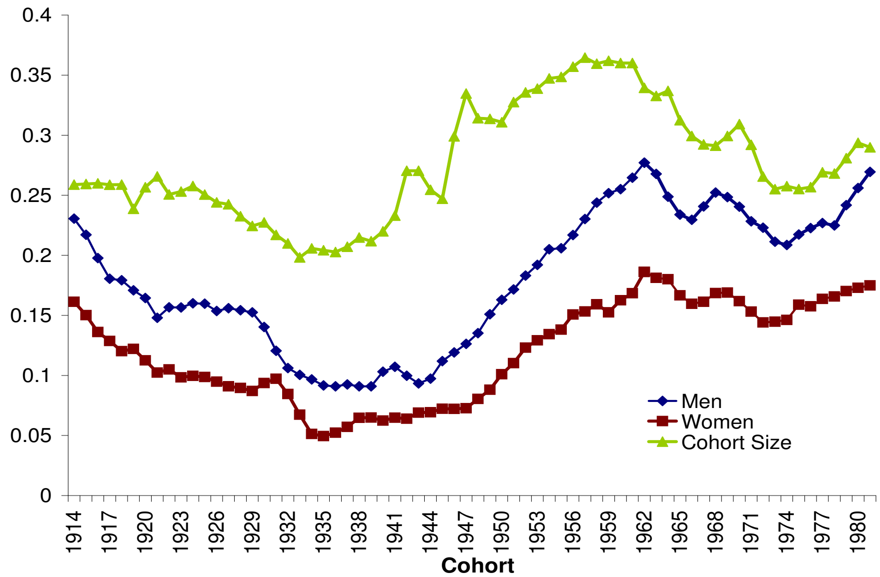
(a) White



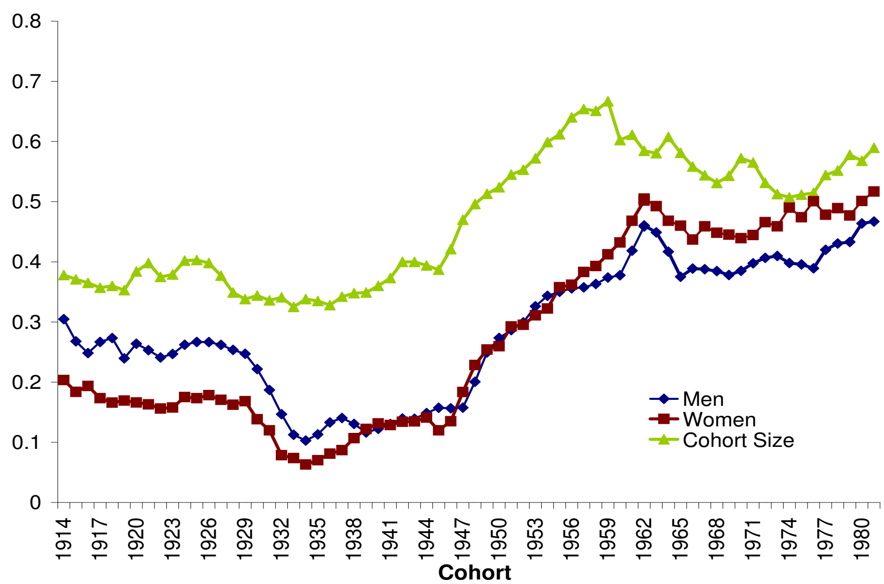
(b) Black

Notes: The vertical axis represents both the percentage of individuals ever married as well as normalized cohort size. In Panel A we normalize cohort size by dividing by 10,000,000; in Panel B by 1,000,000. For share ever married, we graph three-year moving averages. Sources: Vital Statistics of the United States; IPUMS CPS, 1962-2011; IPUMS USA, 1960-1970.

Figure 3.4: Share Never Married and Not Cohabiting By 30



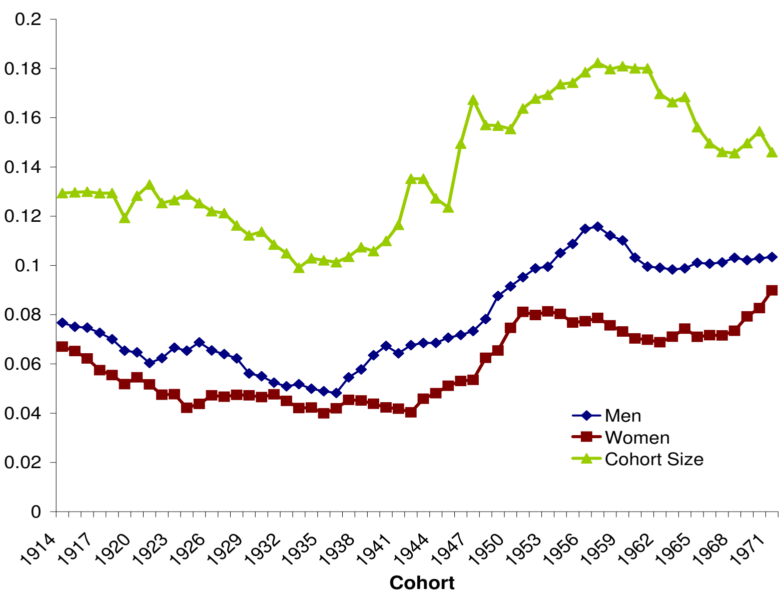
(a) White



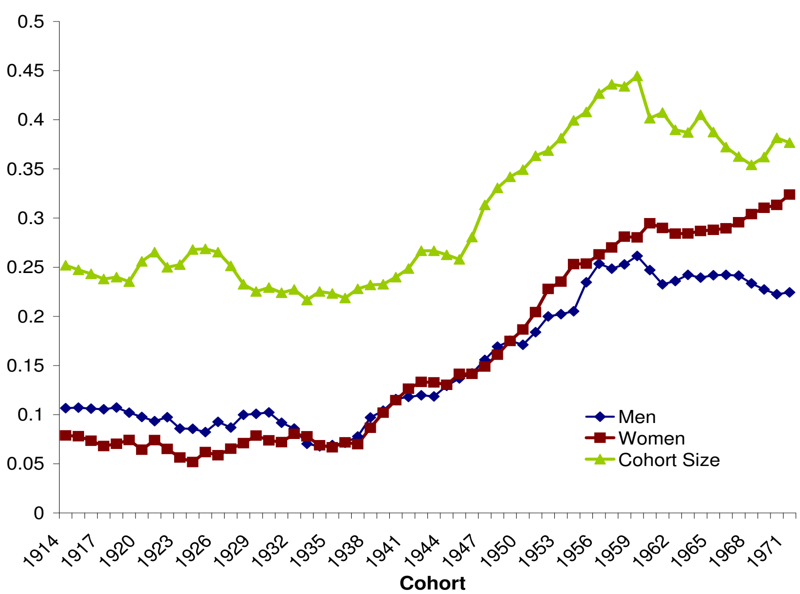
(b) Black

* See note in Figure 3.3.

Figure 3.5: Share Never Married and Not Cohabiting By 40



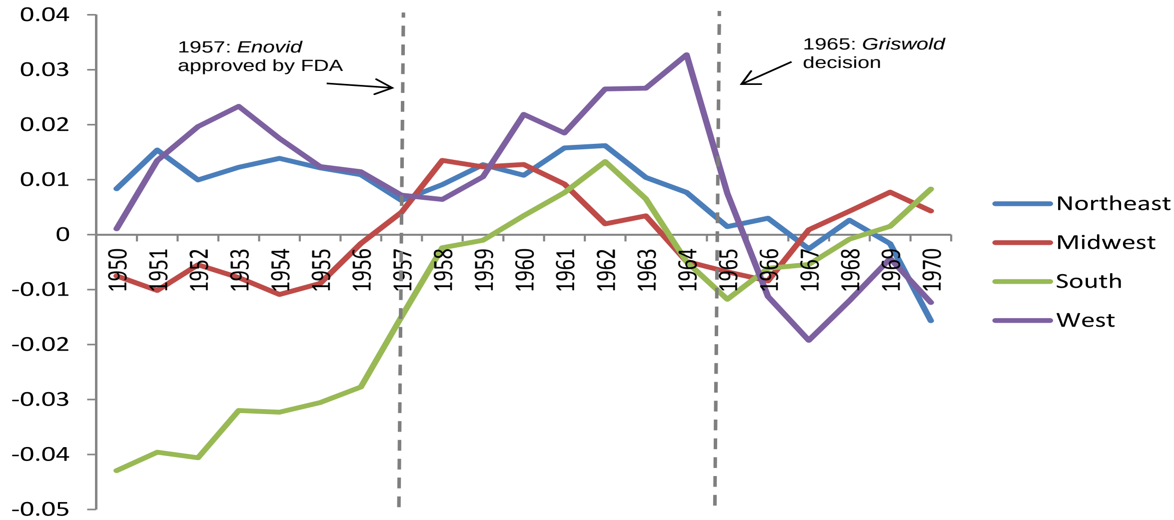
(a) White



(b) Black

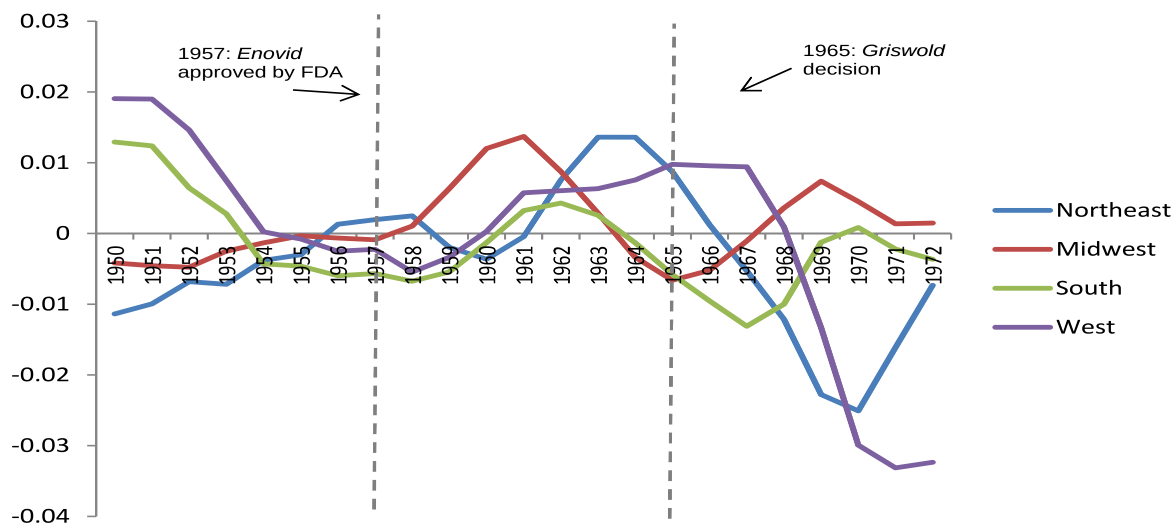
* See note in Figure 3.3.

Figure 3.6: Growth of Total Births, by Region: States with Sales Bans - States without Sales Bans



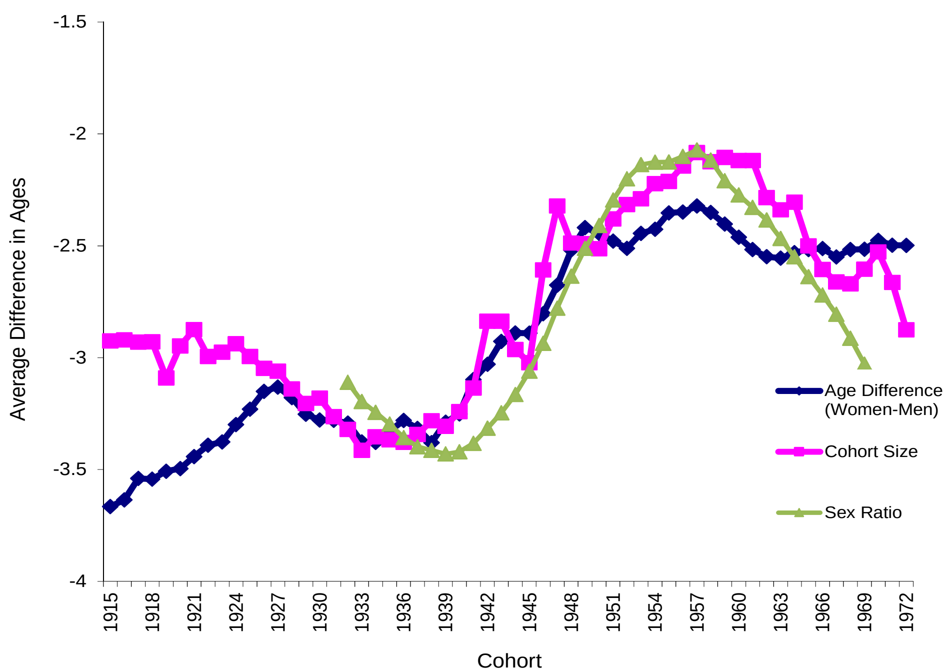
Notes: Three-year moving averages. Sources: National Vital Statistics. Bailey (2010).

Figure 3.7: Growth of Total Adult Population at Age 20, by Region: States with Sales Bans - States without Sales Bans



Notes: Three-year moving averages. Sources: NIH SEER population estimates. Bailey (2010).

Figure 3.8: Age Difference Between Spouses by Cohort, Whites



Sources: Vital Statistics of the United States; IPUMS CPS, 1962-2011; IPUMS USA, 1950-1960.

Figure 3.9: Observed and Simulated Never Married Rates for Women

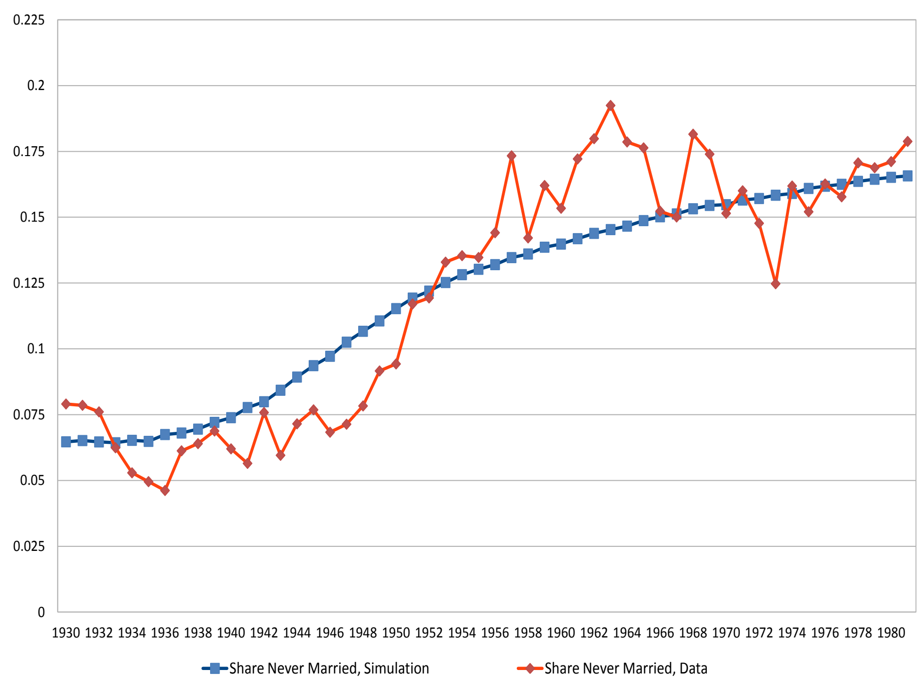


Table 3.1: Time Series Regression of Log Share Ever Married on Log Cohort Size

	White Men	Black Men	White Women	Black Women
Ever married	-0.294***	-0.592***	-0.193***	-0.870***
by age 30	(0.037)	(0.048)	(0.026)	(0.064)
R^2	0.50	0.70	0.46	0.74
Ever married	-0.182***	-0.440***	-0.111***	-0.560***
by age 35	(0.015)	(0.031)	(0.011)	(0.043)
R^2	0.71	0.77	0.65	0.74
Ever married	-0.107***	-0.322***	-0.066***	-0.453***
by age 40	(0.008)	(0.019)	(0.007)	(0.026)
R^2	0.77	0.84	0.58	0.85

*** Significant at 1%. Notes: Newey-West standard errors in parentheses. Each coefficient is the outcome of a separate regression. Regressions include cohorts born after 1914 until the most recent cohort observed at a given age in 2011. The number of observations in each regression is equal to 68 for the share ever married by 30, 63 for the share ever married by 35, and 58 for the share ever married by 40. Sources: IPUMS CPS 1962-2011, IPUMS Census 1960-1970.

Table 3.2: Regression of Change in Log Share Ever Married on Change in Log Cumulative Cohort Size

	1-Yr.	2-Yr.	3-Yr	4-Yr	5-Yr.	7-Yr.	10-Yr.
Men	0.049	0.017	-0.157**	-0.220***	-0.233***	-0.238***	-0.277***
By Age 30	(0.037)	(0.062)	(0.061)	(0.049)	(0.045)	(0.040)	(0.047)
Men	-0.054	-0.057	-0.090**	-0.101***	-0.098**	-0.115***	-0.163***
By Age 35	(0.054)	(0.044)	(0.040)	(0.034)	(0.026)	(0.023)	(0.017)
Men	0.073***	0.037	-0.018	-0.048**	-0.054**	-0.052**	-0.085***
By Age 40	(0.016)	(0.038)	(0.029)	(0.022)	(0.023)	(0.023)	(0.021)
Women	-0.064***	-0.077***	-0.101***	-0.095***	-0.103***	-0.125***	-0.156***
By Age 30	(0.024)	(0.037)	(0.014)	(0.019)	(0.021)	(0.024)	(0.028)
Women	-0.005	-0.053**	-0.070***	-0.084***	-0.081***	-0.090***	-0.113***
By Age 35	(0.056)	(0.021)	(0.011)	(0.012)	(0.015)	(0.016)	(0.015)
Women	-0.011	-0.005	-0.008	-0.022*	-0.036**	-0.056***	-0.073***
By Age 40	(0.021)	(0.021)	(0.012)	(0.013)	(0.014)	(0.010)	(0.006)

* Significant at 10%. ** Significant at 5%. *** Significant at 1%. See notes in Table 1. Newey-West standard errors in parentheses.

Table 3.3: Cross-Sectional Regression of Log Share Ever Married by 30 or 40

Dependent Variable: 10-Yr. Difference in Log Share Ever Married

10-Yr. Dfference in Log Cohort Size	OLS	IV (1)	IV (2)
Men	-0.041**	-0.123***	-0.168***
By Age 30	(0.019)	(0.041)	(0.056)
R^2	0.807	0.791	0.819
Men	-0.029***	-0.047**	-0.040
By Age 40	(0.010)	(0.018)	(0.028)
R^2	0.555	0.551	0.595
Women	-0.041***	-0.104***	-0.115***
By Age 30	(0.010)	(0.023)	(0.036)
R^2	0.555	0.765	0.796
Women	-0.032***	-0.058***	-0.048**
By Age 40	(0.010)	(0.017)	(0.023)
R^2	0.539	0.528	0.605
First Stage Results			
Log Cohort Size at Birth			0.440***
			(0.073)
F-test			36.44
R^2			0.831

* Significant at 1%. ** Significant at 5%. *** Significant at 1%. Notes: Robust standard errors in parentheses. $N = 288$. Each coefficient is the outcome of a separate, population-weighted regression. We control for cohort fixed effects, and for cohort-region fixed effects in IV-(2). Regressions are for decennial cohorts born between 1910 and 1970. Data includes all states except Hawaii and Alaska. Sources: US Census Population Counts, 1910-1990; IPUMS USA, 1940-2010.

Table 3.4: Comstock Laws and N-Year Differences in Log Cohort Size

Dependent Variable: N-Yr. Difference in Log Cohort Size

	1-yr	3-yr	5-yr	7-yr
$Ban * Pill_{c,s}$	0.012**	0.035***	0.041***	0.040**
	(0.005)	(0.011)	(0.015)	(0.017)
Ban_s	-0.002	-0.003	0.006	0.018
	(0.005)	(0.010)	(0.012)	(0.014)
N	1248	1152	1056	960

* Significant at 10%. ** Significant at 5%. *** Significant at 1%. Notes: Robust standard errors in parentheses. Regressions are weighted by population and include controls for physician exception, physician exception interacted with "pill", advertising bans, and cohort-region fixed effects. Source: IPUMS USA, 1980-2000. NIH SEER Population Counts.

Table 3.5: N-Year Differences in Log Share Ever Married and N-Year Differences in Log Cohort Size

Dependent Variable: N-Yr. Difference in Log Share Ever Married (Men)

	1-yr	3-yr	5-yr	7-yr
N-yr Difference	-0.389*	-0.241**	-0.239***	-0.260***
in Log Cohort Size	(0.206)	(0.101)	(0.091)	(0.080)
N	1248	1152	1056	960

Dependent Variable: N-Yr. Difference in Log Share Ever Married (Women)

	1-yr	3-yr	5-yr	7-yr
N-yr Difference	-0.441***	-0.311***	-0.302***	-0.304***
in Log Cohort Size	(0.161)	(0.093)	(0.0845	(0.076)
N	1248	1152	1056	960

* See note in Table 3.4.

Table 3.6: Comstock Laws and Birth Weight

Dependent Variable: Birth Weight

	Share with Birth Weight < 1500	Share with Birth Weight < 2500
$Ban * Pill_{c,s}$	0.0001025	-0.0004241
	(.0002037)	(.0007311)
Ban_s	-0.000462	0.0006242
	(0.0003064)	(0.0006155)
N	1006	1006

* See note in Table 3.4.

Table 3.7: Regressions: Log Age Difference and Log Cohort Size

	Time Series, 1930-1975	Cross-State, 1910-1970
Log Cohort Size	-0.592***	
	(0.043)	
10-Yr. Difference in Log Cohort Size		-0.241***
		(0.092)
N	46	288
R-squared	0.81	

*** Significant at 1%. Notes: Robust standard errors in parentheses. The cross-state IV regression controls for cohort-region fixed effects; data includes all states except Hawaii and Alaska. Sources: US Census Population Counts, 1910-1970; IPUMS USA, 1940-2000.

Table 3.8: Estimated Parameters

Parameters	Estimates	Standard Errors
First Shape Parameter	0.020	[0.011]
Second Shape Parameter	0.072	[0.044]
Value of Being Single	0.107	[0.039]
Fraction of Men Unwilling to Marry	13.3	[0.159]

Table 3.9: Data Sets Used in the Construction of the Main Variables

Variable	Variation of Interest	Dataset
Cohort Size at Birth	National, Yearly	Vital Statistics of the United States (1909-1980)
	State Level, Decennial Years	U.S. Decennial Census, 1910-1970
Cohort Size at Adulthood	National and State Level, Yearly (Around the Introduction of the Pill)	NIH/SEER Population Estimates (1969-2010)
	State-Level, Decennial Years	U.S. Decennial Census, 1930-1990
Share Ever Married	National, Yearly	CPS (1962-2011), U.S. Decennial Census, 1960-1970
	State Level, Decennial Years	U.S. Decennial Census, 1940-2000
Age Differences of Spouses	National, Yearly	CPS (1962-2011), U.S. Decennial Census, 1960
	State Level, Decennial Years	U.S. Decennial Census, 1970-2000

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