Cohort Size and the Marriage Market:

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

Mary Ann Bronson† Maurizio Mazzocco‡

March 2021

Abstract

We document that the U.S. marriage market is characterized by two systematic empirical patterns. First, there is a quantitatively large, strong and persistent negative relationship between changes in cohort size and marriage rates of women. Second, the same negative correlation holds for men. This relationship accounts for a large share of the variation in marriage rates over time and across states. We then establish the features a model should possess to generate the two patterns. We start with a standard matching model with search frictions. We show that it is rejected by the data because it produces a negative relationship for women, but a positive relationship for men. However, the standard model can rationalize both patterns if either of the following two features is added: (i) the marriage surplus deteriorates with cohort size; (ii) cohort size reduces the efficiency of the matching function. We show that the two models have different predictions for the relationship between changes in cohort size and out-of-wedlock births, with model (i) predicting a positive relationship and model (ii) a negative relationship. We find that only the model in which cohort size affects marriage surplus is not rejected.

*We are grateful to participants at various seminars and conferences for helpful comments. Research reported in this paper was supported by the National Institutes of Health under Award Numbers 2T32HD7545-11 and 5T32HD7545-10. The content is solely the responsibility of the authors and does not necessarily represent the official views of the National Institutes of Health.

†Georgetown University, Dept. of Economics, Washington, D.C. Email: mary.ann.bronson@gmail.com.

‡UCLA, Department of Economics, Bunche Hall, Los Angeles, CA. Email: mmazzocc@econ.ucla.edu.

1 Introduction

Why do marriage rates vary over time? Since marriage rates have important implications for many socio-economic variables – including fertility rates, children's welfare, labor force participation, and income inequality – providing an answer to this question is important for all social scientists engaged in studying their evolution.[1] This paper makes two main contributions that can help answer the question.

The first contribution is to document that the U.S. marriage market is characterized by two empirical patterns. First, there is a large, strong, and persistent negative relationship between changes in cohort size – the number of people in a given birth year – and changes in marriage rates for women (fact 1). Second, a similar negative relationship holds for men (fact 2). We document these two empirical patterns using time-series variation, cross-state variation, and arguably exogenous variation generated by the introduction of the birth-control pill. Any theory of the evolution of marriage rates must be able to explain these two patterns.

The second contribution is to establish the features a model should possess to rationalize these facts. We start with a standard dynamic matching model of the marriage market with search frictions, in which men and women differ in their fertile lifespans. In the model, changes in cohort size affect marriage rates by varying the supply of men and women in the marriage market. We rely on a dynamic model with search frictions instead of a static frictionless matching model à la Becker, because it enables us to better account for the dynamic nature of the marriage market.

Using the model we establish the following results. First, it can explain fact 1, but not fact 2, since it predicts a positive relationship between changes in cohort size and marriage rates of men. It is therefore rejected by the data. Second, the matching model can explain both facts if one of the following two features is added: (i) the marriage surplus – the difference between the value of being married and the values of being single – declines with cohort size; or (ii) the efficiency of the matching function – which determines meeting probabilities – declines with cohort size. Lastly, we show that the two augmented models have opposite predictions about the relationship between changes in cohort size and out-of-wedlock births. In the data, these two variables display a positive relationship, indicating that only model (i) is not rejected.

[1] Killingsworth and Heckman (1986), Moffitt (2000), Angrist and Lavy (1996), Gruber (2004), McLanahan and Percheski (2008) are examples of papers analyzing the relationship between marriage rates and these variables.

The previous studies closest to ours are Akers (1967), Schoen (1983), and Bergstrom and Lam (1989). These papers study the effects of changing cohort size on marriage behaviors, but do not establish the systematic empirical relationship with marriage rates (facts 1 and 2). Using demographic models, Akers and Schoen predict the same positive relationship between changes in cohort size and men's marriage rates that is predicted by our standard matching model. Their models are therefore rejected by the data. Bergstrom and Lam can only study the effects of changing cohort size on age differences between spouses, since they assume constant marriage rates.

A large theoretical and empirical literature provides alternative explanations for the variation over time and across space in marriage rates: changes in income (Cherlin (1981), Hill (2011)); advances in fertility technologies (Goldin and Katz (2002), Akerlof, Yellen, and Katz (1996)) and household production technologies (Greenwood and Guner (2009)); changes in labor market opportunities for men and welfare aid policies (Wilson (1987), Ellwood and Crane (1990)); shifts in incarceration policies (Charles and Luoh (2010)); the relative income hypothesis (Easterlin (1987)); changes in sex ratios - the relative number of women in the marriage market – (Angrist (2002), Seitz (2009), Abramitzky, Delavande, and Vasconcelos (2011), Knowles and Vandenbroucke (2016)). While these explanations may account for part of the variation in marriage rates over time and across states, our paper indicates that there are general and persistent patterns in marriage rates that social scientists must take into account to understand their evolution.[2]

The paper proceeds as follows. In Section 2, we describe the data sets used to derive the empirical results. Section 3, documents our empirical findings. In section 4, we develop the dynamic model of the marriage market, present the theoretical results, and report the outcome of the test. Section 5 concludes.

2 Data

Our empirical analysis employs two main variables: the share ever married by 30, 35, and 40, and cohort size. The variable share ever married is constructed by cohort using IPUMS USA (1940-2010) or IPUMS CPS (1962-2011). We use the share ever married by a given age as our main measure of the evolution of the marriage market, because of the limitations that characterize alternative variables. Bronson and Mazzocco (2017) provide evidence that the two alternative measures commonly used in the literature – the number of new marriages

[2]Appendix B provides a detailed discussion of the related literature.

per population and the share of individuals currently or ever married within an age range – generally lead to misleading inference when used to study the probability someone marries during his or her life or fertile life, how it evolves, and how it differs across populations.

Cohort size at marriageable age (proxied by its size at ages 20-25) is constructed using Census Total Population Counts (1910-1980), and the Survey of Epidemiology and End Results (SEER) population estimates (1969-2000). Information on cohort size at birth, used in some of our IV results, is obtained from the National Vital Statistics (1909-1980) and Census Total Population Counts (1910-1980). Cohort size at birth is also the main independent variable for our longitudinal analysis, as the variable cohort size at marriageable age cannot be constructed for most cohorts born prior to 1940. In the appendix we document that, when cohort size is computed for the U.S. population, differences between cohort size at birth and cohort size at marriageable age are small, since net migration to the U.S. was limited during the time period we consider (Figure A.1). Appendix A describes the exact procedure used to construct all the variables employed in the analysis.

3 Empirical Results

In this section, we document the negative relationship between cohort size and marriage rates for both women and men first using time-series variation and then variation across states. We only consider the white population because 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 due to higher mortality and incarceration rates. We leave the investigation of the marriage market for the black population for future research.

3.1 Changes in Marriage Rates Over Time

Figure 1 provides graphical evidence on the two empirical patterns. In Panel A, 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. We plot the share *never* married because visually it is easier to detect a positive correlation between the two variables. For cohorts born before 1960, there is a strong and positive correlation between cohort size and the share never married for both women (fact 1) and men (fact 2). 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" between the mid-1940s and early 1960s, the period in which the cohorts born in the twenties and thirties were active

in the marriage market. Between 1946 and 1959, rapidly growing cohort size during the post-war baby boom is associated with a share never married that nearly tripled.

The correlation between our two main variables weakens for cohorts born after 1960. These cohorts were active in the marriage market from the 1980s, when cohabitation started to become a popular form of household formation and potentially a close substitute for marriage. In Panel B, we plot the variables reported in Panel A, with the exception that cohabiting individuals are treated as married instead of being grouped with never-married people. Once we account for cohabiting households, the relationship between cohort size and household formation is observed over the whole period. In Figure B.1 in the Appendix, we report a similar positive and strong correlation between cohort size and share never married or cohabiting by 40.

In column one of Table 1, we report the average relationship between marriage rates and cohort size at different age cutoffs in our longitudinal data, obtained by estimating the coefficients from a regression of log share ever married or cohabiting on log cohort size.[3] For ease of exposition, in the rest of the paper we will consider the effect of cohort size on the share ever married instead of the share never married. Moreover, given the patterns documented in Figure 1, Panel B, we continue to treat cohabiting households as married.[4] The elasticities decrease with age and are higher at 30, for both sexes, indicating that an increase in cohort size is associated with two effects: (i) a decrease in the eventual share ever married or cohabiting by 40; and (ii) an increase in the age at first marriage. At the cutoff age of 40, we estimate that the cohort size variable accounts for more than half of the variation observed in marriage rates.

Columns 2-4 of Table 1 evaluate whether year-on-year variation in cohort size affects marriage rates, or whether more persistent changes are required. The columns describe the effect of n-year cumulative differences in cohort size on n-year differences in marriage rates. To account for the effect of adjacent cohorts, for $n > 1$ we use differences in cumulative cohort size as our independent variable, where cumulative size for the cohort born in period t for the n-year difference is constructed by adding up cohort size from $t-n+1$ to t.[5] Column two reports the average of the estimates obtained using the one- and two-year differences, which we interpret as the short term effect; column three and four describe the average

[3] We use Newey-West standard errors for all significance tests. We verify using a Johansen test that the non-stationary time series we regress are cointegrated. The test rejects the null hypothesis that the series are not cointegrated at the one-percent level and that the estimated relationship is spurious.

[4] Our results are similar if cohabiting couples are excluded.

[5] 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.

estimates for the medium term – three- to five-year differences –, and the long term – seven- and ten-year differences. In Table B.1 in the Appendix, we report the estimated coefficients for each n-year difference.

Coefficients on cohort size are small and insignificant for short-term changes, but they become large, negative, and significant for persistent medium-term changes, and further increase for long-term changes. This indicates that changes have to cumulate for longer than one or two years to generate significant fluctuations in marriage rates. The long-run effects, by age 40, are large and imply that a standard deviation increase in cohort size decreases the share ever married by a third of a standard deviation.

3.2 Changes in Marriage Rates Across States

We now provide evidence on facts 1 and 2 using cross-state variation. If changes in cohort size influence marriage rates, then states with larger increases in cohort size should experience larger drops in marriage rates.

Changes in cohort size at marriageable age across states are partly driven by cross-state migration. Migration flows are typically related to differences across states in economic and social conditions, which affect marriage rates. Migration can also influence marriage decisions by skewing sex-ratios. Cohort size at marriageable age across states has therefore the potential to be endogenous. To address this issue, we use as an instrument for the size of the marriage market the number of births in a given cohort and state, which are arguably unaffected by the endogeneity concerns discussed above.

To analyze state-level changes, we rely on the decennial Census, since it is the only dataset with sufficiently large sample sizes for all states. In Section 2, we describe 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. We focus on decennial cohorts born between 1910 and 1970, since in the Census the share ever married at 30 and 40 can only be computed for them. Specifically, we perform the empirical analysis by first constructing ten-year log differences for the share ever married, cohort size at birth, and cohort size at marriageable age for each state and cohort. We then pool all cross-sections, because of the small number of observations, and regress differences in log share ever married on differences in log cohort size, where in our main specifications the latter is instrumented using log cohort size at birth. In all regressions, we add cohort fixed effects.

Table 2 presents the results using standard OLS regressions (column 1), IV regressions

(column 2), and IV regressions that control additionally for cohort-region fixed effects (column 3). Similarly to the findings obtained using longitudinal variation, the OLS coefficients on changes in cohort size are negative and statistically significant for both men and women. As argued above, the OLS estimates may be biased if migration is partially driven by the desire to find better earning opportunities. If higher income also increases marriage probabilities, the estimates will be positively biased, away from the strongly negative relationship we expect. The same is true if individuals migrate to find a spouse. The IV regressions address these issues.

The first stage results are reported underneath the second stage results. The instrument cohort size at birth explains a large fraction of cohort size at marriageable age. Nevertheless, the coefficient of 0.44 on log cohort size at birth is well below 1, suggesting that cross-state migration significantly affects changes in cohort size at marriageable age. Our second stage estimates are negative, statistically significant and, as expected, larger in size than the OLS estimates. They indicate that the effects on marriage rates are substantial. For example, from 1940 to 1950, and from 1950 to 1960, cohort size increased by 42% and 17%, respectively. Our estimates imply that the change from the 1940 to 1950 cohorts should be associated with a 3.4 percentage point decrease in the share of women ever married by 30. The corresponding, implied percentage point decrease for men is 3.8. The observed increase in cohort size from 1950 to 1960 implies a 1.5 percentage point decrease in marriage rates for women and a 1.4 decline for men. This is equivalent, on average, to about 55% and 39% of the actual changes in share ever married by 30 observed for women and men over this period.

Using number of births as an instrumental variable avoids reverse causality problems and endogeneity concerns due to migration. But one may nevertheless worry about omitted variables that drive changes in birth rates in a state in one decade as well as changes in marriage decisions of individuals from that birth cohort 20 to 30 years later. Such variables would have to be highly persistent shocks, e.g. positive trends in earnings in some states. If children are a normal good, states with such positive trends may see increased births in one decade, and a greater share of high-earning, marriageable men thirty years later. Alternatively, states may differ in their religiosity or preferences for forming a family. In states with weaker preferences for family, one might expect depressed birth rates as well as lower marriage rates in the future. In these and most credible cases we typically expect an increase in both births and subsequent marriage rates or a decline in both variables, resulting

in a positive coefficient on cohort size, biased away from the negative coefficient we find. Nevertheless, without exogenous variation in cohort size we cannot entirely eliminate the possibility that some biases could work in our favor.

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

The instrument exploits three historical events. The first event is the introduction of laws banning so-called "obscenities" across U.S. states, starting in the early 1870s. These statutes varied in their specific wording, with many states explicitly naming contraceptives as an obscenity. Therefore, when *Enovid* was introduced in 1957, the second historical event, only some states allowed its sale for contraception. Anti-obscenity laws lasted until they were struck down by the 1965 U.S. Supreme Court's decision in *Griswold v. Connecticut*, the third historical event, after which contraceptives could be sold everywhere. Following Bailey, we therefore employ an IV strategy that exploits the fact that between 1957 and 1965, the availability of *Enovid* in only some states – due to historical anti-obscenity laws – generates what is arguably exogenous variation in total births, and therefore cohort size. Bailey provides detailed evidence that whether or not a historical statute included wording banning sales of contraceptives appears to be idiosyncratic.[6]

Twenty-four states had statutes that explicitly banned the sale of products for contraception. We refer to these states as having a sales ban. Seven of those states granted some exceptions to physicians and pharmacists, which we control for explicitly. All remaining states either banned only advertisement (but not sale) of contraceptives, or did not explicitly list contraceptives among obscenities.

We start by giving graphical evidence on the effect of the source of variation described above on our variables of interest. We follow Bailey (2010) and present the results separately for the four Census regions. In Figure 2, Panel A, 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. Two features generate the hump shape that characterizes all census regions. First, in all regions, after the introduction of the pill, states with the ban experienced larger growth

[6]There is no apparent relationship between how conservative a constituency is and how strict the enacted anti-obscenity law is. California and Washington, two of the states that repealed anti-abortion laws before the Roe v. Wade decision, enacted the strictest version of the contraceptive bans whereas Alabama, a generally conservative state, did not categorize contraceptives as obscenities.

in cohort size at birth. In the South, states with the ban were already growing relative to states without a ban before 1957, but the process was expedited by the introduction of the pill. Second, 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.

In Figure 2, Panel B, 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 in Panel A, with the differences in growth in cohort size following the same familiar hump shape between 1957 and 1965. Thus, the differences in cohort size at birth driven by the Comstock laws and the introduction of the Pill persisted to generate differences in cohort size at marriageable age.

We now formally use the introduction of the pill interacted with the sales bans as an instrument for cohort size at marriageable age. 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. In the first stage, we then regress the n-year difference in log cohort size at marriageable age on the two dummy variables and a set of controls, i.e.

$$\log \frac{y_{c,s}}{y_{c-n,s}} = \alpha + \beta_1 ban_s + \beta_2 ban * pill_{c,s} + \sum_{c,r} \pi_{c,r} + X'\gamma + \varepsilon_{c,s}, \tag{1}$$

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 first stage results, presented at the bottom of Table 3, indicate that, between 1957-1965, the sales ban on contraceptives had a positive and statistically significant effect on cohort size at marriageable age, which increases as we go from a 1-year difference to a 5-year difference. The coefficients for the 5-year and 7-year differences are similar in size, and indicate 4% higher growth in cohort size in states with the sales ban between 1957 and 1965, consistent with what Bailey finds for birth rates. The F-tests to evaluate the strength of the instruments are between 10.11 and 19.22 in our four specifications. Finally, the coefficient

on ban_s is always small and statistically insignificant. This suggests that the sales ban had no effect on cohort size before the introduction of the pill, consistent with Bailey's results, and indicates limited effect of the ban on other forms of contraception.

In the second stage, we use a specification similar to the one employed previously in the cross-state analysis, 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 $ban * pill_{c,s}$ rather than with number of births.

To construct the share ever married we must use the decennial Censuses since the CPS does not have enough state-level observations. 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. Without this recall variable, the share ever married cannot be computed directly for each cohort, because in each decennial Census we observe different cohorts at different ages. To address this limitation, we rely on 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, and 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 at the top of Table 3 for men and women separately. The coefficient estimates have the expected negative sign, are statistically different from zero, and large in magnitude for both women and men. 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 percent. The point estimates in the IV regressions are larger in size than the corresponding estimates obtained using longitudinal and cross-sectional variation. This result should be expected given the previous discussion on potential endogeneity concerns, which suggests that the most plausible omitted variables would bias the coefficients positively toward zero. The findings using IV regressions are therefore consistent with the longitudinal and cross-state results which indicate that fact 1 and 2 are important patterns of the marriage market.

We conclude with a discussion of potential threats to our identification strategy. We

begin by noting that the evidence presented in Table 3 rules out a weak instrument problem. The main threat to our IV strategy is therefore a failure of the exclusion restriction to hold. One possibility is that the introduction of Enovid in 1957 and the repeal of contraceptive bans in 1965 coincided with other events or policy changes that potentially affected marriage, two or three decades later, differentially in states that did and did not ban contraceptives. We are not aware of any such policies or coinciding events. Changes to divorce, age of majority, and abortion laws all occurred starting in the late 1960s and early 1970s, after the period we study. By the time these latter changes occurred, contraceptive bans had already been systematically repealed. A second threat to identification is that the historical Comstock laws were correlated with how conservative states were on later policies, such as access to abortion. Our discussion in footnote 6 indicates that there is no evidence of such correlation.

A last potential threat is that our instrumental variable affects subsequent marriage behaviors through mechanisms other than cohort size alone. One possibility is that the negative relationship between cohort size and share ever married is generated by the selection process governing who becomes a mother in states without the ban after the introduction of the pill. If mothers in states without the ban gave birth to fewer children who are positively selected along some dimension and these children are more likely to marry, our IV regressions will estimate a negative relationship between our two main variables. We test this hypothesis by following the analysis in Ananat and Hungerman (2012). They evaluate the effect of state legislations that changed age of majority, and hence access to oral contraceptives, on the share of children born with low weight – a strong predictor of future economic outcomes – to women under age 21. They find that expanded access to the pill produced by laws that reduced the age of majority had the adverse effect of increasing the share of children born with low weight to young women.

Following their empirical strategy, we test whether children born in states where the pill was initially banned are more or less likely to have low birth weight. We employ the same specification we use in the first stage of the IV estimation except that, to make our results comparable with Ananat and Hungerman (2012)'s findings, we use levels instead of differences for the following 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 results reported in Table 4 indicate that the introduction of the pill had no effect on the

fraction of children born with low weight to women of all ages. These findings are in line with those documented in Bailey (2013), who similarly finds no evidence that the birthweight of infants born in the 1960s changed differentially in states where selling the Pill was legal.

There are two explanations for the differences between our and Ananat and Hungerman's results. First, we focus on contraceptive bans that primarily affected women who had already reached the age of majority or were married. By contrast, Ananat and Hungerman focus on laws that reduced the age of majority below 21 and increased access to the pill for unmarried minors. We therefore consider a different and broader population. A second possible explanation is that our analysis focuses on the 1960s and not on 1970s, the period studied by Ananat and Hungerman. In the 1970s, states started to legalize abortion. Ananat and Hungerman's results may therefore be driven not just by increased access to the pill by minors, but also by increased access to abortion. The results reported by Myers (2017) support this conclusion. She documents that the age of majority laws passed in the early 1970s affected primarily young women's take-up of abortion rather than contraceptives' use, reversing the findings by Goldin and Katz (2002) on the effects of early access to contraceptives on marriage and education decisions. Since Ananat and Hungerman's paper uses the same policy variation as Goldin and Katz, changes in fertility patterns attributed to the effects of the pill may in fact be attributable to abortion.[7]

4 A Dynamic Matching Model with Search Frictions

What factors can explain facts 1 and 2? To answer this question, we introduce a standard dynamic matching model of the marriage market with search frictions.

We consider an economy populated by $T + 1$ overlapping generations. In each period t a new cohort of N men and women is born and lives for $T + 1$ periods. We assume that

[7] Ananat and Hungerman (2012) also find weak evidence that reduction of age of majority had the effect of increasing the share of children born in poor families, suggesting that the young women in their sample period who prevented or reduced unwanted pregnancies may have been positively selected. Bailey (2013) tests for such selection during our relevant sample period in the 1960s using data from the Integrated Fertility Survey Series (IFSS). Bailey documents that Pill usage in the early 1960s was concentrated among women in married households, and does not find evidence that reductions in unwanted births were higher for highly-educated women. It is important to note that, while we find no evidence of an increase in the share of children born in poor families in states with contraceptive bans, an increase of the kind documented in Ananat and Hungerman (2012) 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. The literature on household formation appears to rule out this alternative. 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.

women and men have the same cohort size, because it is a good approximation for the white population. If single, a man or woman experiences within-period utility δ, which is constant across individuals. We denote the within-period utility of being married for a couple as a whole by $\hat{\eta}$ (match quality), and the marital surplus by $\eta = \hat{\eta} - 2\delta$. Match quality $\hat{\eta}$ is drawn from a probability distribution $\hat{F}(.)$ and the corresponding distribution for marital surplus η is denoted by $F(.)$. The distributions do not vary across couples or over time.

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 individuals then decide whether to marry with the objective of maximizing their lifetime utility. Once married, they make no further decision. 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 or until they divorce. Divorce can occur in each period with probability q, which is independent of within-period utility, time, and age.[8] A couple can freely divide the gains from marriage using a Nash-bargaining solution, with the parameter $\gamma \in [0, 1]$ determining how the surplus is divided. Future utilities are discounted at the discount factor $\beta \leq 1$.

Single women meet a potential spouse with a positive probability only in the first period of their adult life, while single men meet a partner with a positive probability in the first two periods. This assumption is based on two observations. First, women's fertile lifespan is shorter than men's period of fertility. Second, an important benefit of marriage is that it is an effective arrangement for having and raising children. These two observations imply that the value of getting married for a woman declines faster with age than the value for a man. Our assumption that this value is zero for a woman in the second period of her adult life is a special case of this general idea. Under this assumption, the marriage market is populated by younger women (age 0) and younger and older men (age 0 and 1); and changes in cohort size affect the composition of the marriage market by changing the probabilities with which men and women meet. Allowing women to marry for $\tau - 1$ periods and men to marry for τ periods adds complexity without changing the qualitative nature of the results.[9]

The solution of matching models with search frictions is generally provided in terms of reservation values. In our model, the relevant reservation value is the marital surplus η at which a pair of potential spouses is indifferent between marrying or staying single. In our context, two types of couples can form: couples in which the woman is younger than

[8] Gousse, Jacquemet, and Robin (2017) consider a matching model with endogenous divorce.

[9] This is not the first paper to use differences between women's and men's years of fertility to develop a model of the marriage market. For instance, Siow (1998) uses a similar idea.

the man; and couples in which the woman and the man are both of age 0. Our model is therefore characterized by two reservation values.

In Appendix D.1, we show that a woman of age 0 and a man of age 1 have a reservation value equal to zero, i.e. $\underline{\eta}_{1,t} = 0$. This is intuitive. Since this is their last opportunity to meet a potential spouse, they marry if they draw a match quality $\hat{\eta}$ greater than the sum of their values of staying single, 2δ, and hence if $\eta = \hat{\eta} - 2\delta > 0$.

The reservation value of a couple in which the woman and man are both of age 0 is slightly more complicated to derive because a man of age 0 has the option value of waiting until next period and drawing a new potential spouse. In Appendix D.1, we show that the existence of this option value generates a reservation value equal to

$$\underline{\eta}_{0,t} = B\theta^m_{1,t+1}, \tag{2}$$

where $\theta^m_{1,t+1}$ is the probability that a man meets a woman when older, which corresponds to the ratio of women to men in the marriage market, and B is a positive constant that depends only on the parameters of the model. The derivation of (2) shows that the option value is included in the term $B\theta^m_{1,t+1}$, which measures the probability that a younger man will meet a woman when older, multiplied by the share of the expected marital surplus he will receive, times the probability he will choose to marry her. Thus, the reservation value for this type of couple increases with the probability that a man meets a woman when he is older.

We now characterize the steady state equilibrium in the marriage market. The following Proposition determines its main features.

Proposition 1 *In steady state, there is a unique reservation value for a couple formed by a younger woman and a younger man, which does not depend on cohort size and takes the following form:*

$$\underline{\eta}_{ss} = B\frac{1}{1 + F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}}.$$

The marriage market has therefore a unique steady state equilibrium.

Proof. See Appendix D.2. ■

We now study the effect of a change in cohort size on marriage rates. We will focus on the case in which the shock to cohort size is unexpected and permanent. Similar results apply if the permanent shock is known with certainty. We consider the case of permanent shocks because in the data changes in cohort size tend to be persistent and even reinforcing.

Suppose the steady state economy is hit by a shock in period $t = \tau$ that changes permanently cohort size from N to $N + \Delta$. In Appendix D.3, we show that, if the shock is positive, the reservation value of younger men $\eta_{0,\tau}$ increases, whereas a negative shock has the opposite effect. The intuition behind this result is straightforward. With a permanent increase in cohort size, men are more likely to meet a woman when older. Thus, the option value of waiting until next period for younger men increases and with it their reservation value.

Using this result, the following Proposition establishes the effect of a shock to cohort size on the marriage rate of women.

Proposition 2 *A positive and permanent shock to cohort size in period τ reduces the fraction of cohort-τ women who get married. A negative shock in period τ has the opposite effect.*

Proof. See Appendix D.4. ■

For insight behind this result, note that an increase in cohort size has two effects. First, older men become a scarce resource. As a consequence, the fraction of women who marry mechanically declines because they are less likely to meet older men, who have lower reservation utilities. Second, younger men become more selective because they will have a larger group of women to choose from when they are older. Thus, the second effect also generates a decrease in the fraction of women who marry. 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 matching model can explain fact 1.

The following Proposition establishes the sign of the relationship between changes in cohort size and changes in marriage rates for men.

Proposition 3 *A positive and permanent shock to cohort size in period τ increases the fraction of cohort-τ men who get married. A negative shock in period τ has the opposite effect.*

Proof. See Appendix D.5. ■

Proposition 3 contains a negative result. Since the matching model generates a positive relationship for men between changes in our two main variables, it cannot produce fact 2. This result is not obvious a priori. An increase in cohort size has two different effects on men. First, men's probability of meeting a woman increases. Second, younger men become more selective. The first effect goes against our empirical findings since it implies an increase

in the marriage rate of men when cohort size increases. But the second effect is in favor of the empirical pattern, since it generates a decline in the marriage rate of men. Proposition 3 establishes that the first effect always dominates, therefore rejecting the standard matching model.

Generalizing the Matching Model with Search Frictions. In the standard model, cohort size affects the probability of marriage only through the probability that a man and a woman meet, which is determined by their relative number in the marriage market. Proposition 3 establishes that, if this is the only effect of cohort size, the marriage patterns documented in this paper cannot be matched. To match them, cohort size has to influence marriage through additional channels.

The structure of the matching model points to two plausible channels. Cohort size could affect the efficiency of the matching technology, e.g. through congestion effects (Petrongolo and Pissarides (2001)). Or, the probability distribution of marital surplus could depend on cohort size. This may be true for several reasons. Larger cohorts may have lower educational returns (Card and Lemieux (2001)); experience lower parental investments (Black, Devereux, and Salvanes (2005)); or make fewer pre-marital investments (Angrist (2002)).

To account for these channels, we develop two modified versions of the matching model. The first one generalizes the matching technology by allowing the matching probabilities to depend on a function of cohort size $\lambda(N)$, which determines the efficiency of the matching technology at different N. Specifically, we replace the matching probabilities in the standard model $\theta^i_{a,t}$ with the meeting probabilities $\lambda(N)\,\theta^i_{a,t}$, with $0 \leq \lambda(N) \leq 1$ and $a = 0, 1$.

The second model generalizes the probability distribution of marital surplus, $F(\eta)$. To simplify the exposition, we consider the case in which a change in cohort size shifts the distribution of marital surplus to the left, i.e. $\eta' = \eta - \phi(N)$, for some function $\phi > 0$.

The following Proposition determines the assumptions under which these generalized models can generate both fact 1 and fact 2.

Proposition 4 *Suppose that the matching model with search frictions is generalized in one of the following two ways:*

1. *The distribution of marital surplus deteriorates when cohort size increases, i.e.* $\frac{\partial \phi(N)}{\partial N} > 0$.

2. *Or, the probability that a younger man meets a woman decreases when cohort size rises, i.e.* $\frac{\partial\left(\lambda(N)\theta^i_{0,t}\right)}{\partial N} < 0$.

Then, the model can generate a negative relationship between changes in cohort size and marriage rates for both women and men.

Proof. See Appendix D.6. ■

The intuition behind this result is that both generalizations introduce additional negative effects of an increase in cohort size on marriage rates, relative to the standard model. The result for women is therefore obvious, as in the standard model their marriage rates are already decreasing with cohort size. For men, the supplemental negative effect must be sufficiently strong to dominate the positive effect generated by the standard model. In the congestion model, the additional negative effect is produced by a matching function that reduces the probability of meeting a potential spouse when cohort size increases. The surplus model generates the supplemental negative effect by reducing the gains from marriage, and hence the probability of marriage, when cohort size grows. If the additional negative effects are sufficiently large, the two augmented models can rationalize both fact 1 and fact 2.

The Congestion- vs. Surplus-Model: a Testable Implication. Since both variations of the matching model with search frictions can explain facts 1 and 2, it is important to derive a testable implication to separate them. To do this, we introduce in the model an additional feature of the marriage market: out-of-wedlock births. When a woman and man meet, they give birth to a child during the dating period with a probability σ. We assume that σ does not depend on cohort size to avoid biasing the framework in favor of the surplus or congestion model. If the couple chooses not to marry after giving birth to the child, we consider this case an out-of-wedlock birth. Note that, since σ is not affected by cohort size, Propositions 2-4 hold conditional on having and not having an out-of-wedlock birth. They therefore hold also unconditionally.

The following Proposition establishes that the two models have opposite predictions about the effect of a change in cohort size on the probability that an out-of-wedlock birth occurs. Since in the data we observe out-of-wedlock births only for women, the Proposition considers only women.

Proposition 5 *In the surplus model, a rise in cohort size increases the probability that a woman has an out-of-wedlock birth. In the congestion model, an increase in cohort size reduces this probability.*

Proof. See Appendix D.7. ■

The different implications of the surplus and congestion models for out-of-wedlock births are related to the distinct ways the two models generate the additional reduction in marriage rates when cohort size increases. In the surplus model, gains from marriage decrease when cohort size increases. Therefore, conditional on meeting, the probability that two individuals marry declines. Moreover, in this model, growing cohort size increases the probability that a woman and a man meet. These two features taken together imply an increase in the probability that a woman experiences an out-of-wedlock birth when cohort size rises. By contrast, the congestion model reduces the probability that a woman meets a man when cohort size increases. This feature has the additional effect of reducing the reservation value that determines whether a younger man marries a woman, without changing the reservation value for an older man. Thus, when cohort size increases, in the congestion model women are less likely to meet a man and, when they meet him, they are more likely to marry, generating a decline in out-of-wedlock births.

The Congestion vs. Surplus Model: a Look at the Data. As data on out-of-wedlock births are limited, especially if one wishes to use variation across states, we rely on the decennial Census to test the implication of Proposition 5. Because the Census does not contain fertility and marital histories, we only observe an out-of-wedlock birth if a never-married woman has a child in the household at the time of the Census. We therefore consider women that are between the ages of 18 and 22 to maximize the probability of observing an out-of-wedlock birth.

To test the implication, we use the same approach that we employ in the cross-state analysis. We first construct 10-year differences in out-of-wedlock births and in cohort size. We then regress the first variable on the second controlling for cohort fixed effects and census-region/cohort fixed effects using both OLS and IV, where we instrument for cohort size in a state by its size at birth.[10] The results are reported in Table 5.

The OLS coefficient on cohort size is negative and significant at -0.047, appearing to reject the surplus model in favor of the congestion model. However, the OLS coefficient is potentially downward biased due to migration decisions. States with high migration inflows are likely to have high cohort size at marriageable age and better economic opportunities, which should produce fewer out-of-wedlock births. When we control for this endogeneity issue using cohort size at birth, we find that the downward bias in the OLS estimate is large.

[10]Since the Censuses are available every 10 years, we cannot use the introduction of the pill interacted with the contraception bans as an instrument with these data.

Indeed, the estimated coefficient becames positive in both IV specification, and statistically significant and large in size at 0.070 in our specification with cohort-region fixed effect. These results indicate that the congestion model is rejected in favor of the surplus model.

5 Conclusions

We document two empirical facts that any theory of the evolution of marriage rates must be able to explain. First, there is a there is a strong and quantitatively large negative relationship between changes in cohort size and marriage rates of women. Second, the same relationship applies to men. We show that a standard dynamic matching model of the marriage market with search frictions can explain the first fact, but it fails to explain the second fact. A generalized matching model can rationalize both facts if cohort size affects either the efficiency of the matching technology or the distribution of marital surplus. We test these two models based on their opposite predictions about the relationship between changes in cohort size and the probability of an out-of-wedlock birth. Empirically, the fraction of out-of-wedlock births increases with cohort size, rejecting the congestion model in favor of the surplus model.

Our findings help shed light on the drivers of marital changes over the last century in the U.S, but they are also relevant in other settings. For example, large increases in births immediately after World War II affected growth in cohort size in virtually every developed country. This may explain at least partly why large declines in marriage rates throughout the 1970s – when these cohorts reached marriageable age – were observed in all of those countries. This is true even in Japan, which legalized oral contraceptives – a predominant alternative explanation behind falling marriage rates in the 1970s – only in the late 1990s. Finally, our findings help clarify why complementary explanations about changes in marriage rates account in many periods for a relatively limited share of overall variation in marriage rates, and in some periods appear irrelevant. For example, the significant increase in women's labor supply during and after World War II (Acemoglu et al. (2004)) and the improvements in home technologies in the 1940s and 1950s were accompanied by large increases in marriage rates, counter to the predictions of those theories. This does not necessarily indicate that these theories are wrong, but rather that other factors – in particular the large demographic changes affecting the marriage market in those decades – had a greater effect and pushed marriage rates upward.

Our findings also have policy implications. Both at the state and the federal level, politi-

cians and policy makers have discussed and implemented 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, such proposals allocated up to 1.5 billion dollars to implement and evaluate policies aimed at promoting marriage.[11] Our results suggest that these policies may prove 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.

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

References

Abramitzky, Ran, Adeline Delavande, and Luis Vasconcelos. 2011. "Marrying Up: The Role of Sex Ratio in Assortative Matching." *American Economic Journal: Applied Economics* 3 (3): pp. 124–157.

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

Akers, Donalds. 1967. "On Measuring the Marriage Squeeze." *Demography* 4 (2): 907–924.

Ananat, Elizabeth Oltmans, and Daniel M. Hungerman. 2012. "The Power of the Pill for the Next Generation: Oral Contraception's Effects on Fertility, Abortion, and Maternal and Child Characteristics." *The Review of Economics and Statistics* 94 (1): pp. 37–51.

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

Angrist, Joshua D., and Victor Lavy. 1996. "The Effect of Teen Childbearing and Single Parenthood on Childhood Disabilities and Progress in School." *NBER Working Paper*, no. 5807.

Bailey, Martha J. 2010. ""Momma's Got the Pill": How Anthony Comstock and Griswold v. Connecticut Shaped US Childbearing." *The American Economic Review* 100 (1): pp. 98–129.

———. 2013. "Fifty Years of Family Planning: New Evidence on the Long-Run Effects of Increasing Access to Contraception." *Brookings Papers on Economic Activity*, pp. 341–395.

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

Bergstrom, Theodore C., and David A. Lam. 1989. "The Effects of Cohort Size on Mariage Market in Twentieth Century Sweden." *Center for Research on Economic and Social Theory, Michigan* 91, no. 6.

Bitler, Marianne, and Lucie Schmidt. 2011. "Marriage Markets and Family Formation: The Role of the Vietnam Draft." *University of California at Irvine Working Paper.*

Black, Sandra E., and Paul J. Devereux. 2010. "Recent Developments in Intergenerational Mobility." *Prepared for the Handbook in Labor Economics 2010.*

Black, Sandra E, Paul G Devereux, and Kjell G Salvanes. 2005. "The More the Merrier? The Effect of Family Composition on Children's Education." *The Quarterly Journal of Economics* 120 (2): 669–700.

Bronson, Mary Ann, and Maurizio Mazzocco. 2017. "An Evaluation of Three Ways of Measuring Marriage Rates." *Journal of Demographic Economics, Forthcoming.*

Card, David, and Thomas Lemieux. 2001. "Can Falling Supply Explain the Rising Return to College for Younger Men? A Cohort-Based Analysis." *The Quarterly Journal of Economics* 116 (2): 705–746.

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

Cherlin, Andrew J. 1981. *Marriage, Divorce, Remarriage.* Cambridge: Harvard University Press.

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

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

Goldin, Claudia, and Lawrence F. Katz. 2002. "The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions." *Journal of Political Economy* 110 (4): pp. 730–770.

Gousse, Marion, Nicolas Jacquemet, and Jean Marc Robin. 2017. "Marriage, Labor Supply, and Home Production." *Econometrica* 85 (6): 1873–1919 (November).

Greenwood, Jeremy, and Nezih Guner. 2009. "Marriage and Divorce since World War II: Analyzing the Role of Technological Progress on the Formation of Households." *NBER Macroeconomics Annual* 23 (1): pp. 231–276.

Grossbard-Shechtman, Amyra. 1984. "A Theory of Allocation of Time in Markets for Labour and Marriage." *The Economic Journal* 94 (376): pp. 863–882.

Gruber, Jonathan. 2004. "Is Making Divorce Easier Bad for Children? The Long-Run Implications of Unilateral Divorce." *Journal of Labor Economics* 22 (4): pp. 799–833.

Hill, Matthew. 2011. "Love in the Time of Depression: The Effect of Economic Downturns on the Probability of Marriage." *UCLA Working Paper*.

Killingsworth, Mark R., and James J. Heckman. 1986. "Female Labor Supply: a Survey." *Handbook of Labor Economics, Edited by O. Ashenfelter and R. Layard,* 1:pp. 103–204.

King, Miriam, Steven Ruggles, J. Trent Alexander, Sarah Flood, Katie Genadek, Matthew B. Schroeder, Brandon Trampe, and Rebecca Vick. 2010. *Integrated Public Use Microdata Series, Current Population Survey: Version 3.0. [Machine-readable database]*. Minneapolis: University of Minnesota.

Knowles, John A., and Guillaume Vandenbroucke. 2016. "Fertility Shocks and Equilibrium Marriage-Rate Dynamics: Lessons from World War 1 in France." *Working Paper*.

McLanahan, Sara, and Christine Percheski. 2008. "Family Structure and the Reproduction of Inequalities." *The Annual Review of Sociology* 34:pp. 257–76.

Moffitt, Robert A. 2000. "Welfare Benefits and Female Headship in U.S. Time Series." *The American Economic Review* 90 (2): pp. 373–377.

Myers, Caitlin Knowles. 2017. "The Power of Abortion Policy: Reexamining the Effects of Young Women's Access to Reproductive Control." *Journal of Political Economy* 125 (6): 2178–2224.

Neal, Derek. 2004. "The Relationship between Marriage Market Prospects and Never-Married Motherhood." *The Journal of Human Resources* 39 (4): pp. 938–957.

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

Petrongolo, Barbara, and Christopher A Pissarides. 2001. "Looking into the Black Box: A Survey of the Matching Function." *Journal of Economic literature* 39 (2): 390–431.

Ruggles, Steven, J. Trent Alexander, Katie Genadek, Ronald Goeken, Matthew B. Schroeder, and Matthew Sobek. 2010. *Integrated Public Use Microdata Series: Version 5.0 [Machine-readable database]*. Minneapolis: University of Minnesota.

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

Seefeld, Kristin S., and Pamela J. Smock. 2004. "Marriage on the Public Policy Agenda: What Do Policy Makers Need to Know from Research?" *PSC Research Report, University of Michigan*, no. 04-554.

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

Siow, Aloysius. 1998. "Differential Fecundity, Markets, and Gender Roles." *Journal of Political Economy* 106 (2): 334–354.

Stevenson, Betsey, and Justin Wolfers. 2007. "Marriage and Divorce: Changes and Their Driving Forces." *The Journal of Economic Perspectives* 21 (2): pp. 27–52.

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

Tables and Figures

Table 1: Time Series Regression of Log Share Ever Married on Log Cohort Size

	Levels	Short Term	Medium Term	Long Term
Women by age 30	-0.193***	-0.071**	-0.100***	-0.141***
	(0.026)	(0.031)	(0.024)	(0.017)
Women by age 40	-0.066***	-0.008	-0.022*	-0.065***
	(0.007)	(0.021)	(0.013)	(0.008)
Men by age 30	-0.294***	0.033	-0.203***	-0.258***
	(0.037)	(0.050)	(0.052)	(0.044)
Men by age 40	-0.107***	0.050	-0.040*	-0.069**
	(0.008)	(0.027)	(0.023)	(0.022)

*** Significant at 1%. ** Significant at 5%. * Significant at 10%. Newey-West standard errors in parentheses. The short-term coefficients are the average coefficients for the one-year and two-year difference regressions. The medium-term coefficients are the average coefficients for the three-year to five-year difference regressions. The long-term coefficients are the average coefficients for the seven-year and ten-year difference regressions. Regressions include cohorts born after 1914 until the most recent cohort observed at a given age in 2015. The number of observations in each regression is equal to 71 for the share ever married by 30, 66 for the share ever married by 35, and 61 for the share ever married by 40. To account for the effect of adjacent cohorts in the cumulative regressions, for $n > 1$ we use differences in cumulative cohort size as our independent variable, where cumulative size for the cohort born in period t for the n-year difference is constructed by adding up cohort size from $t-n+1$ to t. Sources: IPUMS CPS 1962-2015, IPUMS USA 1960-1980.

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

Second Stage: Dependent Variable: 10-Yr. Difference in Log Share Ever Married by 30 or 40			
10-Yr. Dfference in Log Cohort Size	OLS	IV (1)	IV (2)
Men By Age 30	-0.041**	-0.123***	-0.168***
	(0.019)	(0.041)	(0.056)
R^2	0.807	0.791	0.819
Men By Age 40	-0.029***	-0.047**	-0.040
	(0.010)	(0.018)	(0.028)
R^2	0.555	0.551	0.595
Women By Age 30	-0.041***	-0.104***	-0.115***
	(0.010)	(0.023)	(0.036)
R^2	0.555	0.765	0.796
Women By Age 40	-0.032***	-0.058***	-0.048**
	(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

*** Significant at 1%. ** Significant at 5%. * Significant at 10%. Robust standard errors in parentheses. 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). The cross-state regressions (1)-(3) include all decennial cohorts born between 1910 and 1970, in all states except HI and AK, with a total of 288 observations. Sources: IPUMS USA, 1940-2100, NIH SEER Population Counts, US Population Counts, 1910-1990.

Table 3: Comstock Laws, the Pill, and Log Share Ever Married

Second Stage Men: Dependent Variable: N-Yr. Difference in Log Share Ever Married

	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

Second Stage Women: Dependent Variable: N-Yr. Difference in Log Share Ever Married

	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)	(0.084)	(0.076)
Obs.	1248	1152	1056	960

First Stage: 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)
Obs.	1248	1152	1056	960

* Significant at 10%. ** 5%. *** 1%. Clustered 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. Sources: IPUMS USA, 1980-2000, NIH SEER Population Counts.

Table 4: Comstock Laws and 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

Regressions are weighted by population and include controls for physician exception, physician exception interacted with "pill," advertising bans, and cohort-region fixed effects. Clustered standard errors in parentheses. Sources: IPUMS USA, 1980-2000, NIH SEER Population Counts.

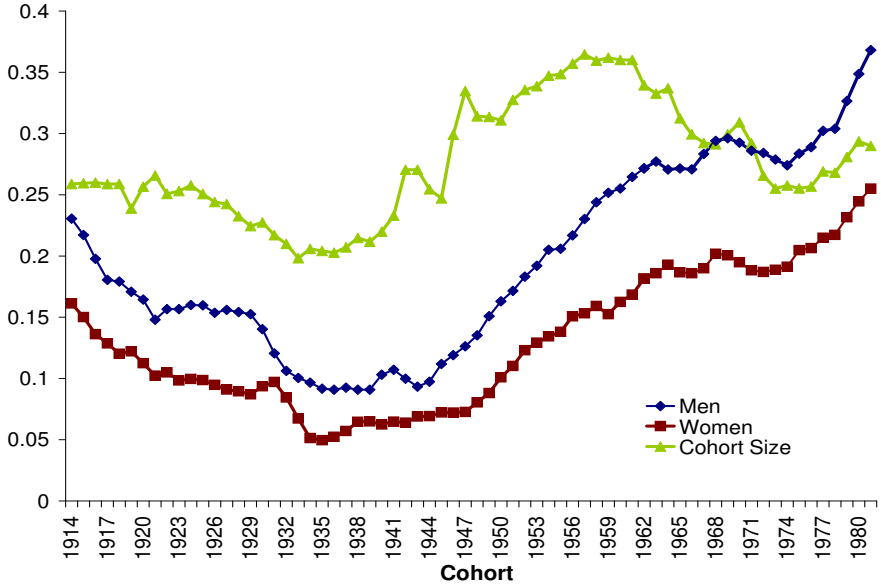
Table 5: Cross-Sectional Regression of Out of Wedlock Births

Second Stage: Dependent Variable: 10-Yr. Difference in Out of Wedlock Births, Women bet. 18 and 22			
10-Yr. Dfference in	OLS	IV	IV
Out of Wedlock Births		(1)	(2)
	-0.047***	0.042	0.070*
	(0.016)	(0.028)	(0.040)
P-values	0.004	0.135	0.079
First Stage Results			
Log Cohort Size at Birth			0.440*** (0.073)
F-test			36.44

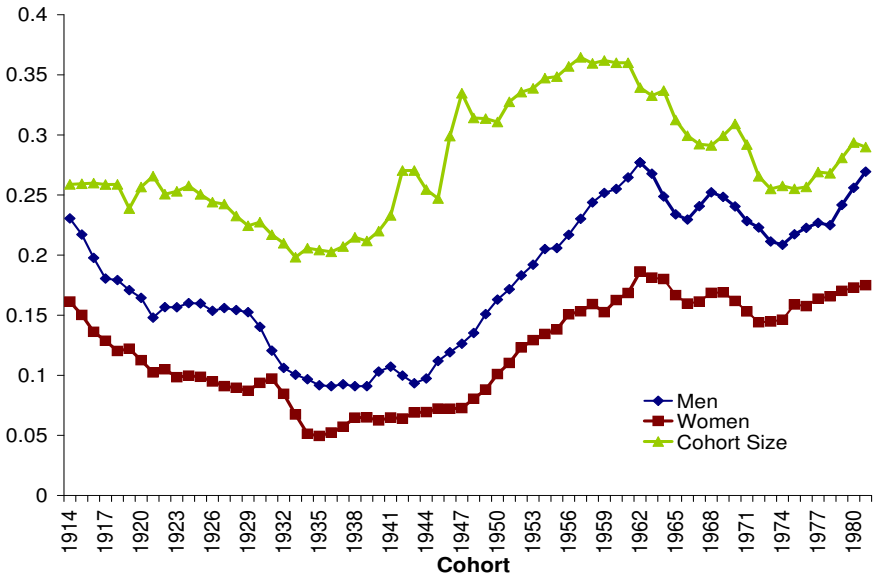
*** Significant at 1%. ** Significant at 5%. * Significant at 10%. Robust standard errors in parentheses. Each coefficient is the outcome of a separate, population-weighted regression. We control for cohort fixed effects, and for cohort-region fixed effects in IV-(2). The cross-state regressions (1)-(3) include all decennial cohorts born between 1910 and 1970, in all states except HI and AK, with a total of 288 observations. Sources: IPUMS USA, 1940-2100, NIH SEER Population Counts, US Population Counts, 1910-1990.

Figure 1: Marriage Rates by 30 and Cohort Size

(a) Married Women and Men



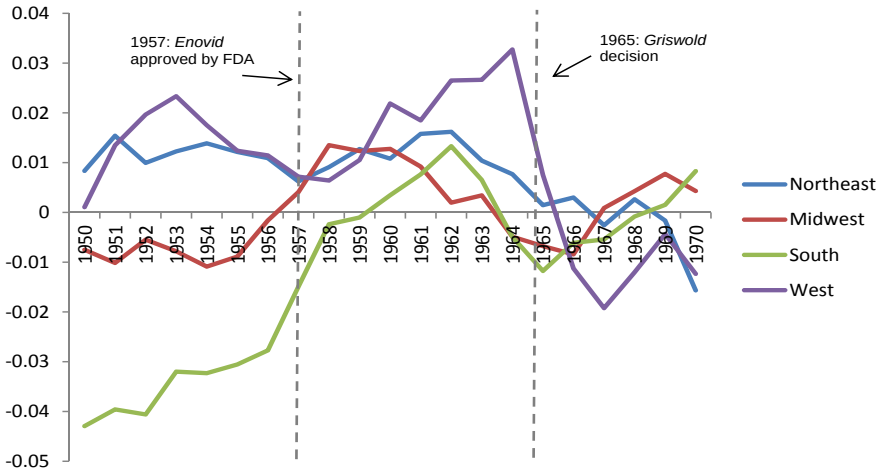
(b) Married and Cohabiting Women and Men



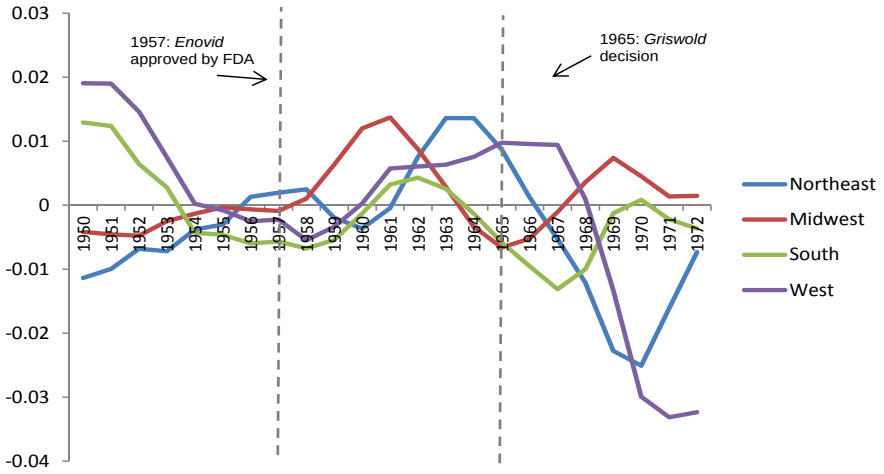
Notes: The vertical axis represents both the percentage of individuals ever married as well as normalized cohort size. We normalize cohort size by dividing by 10,000,000. Sources: Vital Statistics of the United States; IPUMS CPS, 1962-2015; IPUMS USA, 1960-1970.

Figure 2: Growth by Region: States with Sales Bans - States without Sales Bans

(a) Growth of Total Births



(b) Growth of Total Adult Population, at Age 25



* See note in Figure 1.

A Appendix: Data Description

Table A.1 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 share ever married, cohort size at birth, and cohort size at marriageable age.

The variable share ever married 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 1962-2015, 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 variation, we can construct the share ever married annually only for cohorts born after 1914 because the 1960 Census is the first one that records the age at first marriage.

As discussed in Section 3, cohabitation has become a close substitute for marriage in recent decades. Therefore, we use a measure of household formation that accounts for this rise in cohabitation. In particular, whenever we use the shorthand "ever married" in the empirical analysis, we refer to the share ever married or cohabiting by a given age, unless specifically noted otherwise. To identify cohabiting couples, 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. Additionally, none of our cross-state and IV regression results are qualitatively affected by including or excluding cohabitation.

For cohort size, we use two different measures in the paper, as discussed in Section 2: cohort size at marriageable age and cohort size at birth. 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 would be able to construct cohort size at marriageable 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 use the decennial Census for all years in the cross-state analysis. The decennial Census population counts are published for 5-year age groups. We therefore construct cohort size at marriage age by recording the number of individuals between the ages of 20 and 24 in the decennial Censuses 1930-1990. We also experimented with the 5-year age group 30-34 in the 1940-2000 Censuses 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.

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

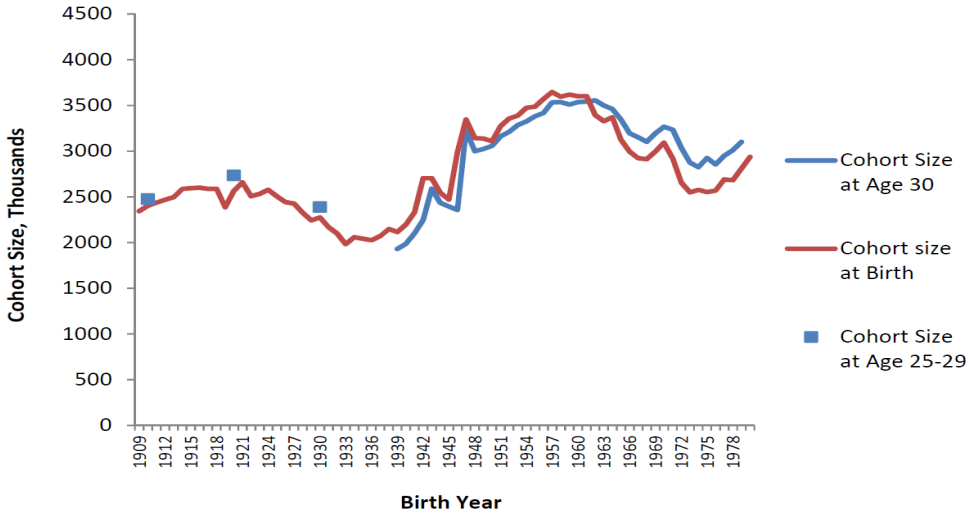
of the pill in states with different anti-obscenity laws. With longitudinal variation we use cohort size at birth as the main independent variable instead of cohort size at marriageable age for two reasons. First, as discussed above, the variable cohort size at marriageable age can be constructed only for decennial cohorts across the whole period of interest. For non-decennial cohorts, cohort size at marriageable age only be constructed for cohorts born after 1945, using SEER data. However, as indicated in Table A.1, cohort size at birth is available annually in the U.S. Vital Statistics by race for every cohort starting in 1909. Second, as shown in Figure A.1, when cohort size is computed for the U.S. population there is little difference between cohort size at birth and cohort size in adulthood, since net migration to the U.S. was relatively limited over this time period, as a share of the population of each cohort. By using cohort size at birth we can therefore consider a larger number of cohorts without significant effect on the analysis.

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 additionally at the state-level from 1940; and the decennial Censuses which provide information on population counts by race and state from the beginning of the twentieth century to 2010. For consistency, rather than combining two different datasets, we use the Censuses over the entire period of interest. Naturally, our results are very similar if we instead use data from the U.S. Vital Statistics for the cohorts born after 1940. As mentioned earlier, decennial Censuses publish population counts for 5-year age groups. From each decennial Census, we therefore record the number of individuals between the ages of 0 and 4 and use it to construct the cohort size at birth. 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 by state, race, and year to compute cohort size at birth for all of them.

Table A.1: 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 (1962-2015), U.S. Decennial Census, 1960-1970
	State Level, Decennial Years	U.S. Decennial Census, 1940-2000

Figure A.1: Cohort Size at Birth and at Age 30



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

B Existing Explanations

In this section, provide additional detail about findings in the existing literature on changes in cohort size and their effect on marriage behaviors, and how our findings compare to this literature.

The idea that changes in population growth introduce variation in the sex composition of the marriage market dates back to the late 1960s. As baby boom cohorts entered marriageable age, demographers argued that women in these cohorts, and in any rapidly growing population, would experience a "marriage squeeze." The classic studies on this topic are by Akers (1967) and Schoen (1983). They analyze the effect of cohort size on marriage rates using demographic models of the marriage market and data simulated from their models. In these models, changes in cohort size modify the supply of women relative to men in the marriage market, which in turn changes the marriage rate. The results of these papers suggest that an increase in cohort size should reduce the marriage rate of women and increase the marriage rate of men. Since those findings are the outcome of the developed models and of the corresponding assumptions, they should be interpreted as theoretical insights on the possible effect of changes in cohort size on marriage decisions. Our paper tests this idea formally using nearly a century of longitudinal data as well as cross-state variation. Our findings indicate that those models are rejected by the U.S. data since increases in cohort size reduce the marriage rates for both men and women. Our results also indicate that the effects of changes in cohort size on the eventual share ever married are substantially larger than those predicted by those models.

A second well-known paper studying the effects of cohort size on sex ratios and marriage behaviors is by Bergstrom and Lam (1989). The authors observe that predictions about the effects of changes in cohort size on sex ratios may overstate the effects on marriage rates, since age differences between spouses can adjust. Even if individuals prefer older or younger spouses, they can always marry within their own cohort, which generally is composed of equal numbers of men and women. To illustrate this point, they calibrate a marriage matching model following Becker (1973) in which marriage rates across cohorts are held constant, but age differences between spouses are allowed to vary. They provide evidence using Swedish data that average age differences between spouses in Sweden do in fact adjust in response to changes in cohort size, in line with the predictions of their model. However, their model assumes that marriage rates never change in response to increases or decreases in cohort size. As we document, this assumption is empirically rejected. Our findings using U.S. data indicate that age differences adjust in response to changes in cohort size, as one might expect, but marriage rates are also strongly affected.

An extended literature related to the previous two papers makes use of changes in the sex ratio, computed as the number of men divided by the number of women in the marriage market, to understand changes in marriage rates. Almost all papers in this literature rely either explicitly or implicitly on a two-sided matching model of the marriage market, as formalized by Becker (1973). A testable implication of this model is that a reduction in sex ratio should reduce women's and increase men's marriage rates. Several papers have attempted to test the relationship between sex ratios and marriage rates. Angrist (2002), Seitz (2009), Abramitzky, Delavande, and Vasconcelos (2011), and Knowles and Vandenbroucke (2016) are examples of papers in this literature.[12]

In our paper, we do not explicitly analyze the link between sex ratios and marriage rates. However, under the standard and empirically based assumption that women primarily marry men of similar or older ages, an increase in cohort size will reduce the sex ratio, or the number of men relative to women in the marriage market. Therefore, an increase in cohort size should reduce the marriage rate of women, but increase the marriage rate of men. An important finding of our paper is that this prediction of the standard matching model is rejected if one considers U.S. data on cohort size and marriage rates over the past century.

This finding does not imply that a change in sex ratio can never generate a marriage

[12]Several papers have studied the relationship between sex ratios and other economic variables, including rates of single motherhood (Neal (2004)), labor force participation of women (Grossbard-Shechtman (1984)), and birth rates (Bitler and Schmidt (2011)).

market outcome that is consistent with the matching model. For instance, Abramitzky, Delavande, and Vasconcelos (2011) consider variation in the sex ratio due to World War I casualties in France and find that a smaller sex ratio is associated with a lower marriage rate for women and a larger marriage rate for men, in line with Becker's prediction. Our findings simply indicate that, if the change in sex ratio is generated by a change in cohort size, the standard matching model is not consistent with U.S. marriage rate data from the past 100 years.

There is one important paper in the sex ratio literature whose results are consistent with ours. Angrist (2002) uses variation in immigration rates from different European countries to the U.S. at the beginning of the twentieth century to study the relationship between sex ratios and marriage rates. He exploits the fact that the majority of migrants were men and that marriages were often formed between individuals belonging to the same ethnicity. Consistent with the patterns documented here, he finds that ethnicities with lower sex ratios experienced lower marriage rates for women as well as men. In our paper, we document that the positive relationship between sex ratio and marriage rates of women and men applies to a century of data and not only to the specific period considered in Angrist (2002), if the changes in sex ratios are produced by variation in cohort size. Moreover, we provide an explanation for the apparent inconsistency between the findings in Angrist (2002) and the findings in Abramitzky, Delavande, and Vasconcelos (2011).

One alternative explanation for the link between cohort size and marriage rates that has commonality with the one we propose is the Easterlin hypothesis. Easterlin (1987) argues that the relative size of a cohort can explain many variables that determine the economic and social outcomes of that cohort: earnings and unemployment rates, college enrollment rates, divorce, fertility, crime, suicide rates, and marriage. Easterlin's theory is composed of two parts. First, the distribution of income of a cohort is affected by its size, with larger cohorts having worse economic outcomes. Second, when income of a cohort is above its aspiration level, the individuals in that cohort will be optimistic and therefore will have better economic and social outcomes. Researchers who have attempted to test the general idea behind Easterlin's hypothesis have found mixed results (Pampel and Peters (1995)).

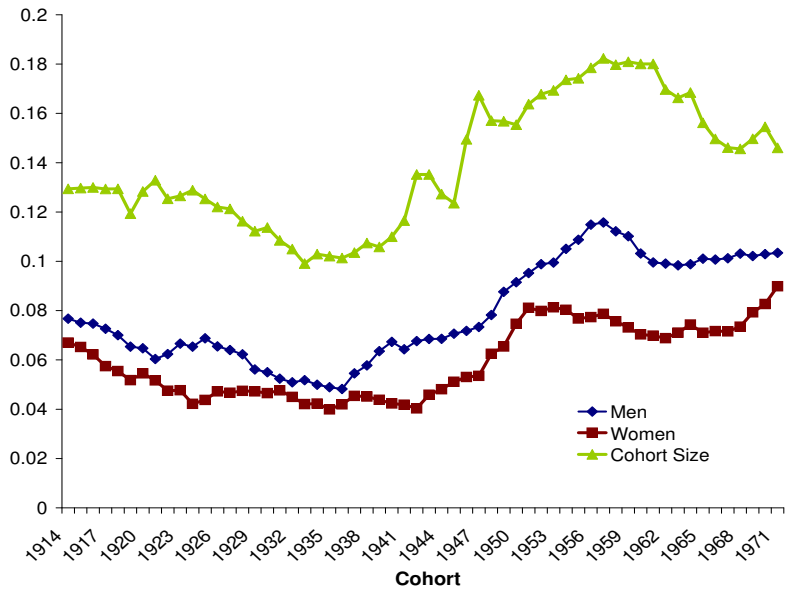
C Appendix: Supplement to Empirical Analysis

In this appendix, we provide supplemental evidence to accompany the analysis in Section 3.

C.1 Supplemental Tables and Figures: Longitudinal Analysis

Figure B.1 shows that changes in marriage rates by 40 similarly follow changes in cohort size. Table B.1 documents the results of the regressions of change in log share ever married on change in log cumulative cohort size, for changes over n years.

Figure B.1: Marriage Rates by 40 and Cohort Size: Married and Cohabiting



* See note in Figure 1.

Table B.1: 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%. ** 5%. *** 1%. See notes in Table 1. Newey-West standard errors in parentheses.

D Appendix: Proofs and Derivations

D.1 Reservation Values

We begin by characterizing the decisions of a man of age 1 in period t. If an old man chooses to be single in the second period, his lifetime utility takes the following form:

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

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

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

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 or until they divorce. Divorce can occur each period with a probability $q = 1 - p$. If a couple divorces, each individual receives the value of being single for the remainder of their lifetime. The lifetime utility of a couple of individuals who are both of age 0 and have drawn match quality $\hat{\eta}$ in period t can therefore be written as follows:

$$\begin{aligned} v_{0,0,t} &= \hat{\eta}\sum_{t=0}^{T}\beta^t p^t + 2\delta\sum_{t=1}^{T}\beta^t(1-p^t) = (\hat{\eta}-2\delta)\sum_{t=0}^{T}\beta^t p^t + 2\delta\sum_{t=0}^{T}\beta^t = \\ &= \frac{1-(p\beta)^{T+1}}{1-p\beta}(\hat{\eta}-2\delta) + \frac{1-\beta^{T+1}}{1-\beta}2\delta, \end{aligned}$$

where the last equality follows from the following geometric series formula:

$$\sum_{t=0}^{T} ab^t = a\frac{1-b^{T+1}}{1-b}.$$

If the couple is composed of an older man and a woman, the man will die one period earlier. As a consequence, following the same steps as in the derivation of $v_{0,0,t}$, their lifetime utility takes the following form:

$$v_{0,1,t} = \hat{\eta}\sum_{t=0}^{T-1}\beta^t p^t + 2\delta\sum_{t=1}^{T-1}\beta^t(1-p^t) + \beta^T\delta = \frac{1-(p\beta)^T}{1-p\beta}(\hat{\eta}-2\delta) + \frac{1-\beta^T}{1-\beta}2\delta + \beta^T\delta.$$

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,

$$\begin{aligned} w_{1,t}^{m}(\hat{\eta}) &= v_{1,t}^{m} + \gamma\left[v_{0,1,t} - v_{1,t}^{m} - v_{0,t}^{w}\right] \\ &= v_{1,t}^{m} + \gamma\left[\frac{1-(p\beta)^{T}}{1-p\beta}(\hat{\eta}-2\delta) + \frac{1-\beta^{T}}{1-\beta}2\delta + \beta^{T}\delta - v_{1,t}^{m} - v_{0,t}^{w}\right], \end{aligned}$$

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.

We can 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 marrying $v_{0,1,t}$ is greater than the sum of their lifetime utilities if they choose to stay single $v_{1,t}^{m} + v_{0,t}^{w}$. As a consequence, they will marry if and only if

$$\frac{1-(p\beta)^{T}}{1-p\beta}(\hat{\eta}-2\delta) + \frac{1-\beta^{T}}{1-\beta}2\delta + \delta\beta^{T} \geq \frac{1-\beta^{T}}{1-\beta}2\delta + \delta\beta^{T}.$$

This implies that the reservation value for match quality between a woman and a man of age 1 is

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

or, equivalently, the reservation value for marital surplus $\underline{\eta}_{1,t} = \underline{\hat{\eta}}_{1,t} - 2\delta = 0$. This result indicates that the decision of an older man depends on the drawn marital surplus and not separately on match quality and the value of being single.

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

$$w_{1,t}^{m}(\hat{\eta}) = \delta\frac{1-\beta^{T}}{1-\beta} + \gamma(\hat{\eta}-2\delta)\frac{1-(p\beta)^{T}}{1-p\beta}.$$

where γ is the man's share of marital surplus. As a consequence, the expected value function of an older man expressed in term of marital surplus $\eta = \hat{\eta} - 2\delta$ can be written in the following

form:

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

It is composed of three parts. The first term describes the value for the older man of meeting a woman with a marital surplus η sufficiently high that the couple will choose to marry multiplied by the corresponding probability. The second term characterizes the value of meeting a woman with a marital surplus η that is below the reservation value $\underline{\eta}_{1,t}$ times the probability of this event. Finally, the last term captures the value of not meeting a woman in the current period multiplied by the probability. By replacing $\underline{\eta}_{1,t} = 0$ and simplifying some of the terms, we obtain the following equation for the value function:

$$v_{1,t}^m = \delta\frac{1-\beta^T}{1-\beta} + E\left[\frac{1-(p\beta)^T}{1-p\beta}\gamma\eta\middle|\eta \geq 0\right]\left(1-F\left(0\right)\right)\theta_{1,t}^m. \tag{3}$$

We are now in position to consider the decision of a younger 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

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

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. With this expression, we can now solve for the reservation value of a man of age 0. Substituting for the expected value function of an older man using equation (3) and simplifying some of the terms, we obtain the following equation for the reservation value of a younger man:

$$\hat{\underline{\eta}}_{0,t} = 2\delta + \beta\frac{1-p\beta}{1-(p\beta)^{T+1}}\gamma E\left[\frac{1-(p\beta)^T}{1-p\beta}\eta'\middle|\eta' \geq 0\right]\left(1-F\left(0\right)\right)\theta_{1,t+1}^m,$$

or, equivalently, if we express the reservation threshold in terms of marital surplus,

$$\underline{\eta}_{0,t} = \beta\frac{1-p\beta}{1-(p\beta)^{T+1}}\gamma E\left[\frac{1-(p\beta)^T}{1-p\beta}\eta'\middle|\eta' \geq 0\right]\left(1-F\left(0\right)\right)\theta_{1,t+1}^m. \tag{4}$$

where η is the current realization of martial surplus and η' is one of the possible realizations if the younger man choose to stay single. Using $\underline{\eta}_{0,t}$, one can derive the expected value function for a woman and a younger man. They are presented in Appendix D.8. Analogously to the marriage decision of an older man, the marriage decision of a younger man depends on the drawn marital surplus and not separately on match quality and the value of being single, as long as $\theta^m_{1,t+1}$ depends only on marital surplus, which we will show is the case.

D.2 Proof of Proposition 1

In this subsection, we use the reservation values discussed above to solve for the steady state equilibrium in the marriage market. We provide the proof for the general case in which the cohort size of women and men is allowed to differ. First, we have to derive the probability that a younger man meets a woman $\theta^m_{0,t}$ and the corresponding probability for an older man $\theta^m_{1,t}$. Let $N^i_{a,t}$ be the number of individuals of gender i, age a, and period t who are present in the marriage market. Then $\theta^m_{0,t}$ and $\theta^m_{1,t}$ can be derived by noting that

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

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

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

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

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

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

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

We can now solve for N_1^m and obtain

$$N_1^m = \sqrt{\left(N_0^m\right)^2 - N_0^m N_0^w + N_0^m N_0^w F\left(\underline{\eta}_0\right)}.$$

Generally, men and women have identical cohort size, i.e. $N_{0,t}^m = N_{0,t}^w = N_{0,t}$.[13] In this case the solution for N_1^m simplifies to

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

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

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

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

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

To determine the reservation value of younger men in steady state, we can substitute for θ_1^m in the equation that determines the reservation value (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:

$$\underline{\eta}_{ss} = \beta \frac{1 - (p\beta)^T}{1 - (p\beta)^{T+1}} \gamma E\left[\eta \,|\eta \geq 0\right] \left(1 - F\left(0\right)\right) \frac{N_0^w}{N_0^m + \sqrt{\left(N_0^m\right)^2 - N_0^m N_0^w + N_0^m N_0^w F\left(\underline{\eta}_{ss}\right)}},$$

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

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

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

[13] 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.

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.

D.3 Derivation of the Effect of an Unexpected Shock to Cohort Size on the Reservation Value

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 $N_0 + \Delta$. According to equation (5), the probabilities $\theta_{j,t}^m$ take the following form:

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

and

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

Consider the period in which the shock is realized and notice that $N_{1,\tau}^m$ are the men born in period $\tau - 1$ who did not marry when younger. As a consequence, $N_{1,\tau}^m$ equals the number of older 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

$$\theta_{0,\tau}^m = \theta_{1,\tau}^m = \frac{N_0 + \Delta}{N_0 + \Delta + N_0 F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}} = \frac{1}{1 + \frac{N_0}{N_0 + \Delta} F\left(\underline{\eta}_{ss}\right)^{\frac{1}{2}}}. \tag{7}$$

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 younger 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 younger man.

We can now determine the effect of a shock to cohort size on the reservation marriage surplus of younger men $\underline{\eta}_{0,\tau}$. Notice that in the determination of $\underline{\eta}_{0,\tau}$ a younger 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 older men at $\tau+1$, which can be written as follows:

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

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

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

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

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

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

$$\underline{\eta}_{0,ss} = \beta\frac{1-(p\beta)^T}{1-(p\beta)^{T+1}}\gamma E\left[\eta\,|\eta\geq 0\right](1-F(0))\frac{1}{1+\left(1-\theta^m_{0,ss}\left(1-F\left(\underline{\eta}_{0,ss}\right)\right)\right)}. \tag{9}$$

Earlier in this section we have shown that, with a positive shock to cohort size, $\theta^m_{0,\tau} > \theta^m_{0,ss}$. As a consequence, a simple comparison of the last two equations implies that an increase in cohort size has the effect of increasing the reservation marriage surplus of younger men. Specifically, 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:

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

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

D.4 Proof of Proposition 2

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

probability she meets an older man times the probability she marries him, i.e.

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

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

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

D.5 Proof of Proposition 3

We prove the Proposition in two steps. We first prove that the probability that a man marries when younger in period t increases with cohort size. When then prove that a man marries when younger or older increases with cohort size.

First step. Let P^{ym}_t be the probability that a man marries when younger the period of the shock τ. Since we consider the case of a permanent shock to cohort size we have $N^m_{0,\tau} = N^m_{0,\tau+1}$. Using equation (6), we can write the number of older men in period $t+1$ as follows:

$$N^m_{1,\tau+1} = N^m_{0,\tau}\left(1 - \theta^m_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right)\right) = N^m_{0,\tau}\left(1 - P^{ym}_\tau\right). \tag{10}$$

Using the previous equation and equation (8), the reservation utility of a younger man can be written in the following form:

$$\underline{\eta}_{0,\tau} = B\frac{N_{0,\tau}}{N_{0,\tau} + N_{0,\tau}\left(1 - \theta^m_{0,\tau}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right)\right)} = B\frac{1}{1 + \left(1 - P^{ym}_\tau\right)} = B\frac{1}{2 - P^{ym}_\tau}.$$

In Section D.3, we have established that $\underline{\eta}_{0,ss} < \underline{\eta}_{0,\tau}$. Hence,

$$\underline{\eta}_{0,ss} = B\frac{1}{2 - P_{ss}^{ym}} < B\frac{1}{2 - P_{\tau}^{ym}} = \underline{\eta}_{0,\tau}.$$

The inequality implies that $P_{\tau}^{ym} > P_{ss}^{ym}$. We can therefore conclude that an increase in cohort size increases the probability that a man marries when younger.

Second step. The probability that a man marries when younger or older P_{τ}^{m} can be written as the probability that a man married when younger in period τ plus the probability that the same man marries when older in period $\tau + 1$, i.e.

$$P_{\tau}^{m} = \theta_{0,\tau}^{m}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right) + \left(1 - \theta_{0,\tau}^{m}\left(1 - F\left(\underline{\eta}_{0,\tau}\right)\right)\right)\theta_{1,\tau+1}^{m}\left(1 - F\left(0\right)\right). \tag{11}$$

The first part of the right hand side is the probability that a younger man meets a woman and marries her in period τ, which we denoted with P_{τ}^{ym}. The second part is the probability that a younger man does not marry in period τ, $1 - P_{\tau}^{ym}$, meets a woman in period $\tau + 1$, and marries her. Using equation (10), the probability that an older men meets a woman can be written as follows:

$$\theta_{1,\tau+1}^{m} = \frac{N_{0,\tau+1}^{m}}{N_{0,\tau+1}^{m} + N_{1,\tau+1}^{m}} = \frac{1}{2 - P_{\tau}^{ym}}.$$

As a consequence, equation (11) can be written as follows:

$$P_{\tau}^{m} = P_{\tau}^{ym} + \frac{1 - P_{\tau}^{ym}}{2 - P_{\tau}^{ym}}\left(1 - F\left(0\right)\right).$$

Taking the derivative with respect to cohort size N of both sides and rearranging terms, we have,

$$\frac{\partial P_{\tau}^{m}}{\partial N} = \frac{\partial P_{\tau}^{ym}}{\partial N}\left[1 - \frac{1 - F\left(0\right)}{\left(2 - P_{\tau}^{ym}\right)^{2}}\right] > \frac{\partial P_{\tau}^{ym}}{\partial N}F\left(0\right) > 0,$$

where the first inequality follows from $\left(2 - P_{\tau}^{ym}\right)^{2} > 1$ and the second from the first step of the proof. Hence, an increase in cohort size increases the probability that a man marries.

D.6 Proof of Proposition 4

Since, we only have to prove that the two models *can* generate the decline in marriage rates for men, we will prove the Proposition under the assumption that match quality $\hat{\eta}$ is uniformly distributed in the interval $[0, 1]$. After the proof for each model, we provide an

example to illustrate the result.

Proof for the surplus model. Suppose that an increase in cohort size produces a deterioration of the marital status distribution by shifting it to the left by $\phi(N) > 0$, with $\frac{\partial \phi(N)}{\partial N} > 0$. Let $F^{\phi}(x)$ be the probability distribution of the marital surplus variable $\eta^{\phi} = \eta - \phi(N)$, i.e.

$$F^{\phi}(x) = P(\eta - \phi(N) \leq x) = P(\eta \leq x + \phi(N)) = F(x + \phi(N)).$$

The result is obvious for women, since $\phi(N) > 0$ reduces their marriage rate even more when cohort size increases. We therefore provide the proof only for men. In the model in which marital surplus deteriorates with cohort size, the reservation value for older men for the variable η^{ϕ} is equal to

$$\underline{\eta}^{\phi}_{1,t} = \underline{\eta}_{1,t} - \phi(N) = 0,$$

where $\underline{\eta}_{1,t}$ is the reservation value in the new model for the variable η for older men. For younger men, the reservation value for the variable η^{ϕ} is equal to

$$\begin{aligned}
\underline{\eta}^{\phi}_{0,t} = \underline{\eta}_{0,t} - \phi(N) &= \beta \frac{1-(p\beta)^T}{1-(p\beta)^{T+1}} \gamma E^{\phi}\left[\eta^{\phi} \middle| \eta^{\phi} \geq 0\right]\left(1 - F^{\phi}(0)\right) \\
&\quad \times \frac{1}{1 + \left(1 - \theta^m_{0,t}\left(1 - F^{\phi}\left(\underline{\eta}^{\phi}_{0,t}\right)\right)\right)} \\
&= \beta \frac{1-(p\beta)^T}{1-(p\beta)^{T+1}} \gamma E\left[\eta - \phi(N) \,|\, \eta \geq \phi(N)\right](1 - F(\phi(N))) \\
&\quad \times \frac{1}{1 + \left(1 - \theta^m_{0,t}\left(1 - F\left(\underline{\eta}^{\phi}_{0,t} + \phi(N)\right)\right)\right)} \\
&= B(\phi(N)) \frac{1}{1 + \left(1 - \theta^m_{0,t}\left(1 - F\left(\underline{\eta}^{\phi}_{0,t} + \phi(N)\right)\right)\right)}. \qquad (12)
\end{aligned}$$

Following the same argument used in Proposition 3 to derive the probability that a cohort-t man marries in the standard model, the probability that a cohort-t man marries when younger or older in the surplus model can be written as follows:

$$P^m_t = P^{ym}_t + \frac{1 - P^{ym}_t}{2 - P^{ym}_t}\left(1 - F^{\phi}(0)\right) = P^{ym}_t + \frac{1 - P^{ym}_t}{2 - P^{ym}_t}(1 - F(\phi(N))).$$

Taking the derivative with respect to N, we obtain

$$\frac{\partial P_t^m}{\partial N} = \frac{\partial P_t^{ym}}{\partial N}\left[1 - \frac{1 - F\left(\phi(N)\right)}{\left(2 - P_t^{ym}\right)^2}\right] - \frac{1 - P_t^{ym}}{2 - P_t^{ym}} F'\left(\phi(N)\right)\frac{\partial \phi(N)}{\partial N}.$$

Since the derivative of the cdf $F'\left(\phi(N)\right)$ is equal to the pdf $f\left(\phi(N)\right) > 0$ and $\dfrac{\partial \phi(N)}{\partial N} > 0$,

$$-\frac{1 - P_t^{ym}}{2 - P_t^{ym}} F'\left(\phi(N)\right)\frac{\partial \phi(N)}{\partial N} < 0.$$

Moreover,

$$0 \leq \left[1 - \frac{1 - F\left(\phi(N)\right)}{\left(2 - P_t^{ym}\right)^2}\right] \leq 1,$$

as $0 \leq 1 - F\left(\phi(N)\right) \leq 1$ and $\left(2 - P_t^{ym}\right)^2 > 1$. Hence, if we can show that, with the introduction of $\phi(N) > 0$, $\dfrac{\partial P_t^{ym}}{\partial N}$ can be less than or equal to zero, we have the result.

P_τ^{ym} can be written as follows:

$$P_\tau^{ym} = \theta_{0,\tau}^m\left(1 - F\left(\underline{\eta}_{0,\tau}^\phi + \phi(N)\right)\right).$$

Under the assumption that match quality $\hat{\eta}$ is uniformly distributed in the interval $[0, 1]$, we have that $\eta = \hat{\eta} - 2\delta$ is uniformly distributed in the interval $[-2\delta, 1 - 2\delta]$. Hence, $F\left(\underline{\eta}_{0,\tau}^\phi + \phi(N)\right) = \underline{\eta}_{0,\tau}^\phi + \phi(N) + 2\delta$ and the previous equation can be rewritten as follows:

$$P_\tau^{ym} = \theta_{0,\tau}^m\left(1 - \underline{\eta}_{0,\tau}^\phi - \phi(N) - 2\delta\right).$$

Moreover, using equation (12) the reservation value of younger men takes the following form:

$$\underline{\eta}_{0,\tau}^\phi = B\left(\phi(N)\right)\frac{1}{1 + \left(1 - \theta_{0,\tau}^m\left(1 - \underline{\eta}_{0,\tau}^\phi - \phi(N) - 2\delta\right)\right)} = B\left(\phi(N)\right)\frac{1}{2 - P_\tau^{ym}}.$$

We can therefore write P_τ^{ym} as

$$P_\tau^{ym} = \theta_{0,\tau}^m\left(1 - \frac{B\left(\phi(N)\right)}{2 - P_\tau^{ym}} - \phi(N) - 2\delta\right),$$

or equivalently

$$-\left(P_\tau^{ym}\right)^2 + \left(2 + \theta_{0,\tau}^m\left(1 - \phi(N) - 2\delta\right)\right)P_\tau^{ym} - \theta_{0,\tau}^m\left(2 - B\left(\phi(N)\right) - 2\left(\phi(N) + 2\delta\right)\right) = 0.$$

By taking the derivative with respect to N and rearranging terms we have,

$$\frac{\partial P_\tau^{ym}}{\partial N}\left[2\left(1-P_\tau^{ym}\right)+\theta_{0,\tau}^m\left(1-\phi(N)-2\delta\right)\right]=$$
$$=-\theta_{0,\tau}^m\left(2-P_\tau^{ym}-\frac{\partial B}{\partial\phi}\right)\frac{\partial\phi}{\partial N}+\left[\left(2-P_\tau^{ym}\right)\left(1-\phi(N)-2\delta\right)-B\left(\phi(N)\right)\right]\frac{\partial\theta_{0,\tau}^m}{\partial N}, \qquad (13)$$

Note that

$$\begin{aligned} B\left(\phi(N)\right) &= \beta\frac{1-(p\beta)^T}{1-(p\beta)^{T+1}}\gamma E\left[\eta-\phi(N)\,|\eta\geq\phi(N)\right]\left(1-F\left(\phi(N)\right)\right) \\ &= \frac{K}{2}\left(1-\phi(N)-2\delta\right)^2, \end{aligned}$$

where $\frac{K}{2}=\beta\frac{1-(p\beta)^T}{1-(p\beta)^{T+1}}\gamma$. Hence,

$$\frac{\partial B}{\partial\phi}=-K\left(1-\phi(N)-2\delta\right).$$

We can therefore write equation (13) as follows:

$$\begin{aligned} \frac{\partial P_\tau^{ym}}{\partial N}\left[2\left(1-P_\tau^{ym}\right)+\theta_{0,\tau}^m\left(1-\phi(N)-2\delta\right)\right] &= -\theta_{0,\tau}^m\left(2-P_\tau^{ym}+K\left(1-\phi(N)-2\delta\right)\right)\frac{\partial\phi}{\partial N} \\ &+\left[\left(2-P_\tau^{ym}\right)\left(1-\phi(N)-2\delta\right)-\frac{K}{2}\left(1-\phi(N)-2\delta\right)^2\right]\frac{\partial\theta_{0,\tau}^m}{\partial N} \\ &= -\theta_{0,\tau}^m\left(2-P_\tau^{ym}+K\left(1-\phi(N)-2\delta\right)\right)\frac{\partial\phi}{\partial N} \\ &+\left[2-P_\tau^{ym}-\frac{K}{2}\left(1-\phi(N)-2\delta\right)\right]\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N} \\ &= -\left(2-P_\tau^{ym}\right)\left[\theta_{0,\tau}^m\frac{\partial\phi}{\partial N}-\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}\right]-K\left(1-\phi(N)-2\delta\right)\left[\theta_{0,\tau}^m\frac{\partial\phi}{\partial N}+\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}\right]. \end{aligned} \qquad (14)$$

Since $2\left(1-P_\tau^{ym}\right)+\theta_{0,\tau}^m\left(1-\phi(N)-2\delta\right)\geq 0$, the sign of $\frac{\partial P_\tau^{ym}}{\partial N}$ corresponds to the sign of the right hand side of the previous equation. Note that $2-P_\tau^{ym}>0$ and $K\left(1-\phi(N)-2\delta\right)\geq 0$ and $\left[\theta_{0,\tau}^m\frac{\partial\phi}{\partial N}+\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}\right]>0$. Hence, as $\frac{\partial P_\tau^{ym}}{\partial N}\leq 0$ if

$$\theta_{0,\tau}^m\frac{\partial\phi}{\partial N}\geq\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}.$$

Since,

$$\left(1-\phi(N)-2\delta\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}=\left(1-F\left(\phi(N)\right)\right)\frac{\partial\theta_{0,\tau}^m}{\partial N}\leq\frac{\partial\theta_{0,\tau}^m}{\partial N},$$

a sufficient condition for $\frac{\partial P_\tau^{ym}}{\partial N} \leq 0$, and hence $\frac{\partial P_\tau^{m}}{\partial N} \leq 0$, is that

$$\frac{\partial \phi}{\partial N} \geq \frac{1}{\theta_{0,\tau}^m} \frac{\partial \theta_{0,\tau}^m}{\partial N} = \frac{\partial \ln \theta_{0,\tau}^m}{\partial N}.$$

This condition is satified if one sets for instance $\phi(N) = \ln \theta_{0,\tau}^m + H$, where H is a positive constant choosen to guarantee that $\phi(N) > 0$.

Using this results, we can find an example in which $\frac{\partial P_\tau^{ym}}{\partial N} \leq 0$ and, hence, $\frac{\partial P_\tau^{m}}{\partial N} < 0$. Let cohort size in steady state be $N = 100$ and the increase at time τ be $\Delta N = 1$. Moreover, let $\delta = 0.25$, $\beta = 1$, $p = 1$, and $\gamma = 0.5$. We set $\phi(N) = \ln \theta_{0,t}^m + H$, with $H = 1$. Under these conditions we have that, in steady state, $\theta_{0,ss}^m = 0.5664$, $\phi(N) = 0.3467$, $\underline{\eta}_{0,ss} = 0.0031$, and the number of older individuals is $N_{1,ss} = 92.1833$. The probability that a younger man marries in steady state is therefore $P_{ss}^{ym} = \theta_{0,ss}^m \left(1 - \underline{\eta}_{0,ss} - \phi(N) - 2\delta\right) = 0.0782$ and the probability that a man in a given cohort marries is

$$P_{ss}^m = P_{ss}^{ym} + \frac{1 - P_{ss}^{ym}}{2 - P_{ss}^{ym}} \left(1 - \phi(N) - 2\delta\right) = 0.1517.$$

At τ, the period of the cohort shock, we have $\theta_{0,\tau}^m = \frac{N+\Delta}{N+\Delta+N_{1,\tau}} = \frac{N+\Delta}{N+\Delta+N_{1,ss}} = 0.5228$, which implies that $\phi(N + \Delta) = \ln \theta_{0,\tau}^m + 1 = 0.3515$. The reservation value is the solution of equation (12), which is a quadratic equation in $\underline{\eta}_{0,\tau}$ and has as the only positive solution 0.0029. The probability that a younger man marries at τ is therefore $P_\tau^{ym} = \theta_{0,\tau}^m \left(1 - \underline{\eta}_{0,\tau} - \phi(N + \Delta) - 2\delta\right) = 0.762$ and the probability that a cohort-τ man marries is

$$P_\tau^m = P_\tau^{ym} + \frac{1 - P_\tau^{ym}}{2 - P_\tau^{ym}} \left(1 - \phi(N + \Delta) - 2\delta\right) = 0.1475,$$

which is lower than the corresponding probability before the shock.

Proof for the Congestion Model. The only difference between the standard model and the model with congestion is that the probability that a younger man meets a women become $\lambda(N)\,\theta_{0,t}^i$ and the corresponding probability for an older man take the form $\lambda(N)\,\theta_{1,t}^i$, with $\lambda(N) > 0$ and $\frac{\partial \lambda \theta_{0,t}^i(N)}{\partial N} \leq 0$. Analogously to the proof for the surplus model, this straightforwardly implies that an increase in cohort size reduced the marriage rate of women, as the congestion effects decrease the marriage rate of women more than in the standard model. We will therefore provide the proof only for men.

As mentioned earlier, we will consider the case in which match quality $\hat{\eta}$ is distributed

according to a uniform $[0,1]$ and, hence, marital surplus is distributed according to a uniform $[-2\delta, 1-2\delta]$. Following the steps used in Proposition 3, the probability that a cohort-t man marries a women can be written as

$$P_t^m = \lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right) + \left(1-\lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\right)\lambda(N)\,\theta_{1,t+1}^m\left(1-F(0)\right).$$

Since $P_t^{ym} = \lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)$ and we showed in Proposition 3 that $\theta_{1,t+1}^m = \dfrac{1}{2-P_t^{ym}}$, the probability P_t^m can be written as follows:

$$P_t^m = P_t^{ym} + \lambda(N)\frac{1-P_t^{ym}}{2-P_t^{ym}}\left(1-F(0)\right),$$

and its derivative with respect to cohort size N as

$$\frac{\partial P_t^m}{\partial N} = \frac{\partial P_t^{ym}}{\partial N}\left[1-\frac{\lambda(N)\left(1-F(0)\right)}{\left(2-P_t^{ym}\right)^2}\right] + \frac{1-P_t^{ym}}{2-P_t^{ym}}\left(1-F(0)\right)\frac{\partial\lambda(N)}{\partial N}. \tag{15}$$

The derivative $\dfrac{\partial P_t^{ym}}{\partial N} = \dfrac{\partial}{\partial N}\left(\lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\right)$ takes the following form:

$$\frac{\partial P_t^{ym}}{\partial N} = \frac{\partial\left(\lambda(N)\,\theta_{0,t}^m\right)}{\partial N}\left(1-F\left(\underline{\eta}_{0,t}\right)\right) - \frac{\partial F\left(\underline{\eta}_{0,t}\right)}{\partial N}\lambda(N)\,\theta_{0,t}^m,$$

which under the assumption that $\eta \sim U[-2\delta, 1-2\delta]$, becomes

$$\frac{\partial P_t^{ym}}{\partial N} = \frac{\partial\left(\lambda(N)\,\theta_{0,t}^m\right)}{\partial N}\left(1-F\left(\underline{\eta}_{0,t}\right)\right) - \frac{\partial\underline{\eta}_{0,t}}{\partial N}\lambda(N)\,\theta_{0,t}^m. \tag{16}$$

In the congestion model, the reservation value of a younger man takes the following form:

$$\begin{aligned}\underline{\eta}_{0,t} &= \beta\frac{1-(p\beta)^T}{1-(p\beta)^{T+1}}\gamma E\left[\eta\,|\eta\geq 0\right]\left(1-F(0)\right)\lambda(N)\frac{1}{1+\left(1-\lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\right)}\\ &= B\frac{\lambda(N)}{1+\left(1-\lambda(N)\,\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\right)}.\end{aligned} \tag{17}$$

Consider the derivative of both sides of the previous equation with respect to N. Since $\frac{1}{1+\left(1-\lambda(N)\theta_{0,t}^m\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\right)} = \theta_{1,t+1}^m = \frac{1}{2-P_t^{ym}}$ and η is distributed uniformly, we have,

$$\frac{\partial\underline{\eta}_{0,t}}{\partial N} = \frac{B}{2-P_t^{ym}}\frac{\partial\lambda(N)}{\partial N} - \frac{B\lambda(N)^2\,\theta_{0,t}^m}{\left(2-P_t^{ym}\right)^2}\frac{\partial\underline{\eta}_{0,t}}{\partial N} + \frac{B\lambda(N)\left(1-F\left(\underline{\eta}_{0,t}\right)\right)}{\left(2-P_t^{ym}\right)^2}\frac{\partial\left(\lambda(N)\,\theta_{0,t}^m\right)}{\partial N}.$$

Hence,

$$\frac{\partial \underline{\eta}_{0,t}}{\partial N} = \frac{B\left(2-P_t^{ym}\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\frac{\partial\lambda\left(N\right)}{\partial N} + \frac{B\lambda\left(N\right)\left(1-F\left(\underline{\eta}_{0,t}\right)\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\frac{\partial\left(\lambda\left(N\right)\theta_{0,t}^m\right)}{\partial N}.$$

By replacing for $\frac{\partial \underline{\eta}_{0,t}}{\partial N}$ in equation (16), we have

$$\begin{aligned}\frac{\partial P_t^{ym}}{\partial N} &= \frac{\partial\left(\lambda\left(N\right)\theta_{0,t}^m\right)}{\partial N}\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\left[1-\frac{B\lambda\left(N\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\right] \\ &- \frac{\lambda\left(N\right)\theta_{0,t}^m B\left(2-P_t^{ym}\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\frac{\partial\lambda\left(N\right)}{\partial N}. \end{aligned} \tag{18}$$

Notice that in the equation above

$$\frac{B\lambda\left(N\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m} \leq \frac{B\lambda\left(N\right)}{\left(2-P_t^{ym}\right)^2} \leq \frac{B\lambda\left(N\right)}{2-P_t^{ym}} = \underline{\eta}_{0,t} < 1.$$

Hence, since $\frac{\partial\left(\lambda\left(N\right)\theta_{0,t}^m\right)}{\partial N} \leq 0$, in equation (18), the term

$$\frac{\partial\left(\lambda\left(N\right)\theta_{0,t}^m\right)}{\partial N}\left(1-F\left(\underline{\eta}_{0,t}\right)\right)\left[1-\frac{B\lambda\left(N\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\right] \leq 0,$$

As a consequence, since in equation (15) $1-\frac{\lambda\left(N\right)\left(1-F\left(0\right)\right)}{\left(2-P_t^{ym}\right)^2} \geq 0$, if we can show that $\frac{\partial P_t^m}{\partial N} \leq 0$ for the extreme case where $\frac{\partial\lambda\theta_{0,t}^i(N)}{\partial N} = 0$, the same result applies to the general case. In this case, equation (18) becomes

$$\frac{\partial P_t^{ym}}{\partial N} = -\frac{\lambda\left(N\right)\theta_{0,t}^m B\left(2-P_t^{ym}\right)}{\left(2-P_t^{ym}\right)^2 + B\lambda\left(N\right)^2\theta_{0,t}^m}\frac{\partial\lambda\left(N\right)}{\partial N}.$$

Hence, using equation (15) and $F\left(0\right) = 2\delta$ due to the assumption that $\eta \sim U\left[-2\delta, 1-2\delta\right]$,

the derivative of the probability that a cohort-t man marries can be written as follow:

$$\begin{aligned}
\frac{\partial P_t^m}{\partial N} &= \left[-\left(1-\frac{\lambda\left(1-F\left(0\right)\right)}{\left(2-P_t^{ym}\right)^2}\right)\frac{\lambda\theta_{0,t}^m B\left(2-P_t^{ym}\right)}{\left(2-P_t^{ym}\right)^2+B\lambda^2\theta_{0,t}^m}+\frac{1-P_t^{ym}}{2-P_t^{ym}}\left(1-F\left(0\right)\right)\right]\frac{\partial\lambda}{\partial N} \\
&= \frac{\left[-\left(2-P_t^{ym}\right)^2+\lambda\left(1-2\delta\right)+\left(1-P_t^{ym}\right)\left(1-2\delta\right)\lambda\right]\lambda\theta_{0,t}^m B+\left(1-P_t^{ym}\right)\left(1-2\delta\right)\left(2-P_t^{ym}\right)^2}{\left(2-P_t^{ym}\right)\left[\left(2-P_t^{ym}\right)^2+B\lambda^2\theta_{0,t}^m\right]}\frac{\partial\lambda}{\partial N} \\
&= \frac{\left[-\left(2-P_t^{ym}\right)^2+\left(1-2\delta\right)\left(2-P_t^{ym}\right)\lambda\right]\lambda\theta_{0,t}^m B+\left(1-P_t^{ym}\right)\left(1-2\delta\right)\left(2-P_t^{ym}\right)^2}{\left(2-P_t^{ym}\right)\left[\left(2-P_t^{ym}\right)^2+B\lambda^2\theta_{0,t}^m\right]}\frac{\partial\lambda}{\partial N} \\
&= \frac{\left[-\left(2-P_t^{ym}\right)+\left(1-2\delta\right)\lambda\right]\lambda\theta_{0,t}^m B+\left(1-P_t^{ym}\right)\left(1-2\delta\right)\left(2-P_t^{ym}\right)}{\left(2-P_t^{ym}\right)^2+B\lambda^2\theta_{0,t}^m}\frac{\partial\lambda}{\partial N}.
\end{aligned}$$

As the cohort shock hits the economy when it is in steady state, the previous derivative has to be evaluated at the steady state. Consequently, it takes the following form:

$$\left.\frac{\partial P_t^m}{\partial N}\right|_{t=ss}=\frac{\left[-\left(2-P_{ss}^{ym}\right)+\left(1-2\delta\right)\lambda_{ss}\right]\lambda_{ss}\theta_{0,ss}^m B+\left(1-P_{ss}^{ym}\right)\left(1-2\delta\right)\left(2-P_{ss}^{ym}\right)}{\left(2-P_{ss}^{ym}\right)^2+B\lambda^2\theta_{0,ss}^m}\frac{\partial\lambda}{\partial N}. \tag{19}$$

Since $\frac{\partial\theta_{0,t}^i}{\partial N}>0$, the assumption that $\frac{\partial\lambda(N)\theta_{0,t}^i}{\partial N}=0$ implies that $\frac{\partial\lambda}{\partial N}<0$. Moreover, $P_t^{ym}\leq 1$, $B>0$, $\lambda>0$, and $\theta_{0,t}^m\geq 0$. Hence, $\left.\frac{\partial P_t^m}{\partial N}\right|_{t=ss}\leq 0$ if the numerator of equation (19) is positive. To prove that the numerator is positive note that in the case we consider $\frac{\partial\lambda(N_t)\theta_{0,t}^i}{\partial N}=0$. Hence, $\lambda\left(N_t\right)\theta_{0,t}^i=K$, for some constant $0\leq K\leq 1$. We can therefore rewrite the numerator of equation (19) as follows:

$$\begin{aligned}
d(K) &= \left[-\left(2-P_{ss}^{ym}\right)+\left(1-2\delta\right)\lambda\right]KB+\left(1-P_{ss}^{ym}\right)\left(1-2\delta\right)\left(2-P_{ss}^{ym}\right) \\
&\geq -\left(2-P_{ss}^{ym}\right)KB+\left(1-P_{ss}^{ym}\right)\left(1-2\delta\right)\left(2-P_{ss}^{ym}\right).
\end{aligned}$$

Hence, $d(K)>0$ if it is possible to fine a $0<K\leq 1$ such that

$$-\left(2-P_{ss}^{ym}\right)KB+\left(1-P_{ss}^{ym}\right)\left(1-2\delta\right)\left(2-P_{ss}^{ym}\right)>0,$$

or equivalently

$$-KB+\left(1-P_{ss}^{ym}\right)\left(1-2\delta\right)>0.$$

The probability that a younger man marries can be written as

$$P_{ss}^{ym}=\lambda\left(N\right)\theta_{0,t}^i\left(1-F\left(\underline{\eta}_{0,ss}\right)\right)=K\left(1-\underline{\eta}_{0,ss}-2\delta\right).$$

By replacing for P_{ss}^{ym} in the previous inequality, we have:

$$-KB + \left(1 - K\left(1 - \underline{\eta}_{0,ss} - 2\delta\right)\right)(1 - 2\delta) > 0.$$

In our model, $\theta_{0,t}^{m} = \theta_{1,t}^{m}$ in each t. Hence, in steady state, $\underline{\eta}_{0,ss} = B\lambda\theta_{1,ss}^{m} = B\lambda\theta_{0,ss}^{m} = BK$. We can therefore derive the set of values for K that make $d(K) > 0$ by replacing for $\underline{\eta}_{0,ss} = BK$ in the previous inequality and obtain

$$-KB + (1 - K(1 - BK - 2\delta))(1 - 2\delta) > 0,$$

which is equivalent to

$$(1 - 2\delta)BK^2 - \left(B + (1 - 2\delta)^2\right)K + (1 - 2\delta) > 0. \tag{20}$$

The two roots for K of the equation that correspond to the previous inequality are

$$K^{+} = \frac{1}{1 - 2\delta} \qquad \text{and} \qquad K^{-} = \frac{1 - 2\delta}{2B}.$$

Observe that

$$B = \beta\frac{1 - (p\beta)^{T}}{1 - (p\beta)^{T+1}}\gamma E\left[\eta \,|\eta \geq 0\right](1 - F(0)) \leq E\left[\eta \,|\eta \geq 0\right](1 - F(0)) = \frac{1 - 2\delta}{2}(1 - 2\delta) = \frac{(1 - 2\delta)^2}{2}.$$

Thus, $K^{+} \leq K^{-}$. Since the second derivative of the left hand side of the inequality is positive, this implies that all the values of K in the set $\{K : K < K^{+} \text{ or } K > K^{-}\}$ satisfy the inequality. Lastly, note that $K^{+} > 1$, which implies that we can find a $0 < K < 1$ that satisfies inequality (20) and, hence, makes $d(K) > 0$, the numerator of equation (19) positive, and $\dfrac{\partial P_{\tau}^{m}}{\partial N} < 0$.

Using the steps of the proof, we can find an example that generates $\frac{\partial P_{\tau}^{m}}{\partial N} < 0$. Let cohort size in steady state be $N = 100$ and the increase at time τ be $\Delta N = 1$. Moreover, let $\delta = 0.25$, $\beta = 1$, $p = 1$, and $\gamma = 0.5$. Then $B = \frac{\gamma(1-2\delta)^2}{2} = \dfrac{1}{16}$. Since any $0 < K < 1$ generates the desired result, we choose $K = 0.5$. Under these conditions we have that, in steady state, $\underline{\eta}_{0,ss} = BK = 0.03125$. The number of older men in steady state is therefore $N_{1,ss} = N\left(1 - K\left(1 - \underline{\eta}_{0,ss} - 2\delta\right)\right) = 76.5625$, $\theta_{0,ss}^{m} = N/(N + N_{1,ss}) = 0.5664$, which implies that $\lambda_{ss} = K/\theta_{0,ss}^{m} = 0.8828$. The probability that a younger man marries in steady state is therefore $P_{ss}^{ym} = K\left(1 - \underline{\eta}_{0,ss} - 2\delta\right) = 0.2344$ and the probability that a man in a

given cohort marries is

$$P_{ss}^{m} = P_{ss}^{ym} + \lambda\left(N_{ss}\right)\frac{1-P_{ss}^{ym}}{2-P_{ss}^{ym}}\left(1-F\left(0\right)\right) = 0.4258.$$

At τ, the period of the cohort shock, we have $\theta_{0,\tau}^{m} = \frac{N+\Delta}{N+\Delta+N_{1,\tau}} = \frac{N+\Delta}{N+\Delta+N_{1,ss}} = 0.5688$, which implies that $\lambda_{\tau} = K/\theta_{0,\tau}^{m} = 0.8790$. The reservation value is the solution of equation (17), which is a quadratic equation in $\underline{\eta}_{0,\tau}$ and has as the only positive solution 0.0311. The probability that a younger man marries at τ is therefore $P_{\tau}^{ym} = K\left(1-\underline{\eta}_{0,\tau}-2\delta\right) = 0.2344$ and the probability that a cohort-τ man marries is

$$P_{\tau}^{m} = P_{\tau}^{ym} + \lambda\left(N+\Delta\right)\frac{1-P_{\tau}^{ym}}{2-P_{\tau}^{ym}}\left(1-F\left(0\right)\right) = 0.4250,$$

which is lower than the corresponding probability before the shock.

D.7 Proof of Proposition 5

In the data, we only observe out-of-wedlock births for women. We therefore derive the result for women. Let σ be the probability that a birth occurs during a dating period.

In the congestion model, the probability of observing an out-of-wedlock birth for a woman is given by the probability of meeting a younger man at τ, having a child, and not marry him, plus the probability of meeting an older man at τ, having a child, and not marry him, i.e.

$$P_{\sigma} = P_{\sigma}^{y} + P_{\sigma}^{o} = \sigma\left[\lambda\left(N\right)\theta_{0,\tau}^{m}F\left(\underline{\eta}_{0,\tau}\right) + \lambda\left(N\right)\left(1-\theta_{0,\tau}^{m}\right)F\left(0\right)\right].$$

The derivative of P_{σ} with respect to N takes therefore the following form:

$$\frac{\partial P_{\sigma}}{\partial N} = \sigma\left[\frac{\partial\left(\theta_{0,\tau}^{m}\lambda\left(N\right)\right)}{\partial N}\left(F\left(\underline{\eta}_{0,\tau}\right)-F\left(0\right)\right) + \lambda\left(N\right)\theta_{0,\tau}^{m}\frac{\partial F}{\partial\eta}\frac{\partial\underline{\eta}_{0,\tau}}{\partial N} + \frac{\partial\lambda\left(N\right)}{\partial N}F\left(0\right)\right] < 0,$$

where the inequality follows from $\frac{\partial\left(\theta_{0,\tau}^{m}\lambda(N)\right)}{\partial N} < 0$, $F\left(\underline{\eta}_{0,\tau}\right) - F\left(0\right) > 0$ because $\underline{\eta}_{0,\tau} > 0$, $\frac{\partial F}{\partial\eta} = f\left(\eta\right) > 0$, $\frac{\partial\underline{\eta}_{0,\tau}}{\partial N} \leq 0$ as shown in the proof of Proposition 4, and $\frac{\partial\lambda}{\partial N} < 0$ since $\frac{\partial\theta_{0,\tau}^{m}}{\partial N} > 0$ as shown in previous proofs and $\frac{\partial\left(\theta_{0,\tau}^{m}\lambda(N)\right)}{\partial N} < 0$.

In the surplus model, the probability of observing an out-of-wedlock birth for a woman is given by the probability of meeting a younger man, having a child, and not marry him, plus the probability of meeting a older man, having a child, and not marry him, i.e.

$$P_{\sigma} = P_{\sigma}^{y} + P_{\sigma}^{o} = \sigma\left[\theta_{0,\tau}^{m}F\left(\underline{\eta}_{0,\tau}^{\phi}+\phi(N)\right) + \left(1-\theta_{0,\tau}^{m}\right)F\left(\phi(N)\right)\right].$$

Note that $P_\tau^{ym} = \theta_{0,\tau}^m - \theta_{0,\tau}^m F\left(\underline{\eta}_{0,\tau}^\phi + \phi(N)\right)$ and, hence, $\theta_{0,\tau}^m F\left(\underline{\eta}_{0,\tau}^\phi + \phi(N)\right) = \theta_{0,\tau}^m - P_\tau^{ym}$. We can therefore rewrite P_σ as follows:

$$P_\sigma = \sigma\left(\theta_{0,\tau}^m - P_\tau^{ym}\right) + \sigma F\left(\phi(N)\right)\left(1 - \theta_{0,\tau}^m\right).$$

Taking the derivative with respect to N, we therefore have

$$\begin{aligned}\frac{\partial P_\sigma}{\partial N} &= \sigma\left[\frac{\partial \theta_{0,\tau}^m}{\partial N} - \frac{\partial P_\tau^{ym}}{\partial N} + f\left(\phi(N)\right)\left(1 - \theta_{0,\tau}^m\right)\frac{\partial \phi}{\partial N} - F\left(\phi(N)\right)\frac{\partial \theta_{0,\tau}^m}{\partial N}\right] \\ &= \sigma\left[\frac{\partial \theta_{0,\tau}^m}{\partial N}\left(1 - F\left(\phi(N)\right)\right) + f\left(\phi(N)\right)\left(1 - \theta_{0,\tau}^m\right)\frac{\partial \phi}{\partial N} - \frac{\partial P_\tau^{ym}}{\partial N}\right] > 0,\end{aligned}$$

where the inequality follows from $\frac{\partial \theta_{0,\tau}^m}{\partial N} \geq 0$, $0 \leq 1 - F\left(\phi(N)\right) \leq 1$, $f\left(\phi(N)\right)\left(1 - \theta_{0,\tau}^m\right) > 0$, $\frac{\partial \phi}{\partial N} > 0$, and $\frac{\partial P_\tau^{ym}}{\partial N} \leq 0$ in the empirically relevant surplus model. Therefore, the surplus model that is not rejected by the data predicts that $\frac{\partial P_\sigma}{\partial N} > 0$.

D.8 Expected Value Functions

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

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

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 woman 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 younger times the probability.

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

form:

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

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