Did the baby boom cause the US divorce boom?*

Tereza Ranošová †

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Abstract

The United States experienced two major demographic 'booms' during the second half of the twentieth century, in births after the second world war and in divorces 25 years later. This paper argues that the two booms are linked. As the baby-boom generations were entering marriageable age, men in previous cohorts were faced with exceptionally good remarriage prospects motivating them to rematch. The cohorts who ultimately divorced most were the ones with the biggest increase in remarriage opportunities for men. Using cross-state variation in the size of the baby-boom, I show that marriages in the pre-boom generations were more likely to divorce the bigger the relative supply of young women. This conclusion is robust to instrumenting the size of the baby-boom with WWII mobilization rates. Lastly, I construct a simple dynamic marriage market model which can generate a divorce boom caused by a baby-boom, and can account for between a seventh and a third of the rise in divorces in the 1970s.

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†University of Michigan, tranosov@umich.edu.

1 Introduction

Between 1960 and 1980 the divorce rate in the United states more than doubled. This was also the time when large post World War II cohorts started entering the marriage market. In this paper I propose a causal link between the two concurrent events. I hypothesize that more marriages started to break up during this time, because the marriage market flooded with younger generations much bigger than the previous ones. As men typically match with younger women, a large young cohort presents a supply shock to the remarriage options of husbands in existing partnerships. At the same time sharp cohort size growth results in a shortage of eligible men, making it more worth while for young women to match with already married men.

Divorce is a large and common disruption in the lives of American families. By age 45 one third of first marriages end in divorce (Stevenson and Wolfers, 2007) and 4 out of 10 children in the U.S. will experience the divorce of their parents (Garcia-Moran, 2018). It represents a large income shock, especially for women and children.[1][2] In fact, Moffitt (1992) notes that most exits and entrances into welfare are associated with changes in family structure (not with changes in e.g. labor market circumstances) and Chetty et al. (2014) shows that the share of single-parent families is the strongest predictor of geographical variation in intergenerational mobility. Yet why people divorce is not very well understood.

This paper argues that a major reason for divorce is rematching of men to younger women (whom they prefer) which causes marriages among peers to break up. Survey evidence consistently shows that men prefer younger women, while women prefer men of their own age (or older).[3] Many features of existing marriage and divorce behavior are consistent with asymmetric age preferences causing divorce. Marriages are consistently forming more between younger women and older men, a tendency that is stronger for later-age marriages and for remarriages (England and McClintock, 2009). Divorced people rarely stay single. Stevenson and Wolfers (2007) show that by age 45, 69% of those who divorced have remarried. Men remarry more than women and soon after the divorce (Browning et al., 2011).

Showing that divorces are at least partially driven by preferences of men for younger women also sheds light on an important gender asymmetry in the lifecycle of marriage. In particular, if men prefer younger women while women prefer men of the same age, the remarriage prospects of the wife deteriorate faster relative

[1] Depew and Price (2018), Ananat and Michaels (2008) and others use the sex of the first child as an instrument for divorce to confirm this link is in fact causal.

[2] See Chiappori et al. (2018) for a review of the evidence on negative effects of divorce on reported subjective wellbeing beyond the loss of income. See Del Boca (2003) for a review on the negative effects of divorce on children.

[3] See for example Bozon (1991), Kenrick et al. (1996). Kenrick and Keefe (1992) shows evidence of age preferences from dating newspaper advertisements. Recently Low et al. (2018) showed that this result holds even when controlling for physical appearance. These preferences are often rationalized with gender differences in fecundity by age (for example inGiolito (2004), Díaz-Giménez and Giolito (2013) and Low (2017)). Shephard (2019) shows using an estimated a random search steady state matching model with limited commitment that the increase in utility for a man whose wife is around five years younger is equivalent to his private consumption being 50% higher.

to the husband as a couple ages together. Consistent with this hypothesis, Bruze et al. (2015) finds that as men and women get older, husbands receive a larger share of the marital surplus. This mechanism has consequences for the inherent stability of different marriage matchings, the importance of common property as an insurance mechanism (Lafortune and Low, 2017) and the welfare implications of policies concerning division of property at divorce, joint taxation, spousal income effects on social security and others.

In this paper I provide evidence that the entry of baby-boomers on the marriage market in the 1970s was at least partially responsible for the rise in divorces. First, in section 2 I show that on the aggregate the timing of the divorce boom and the incidence among cohorts matches up with this hypothesis. Divorce rates started increasing sharply when the cohorts who should be most affected by the entry baby-boomers were in their prime divorcable age. Moreover, these cohorts also had the highest chances of ultimately having divorced. Section 3 presents cross-state evidence that the pre baby-boom cohorts divorced more by 1980 if born in a state with a bigger baby-boom between 1930 and 1950. I show that this correlation is robust to controlling for differences in a variety of socio-demographic characteristics. Lastly, I confirm that the estimated effect of a steep cohort size increase on divorces of preceding cohorts is in fact stronger if the size of the early baby-boom is instrumented with WWII mobilization rates (a strategy motivated by Doepke et al. (2013)). This provides further evidence that the cross-sectional correlation is causal and not driven by common state specific factors such as trends in pro-family values. Section 4 shows that a simple repeated matching marriage market model, where divorce is driven by remarriage, can generate a divorce boom out of a baby-boom entry. Despite its simplicity, this approach is novel in capturing the proposed mechanism. Models of the marriage market currently used in the literature do not provide a useful starting point for this hypothesis, because they are either ignoring divorce and remarriage altogether, or divorce is not attributed to remarriage and rematching is only possible out of singlehood.[4] The within period matching market is based on Choo and Siow (2006).[5] The model matches the timing of the boom and the incidence across cohorts, but misses the persistence of the divorce bust after 1990. Quantitatively, it explains between a fifth and a third of the divorce boom (depending on the metric chosen).

This paper contributes to the literature studying the effects of cohort size variation on the marriage markets. A long tradition in demography since Groves and Ogburn (1928), and most recently Abramitzky et al. (2011), has recognized that a variation in cohort size combined with a common age gap in marriages causes

[4]Notable exceptions of models that do allow for rematching 'on-the-job' (building on Mortensen (1988)) are Cornelius (2003) and Burdett et al. (2004), who work with stylized theoretical random search models with non-transferable utility. In spirit, the model closest to the hypothesis in this paper is Shephard (2019), which builds a random search model with rematching with age as a key variable of interest. It recognizes that divorce at older ages can happen, because men realize their preferred (younger) partners are now available. However, this paper still does not allow rematching straight from marriage (that would motivate divorce) and the model is too complex to study dynamics (the paper only studies a steady state).

[5]Choo and Siow (2007) and Choo (2015) also extend Choo and Siow (2006) to a dynamic setting. In many aspects these models are more complex, yet none of them allows divorce motivated by rematching (in fact they assume exogenous divorce).

a 'marriage squeeze', affecting *marriage rates*. Bergstrom and Lam (1991), Bergstrom and Lam (1994) and more recently Ni Bhrolcháin (2001) show that the effects on marriage rates are largely mitigated by adjustments in the age gap itself. Brainerd (2017), and indirectly Bronson and Mazzocco (2018), provide evidence that *men investing* more or less in marriageable capital such as education and experience can also be a margin on which the marriage market adjusts to a cohort size variation. I propose that *divorce and remarriage* (serial monogamy) can help the marriage market absorb a large increase in cohort size.

In addition, this paper adds to the literature trying to understand the rise in divorces in the US in the 1970s. Most existing explanation relate to the concurrent decrease in the wage-gap/ increase in female labor supply (e.g. Ruggles (1997), Weiss and Willis (1997)).[6] McKinnish (2007) shows that sexually integrated workplaces cause divorce (through remarriage) and Greenwood et al. (2016) shows that a technological progress in the household sector can be behind both an increase in divorces and an increase in female labor supply. A large literature discusses whether the boom in divorces is caused by liberalized divorce laws in the early 1970s (see e.g. Friedberg (1998), Wolfers (2006)), though Stevenson and Wolfers (2007) concludes that despite apparent conflict in this literature, liberalized divorce laws had at most a small effect on divorce rates.[7][8]

2 Aggregate evidence

The United States experienced two major demographic 'booms' during the second half of the twentieth century. Between 1945 and 1960 the number of births increased sharply (see figure 1b). Approximately 25 years later divorces started to rise, peaking around 1980 (see figure 1a). This paper argues that the two booms are linked. As the baby-boom generations were entering marriageable age, men in previous cohorts were faced with exceptionally good remarriage prospects. This hypothesis requires that the divorce boom must have a strong cohort component, because the entry of baby-boom generations is expected to affect the cohorts immediately preceding more than others. Indeed, figure 2a shows that there is a boom and bust also in the share of men ever-divorced by cohort. Men born between 1930 and 1940 were the ones for whom the chances of divorce increased sharply. The cohorts of men born between 1940 and 1950 had the highest chance of ever experiencing a divorce. By the time the divorce rate started increasing, between the years 1960 and 1980, these men were entering their 30s, which is a common age to get divorced.[9] Figure 2a shows

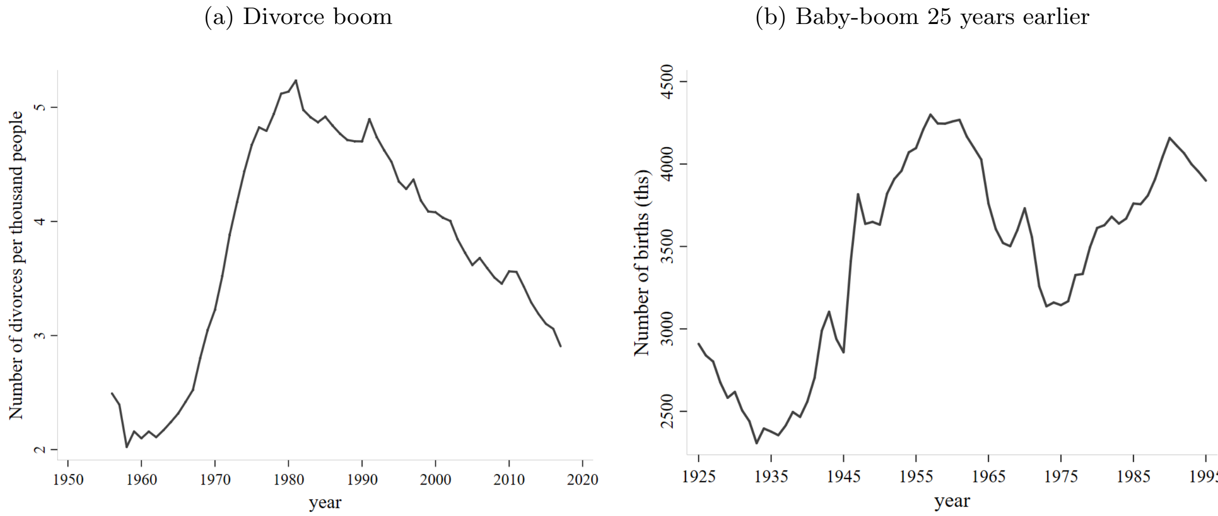
[6]However, it is also natural to suspect reverse causality from divorce to female labor supply, as showed in Johnson and Skinner (1986) and most recently by Goldin and Katz (2016).

[7]Interestingly, this paper suggests divorce laws might have been liberalized because because of a mounting demand for divorces as the baby-boomers were entering marriageable age. De La Croix and Mariani (2015) hypothesizes that throughout history marriage norms have been changing to accommodate demand for polygyny in otherwise monogamous societies. This theory is consistent with the divorce laws changing in the 1970s.

[8]Other potential suspects include social attitudes towards marriage (Cherlin, 2004) and generosity of the welfare state (Moffitt, 1997).

[9]See appendix A.2.

Figure 1: Divorce boom and Baby-boom



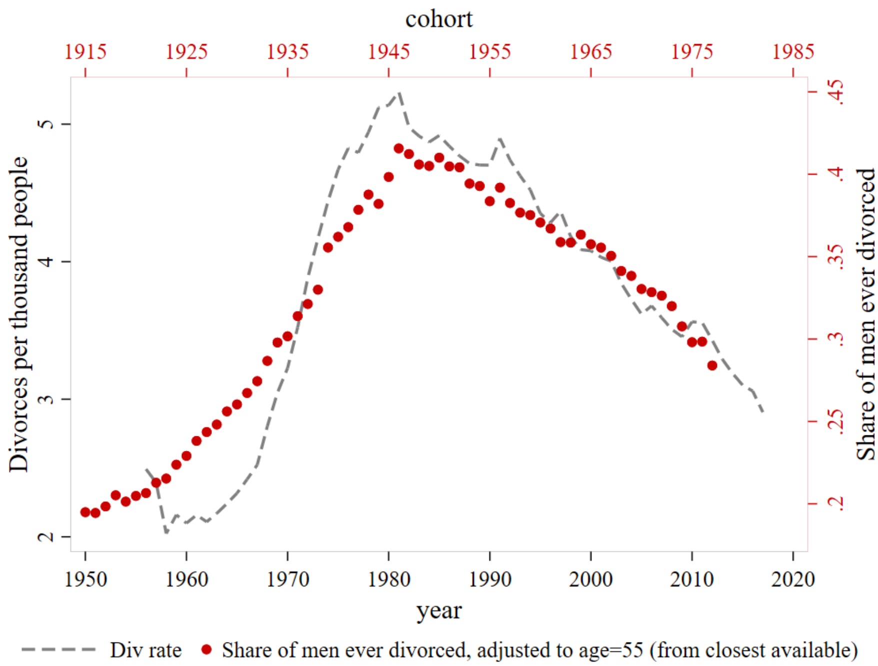
Source: Vital Statistics (divorce rate in panel 1a is based on a subsample of states with coverage over the whole sample).

that the peak of the divorce boom coincides with the years when cohorts who divorced most were 34 years old. This was also the time when the age distribution in the marriage market relevant for men in their 30s shifted dramatically. Figure 2b shows a sharp increase between 1960 and 1980 in the number of people older than 34 compared to the number of people younger than 34 (by up to 6 and 12 years respectively), peaking around 1980. This means that by 1980 the prospects for a 34 year old man to remarry to a younger woman were especially high as there were many more younger women compared to the number of men who would be competing with him. Subtracting the lag of 35 years in figure 2b implies that the remarriage prospects were most favorable for men born between 1940 and 1950, exactly the cohorts who divorced the most.[10]
Lastly, table 1 shows that after the age-gap in men's higher order marriage is systematically bigger than in man's first marriages by about 3 years, confirming the pattern that after divorce men typically remarry to

[10]The divorce rates are also high for the cohorts 1950-1955, even though the age distribution has already been stabilizing. It is very possible that this persistence in divorce risk is still endogenous to the mechanism discussed here. Early baby-boom cohorts were increasingly entering in marriages with a lower age-gap (as shown in figure 20 and predicted by Bergstrom and Lam (1991)) and marriages to divorcees, precisely because young women were in over-supply. These two characteristics correlate with a higher risk of divorce. This can explain why divorce rates remained high for the baby-boomers, despite falling remarriage prospects. Section D in the appendix discusses this hypothesis in more detail. It is also possible that after divorce rates increase, the fundamentals of the marriage market change. For example, Chiappori and Weiss (2003) and Chiappori and Weiss (2007) suggest that a small initial increase in divorce rates can be multiplied and turn permanent, because the remarriage market is now more favorable to everybody. This mechanism is requiring that single divorcees are more available to find a match than married individuals, which I am not able to capture in my simple model.

Figure 2: Cross-state variation.

(a) Divorce rate (number of divorces a year per thousand people) (left), Share men ever divorced (right)



(b) Size of the new entry shock.

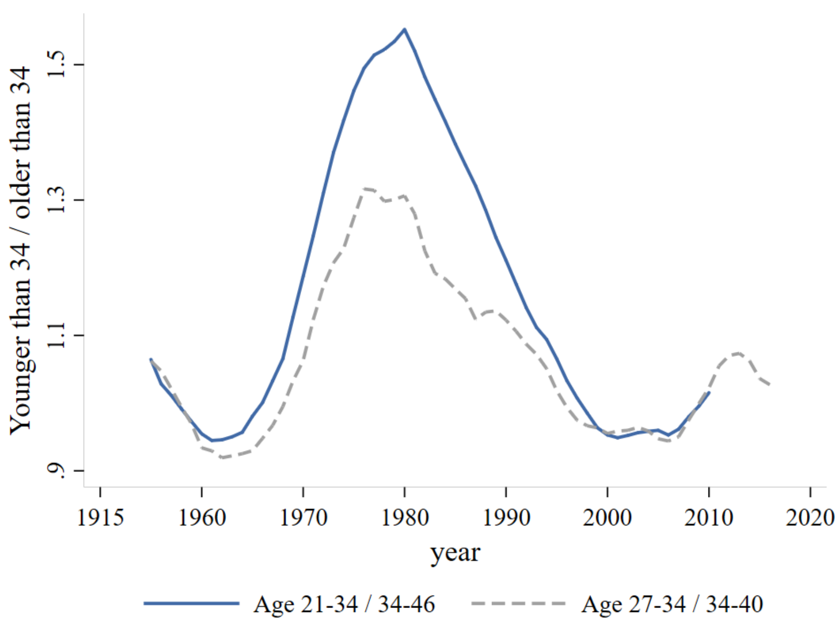


Figure 2a: the divorce rate (on the left axis) by year and the share of men in the cohort born 35 years ago who ever went through a divorce (or remarriage), measured at age 50 or closest available and regression adjusted to age 50 (see section in the appendix). Figure 2b: change in the supply of younger women to older men, measured by the 0 to 6/12 years younger over 0 to 6/12 years older ratio by age 34 (and so for cohorts born 35 ago). This kind of measure is always high whenever the cohort size is rapidly increasing and is equal to 1 in a stable population. Source: IPUMS Census, Vital Statistics.

younger partners.

Table 1: Increase in the share ever-divorced when remarriage options of men improve

	Mean	sd	mean - sd	mean + sd
First marriages, all	2.92	4.66	-1.74	7.58
First marriages, ages 35-44	2.49	4.12	-1.62	6.61
Higher order marriages, all	6.00	8.00	-1.51	13.89
Higher order marriages, ages 35-44	3.85	6.16	-2.31	10.01

Source: 1960 Census

Overall this shows that the timing of the divorce boom and the incidence across cohorts is consistent with the hypothesis that the baby-boom caused the divorce boom. Section A.3 in the appendix shows additional descriptive evidence on the adjustment of the marriage market to the baby-boom. First, the age-gap in marriage was higher and rose more quickly over their lifetime for the men born around 1940 compared to men in other cohorts. These men also stayed single slightly less, almost closing the gender gap in ever getting married. Lastly, compared to men, women born around 1940 who ever went through a divorce remarried less and stayed divorced more.

3 Leveraging cross-state variation in the baby-boom

In this section I present evidence that being from a state with a large early baby-boom (large post 1945 cohorts compared to 1935-1944) caused the treated cohorts (of 1935-1944) to divorce more. I focus on this group for several reasons. These were the cohorts who were most likely responsible for the sharp run up of the divorce boom. They were in their prime divorce age in the 1970s[11] and they ultimately ended up having a 10 percentage points higher chance of ever getting divorced than cohorts 10 years older (the sharpest increase compared to other cohort groups: see figure 2a). Second, by 1980 they were 35-44 years old, which is old enough to limit the worry of selection into marriage, but young enough for differential mortality not to matter. Since remarriages are not observed in the Census for 1990 and 2000, younger cohorts are too young to be studied by 1980 (as many men have not married yet, selection into marriage makes comparison across cohorts problematic). Third, building on Doepke et al. (2013) I will argue that the relative supply of younger partners available to these cohorts (the size of the early baby-boom of 1945-1954 compared to the

[11] See Section A.2 in the appendix.

pre-baby-boom cohorts of 1935-194) can be instrumented by WWII mobilization rates. This allows me to rule out that the correlation between divorce probabilities and remarriage supply is caused by confounding factors such as persistent cultural norm differences across states towards high fertility and few divorces.

To study the share of a given cohort who divorces by a certain age, I primarily use pooled IPUMS Census data for 1960 and 1980 (Ruggles et al., 2019), focusing on men and women age 35 to 44 (by 1980, the cohorts born 1935-1944 reached this age, cohorts born 1915-1924 who were 25-44 years old in 1960 are used as a control). I restrict the analysis to those born in the United States (as immigrants tend to not fully integrate into local marriage markets)[12]. I also exclude those living in group quarters for the same reason.[13] A person is classified as being ever-divorced, if their current marital status is divorced or if their last marriage was not their first.[14] To supplement the main sample, I use IPUMS Census data to construct a state level measure of remarriage opportunities for the cohorts of interest. Specifically, I compute the share of people aged 30-34 in the group 30-44, ${\frac{n_{30-34}}{n_{30-44}}}_{st}$, in each state and year. Between 1960 and 1980 this measure captures the variation in the steepness of the early baby-boom (of cohorts 1945-1949). As a robustness check, I repeat the analysis using a more direct measure of this cohort size increase, the share of children 0-4 among children 0-14 30 years ago, ${\frac{n_{0-4}}{n_{0-14}}}_{s,t-30}$ (using IPUMS Census data from 1930 and 1950). The results are presented in section B.1 in the appendix.

The main hypothesis is that people in their late 30s born in states with a steep improvement in remarriage opportunities for men had a higher chance of ever getting divorced. Cross-state differences in fertility can arise for a multitude of reasons, some of which could be persistent and correlate with the likelihood of divorce (violating the exogeneity of the remarriage opportunity measure). For example, cultural norm differences across states towards high fertility and few divorces would bias the OLS results downwards. To rule out that the results are driven by this kind of omitted variable bias, I instrument the change in the remarriage opportunity measure between 1960 and 1980 with WWII mobilization rates (as used by Acemoglu et al. (2004)). This strategy is based on the evidence in Doepke et al. (2013) suggesting that WWII mobilization was at least partially responsible for the baby-boom. The hypothesized mechanism goes as follow: women (mainly in cohorts 1905-1914, but more broadly in 1905-1924) worked during the war and beyond, this labor supply 'shock' depressed wages of women (being an imperfect substitute for men in the labor market), incentivising the cohorts of 1925-1934 (who were 15-24 years old by 1950) to stay at home and increase

[12]A fact explored in the marriage literature when variation in migration rates can create immigrant specific variation in sex ratios, as in Angrist (2002).

[13]Lastly, people born in Hawaii, Alaska and DC are also excluded from the analysis, as data on WWII mobilization rates are not available for these states.

[14]This information is available in the 1960-1980 Census and the 2008 onwards ACS. Notice it would be unreasonable to use simply the marital status of being divorced, because the main hypothesis is that divorces happened because of a desire to remarry. Unfortunately, this variable is an imperfect proxy, as it also includes people who remarried after their spouse died. This is another reason to use a relatively young age group, for which mortality is still low.

fertility (realized between 1945 and 1960).[15][16]

Figure 3a displays a cross plot of state level mobilization rates on the x-axis with the change in the remarriage opportunity measure on the y-axis, showing a strong positive correlation (confirming that the fertility variation explained by Doepke et al. (2013) did result in cross-state variation in relative cohort size 30 years later). Second, figure 3b shows a positive correlation between the change in the remarriage opportunity measure with the growth in divorce (measured by a change in the share ever-divorced in the relevant age range).

(a) Mobilization rates affecting the size of the baby-boom. (b) Remarriage opportunities for men affecting divorce.

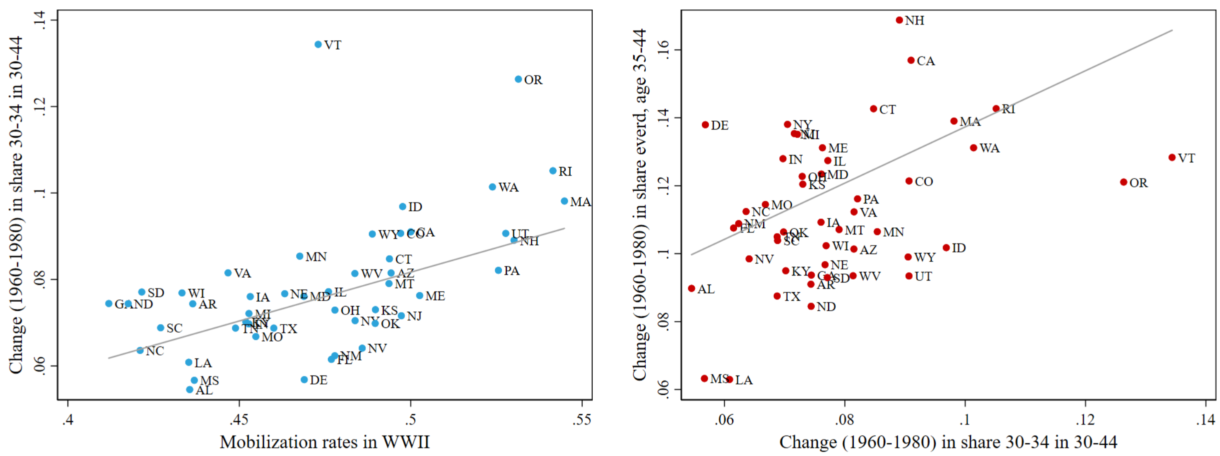


Figure 3: Share of ever-divorced attached by state of birth. Data source: IPUMS Census data from 1960 and 1980, state-level mobilization rates during WWII from Acemoglu et al. (2004). States weighted by population in 1930.

Next, I confirm that this correlation is statistically significant in a regression, and robust to controlling for observable differences among states and to instrumenting change in the remarriage opportunity measure with mobilization rates during WWII. The main specification (equation 1) regresses a dummy variable y_{ist} of being ever-divorced (where i stands for an individual, t for a year $\in \{1960, 1980\}$ and s for a state of birth)[17] on state and year fixed effects, the remarriage opportunity measure $\frac{n_{30-34}}{n_{30-44}}_{st}$, and a set of individual

[15]Notice this mechanism should directly affect the cohorts 1915-1924, which serve as a control in the baseline analysis. Specifically the mechanism predicts that these cohorts have a higher labor supply in high mobilization states. Since female labor supply is, if anything, associated with higher chances of divorce, this potential source of invalidity of the exclusion restriction should bias against finding an effect when comparing the treated cohorts to the control. Section B presents a further discussion of this concern.

[16]Section A.4 in the appendix provides more detail on the variation in mobilization rates.

[17]I classify people based on their state of birth to avoid spurious correlation in divorcing behavior with internal migration decisions, for marriage or after divorce. Moreover, it is not clear whether the relevant marriage market pool is closer to the

level sociodemographic controls. The running null hypothesis is that people born in states with a bigger increase in remarriage opportunities were more likely to get a divorce ($\beta > 0$).

$$y_{ist} = \alpha_s + \alpha_{1980} + \beta \frac{n_{30-34}}{n_{30-44}}_{st} + \gamma X_{ist} + \epsilon_{ist} \tag{1}$$

Table 2 presents the baseline results. All specifications include dummies for age, sex and race. Columns 1 and 2 present the OLS results, confirming what figure 3b has suggested. Individuals born in states with higher remarriage prospects are more likely to go through a divorce (compared to their counterparts 20 years ago). Column 2 adds additional controls, namely dummy variables[18] for levels of education, whether an individual lives in a metro are and farm status. These controls are motivated by baseline cross-state differences (already observed by 1940) between states with high and low mobilization rates (see section A.4 for details). If anything, including these controls makes the OLS results slightly bigger in magnitude.

Columns 3 and 4 mirror columns 1 and 2, but instrument $\frac{n_{30-34}}{n_{30-44}}_{st}$ with WWII mobilization rates interacted with a dummy for 1980. The specification of the first stage is directly motivated by Doepke et al. (2013). Results of the first stage regressions are presented in table 8 in the appendix. Table 2 shows that instrumenting the remarriage opportunity measure only strengthens the results. This suggests that perhaps the correlation between remarriage opportunity and divorce chances is mitigated by omitted variables such as persistent trends in pro-family values. Column 5 shows that the result is robust to adding region times year fixed effects (being identified of within region differences in the growth of remarriage options). Column 6 shows that the main conclusion is robust to only restricting to people who ever got married, so it is not driven by selection into marriage.

Overall, this section presents robust evidence of the main mechanism proposed in this paper. Married men surrounded by an increased opportunity to rematch with a younger spouse were more likely to divorce. On the aggregate $\frac{n_{30-34}}{n_{30-44}}_t$ rose from 0.329 in 1960 to .407 in 1980 a difference of 0.077. As a result, the coefficients presented in table 2 if used to predict the change on the aggregate imply an increase of between 7 and 15 percentage points in the probability of divorcing for 35-44 years old. In the data this measure rose on the aggregate by 12 percentage points. The cross-sectional variation can explain the rise in divorces in this age-group.

state of birth or the state of residence (as individuals can for example temporarily move for a job or to get a degree and then move back 'home' and settle down).As a consequence, the estimates are in spirit closer to measuring an intent-to-treat effect. When state level remarriage opportunities are assigned based on the state of residence, the results remain qualitatively similar and statistically significant, yet slightly smaller in magnitude and much noisier. This suggests that internal migration decisions mainly add noise to the analysis.

[18]Including dummies for all categories of these variables as provided by IPUMS.

Table 2: Increase in the share ever-divorced when remarriage options of men improve

	Ever-divorced					
	OLS		2SLS			
$\frac{n_{30-34}}{n_{30-44}}_{st}$	0.809	0.900	1.485	2.065	1.885	2.011
	(0.238)	(0.296)	(0.387)	(0.457)	(0.429)	(0.444)
X_i:						
sex, race	yes	yes	yes	yes	yes	yes
farm, metro	no	yes	no	yes	yes	yes
educ dummies	no	yes	no	yes	yes	yes
region-year fes	no	no	no	no	yes	no
In ever-married	no	no	no	no	no	yes

N = 2032220 (1902899 in column 5), 48 clusters

t statistics in parentheses. SEs clustered at the state level.

All regressions include year, age and state fixed effects.

Pooled cross-section 1960 and 1980 IPUMS Census data, fitting a linear probability model of being ever-divorced and divorced on $\frac{n_{30-34}}{n_{30-44}}_{st}$, a share of 30-34 year olds in 30-44 year olds (aggregated on the state level). Restricting to a sample of ever-married men and women 35-44 years old. All columns include year and state fixed-effects and demographic controls. In columns 3-6 $\frac{n_{30-34}}{n_{30-44}}_{st}$ is instrumented by WWII mobilization rates interacted with an indicator for 1980.

4 Can a steep cohort size growth cause a boom in rematching?

In this section, I present a simple repeated matching model calibrated to the pre-divorce-boom marriage market characteristics to illustrate that a baby-boom naturally generates a divorce boom, through the simple channel of men rematching with younger women more. The model is set up with the following strategy in mind: I make assumptions necessary to make the decision about marriage effectively static and happening repeatedly in a competitive matching environment based on Choo and Siow (2006), while keeping initial divorce rates sufficiently low. At the same time, aggregate supply of men and women of different ages is dynamic and exogenous (taken to match the data). Overall, this model can generate a divorce boom (and explain between a seventh and a third of the rise in divorce), despite being very simple and focusing solely on divorce motivated by rematching keeping initial motivations for matching by age fixed.

The setup of the model is as follows. Both men and women are only differentiated by age: $i, j \in \{21, ..., 69\}$. Every period starts with the distribution of existing marriages N_{ij} (number of marriages among men of age

i and women of age j) and singles $N_{i\emptyset}$, $N_{\emptyset,j}$. The length of a period is 4 years. Cohorts enter the marriage market in a scattered pattern. Women enter first (a share α_f of women enters at age 21, the rest at age 25). Men enter later (a share α_m of men enters at age 25, the rest at age 29). The main motivation for the staggered entry is to limit initial unnecessary divorces as much as possible.[19] Once a person joins the market, they continue to participate each year (until they die or are widowed). Marriages survive if both partners choose to match with their existing type. Otherwise, a divorce occurs. At the end of a period, the oldest cohort dies. At the start of next period, a new cohort enters the system at age 21.

When a man of type i arrives on the marriage market, he chooses a woman of any age or to stay single, to maximize his lifetime utility from marriage according to the definition in 2.

$$\begin{aligned} &\forall i \in 1, ..., I \\ &V(i, \epsilon_p) = \max_{j \in 1,2,...,J,\emptyset} v_{ij} - \tau_{ijt} + \epsilon_{ijp} \\ &\qquad\qquad + \ \beta V(i+1, \epsilon_p) \end{aligned} \tag{2}$$

v_{ij} are exogenous systematic gains from a match ij for a man of age i who matches with a woman of age j (or no woman $\emptyset$) respectively. When being born every person p is endowed with list of age-gap specific preferences $\epsilon_p = (\epsilon_{ijp})_{\{i \in 1,...,I; j \in \emptyset, 1,...,J\}}$ for single-hood and all possible partner ages, given own age. $\tau_t^i = (\tau_{i1t}, ..., \tau_{iJt})$ is a vector of one time endogenous transfers a man of type i would have to pay when matching with a type j. Cohort size variation influences the individual choice through affecting transfers in a standard general equilibrium sense. The optimal choice of a woman is defined equivalently (3).

$$\begin{aligned} &\forall j \in 1, ..., I \\ &U(j, \epsilon_{p'}) = \max_{i \in 1,2,...,I,\emptyset} u_{ij} + \tau_{ijt} + \epsilon_{ijp} \\ &\qquad\qquad + \ \beta U(j+1, \epsilon_{p'}) \end{aligned} \tag{3}$$

Since the continuation value does not depend explicitly on the current match, the matching decision as effectively static.

This simple assumption has one exception. To avoid a large surge of 'grey' divorces, I remove widows and widowers from the marriage market. If windows return to the market after their husbands die, they create a sudden increase in the supply of older women, changing the incentives for everybody to match and destabilizing many current marriages. If widowers stay of the market (and preferences for age gaps are

[19]The initial entry in the marriage market is consistent with the intuition that different people become 'ready for marriage' at different ages.

stable), existing matches are stable at older ages when new entrants to the market are mostly too young to be relevant rematching partners. However, to preserve the static nature of the matching decision, I assume that this drop-off from the market comes as a surprise to widows and widowers. This way the decision to marry a partner of age $I-1$ is not affected by the future market prospects.[20]

Transfers are determined endogenously in an equilibrium defined as follows:

Definition 1. *Given a sequence of cohort sizes $\{N_t\}$ a dynamic equilibrium in the marriage market consists of sequences of transfers $\{(\tau_{i,j,t})_{\{i=1,...,I;j=1,...,J\}}\}$, numbers of couples of each type $\{(n_{i,j,t})_{\{i=1,...,I;j=1,...,J\}}\}$ and numbers of single men $\{(n_{i,\emptyset,t})_{\{i=1,...,I\}}\}$ and women $\{(n_{\emptyset,j,t})_{\{j=1,...,J\}}\}$ such that*

1. *Given transfers every period the choice of a partner $j^*(i,\epsilon_p)$ by a man who enters the marriage market solves the maximization problem as defined in 2.*
   The choice of a partner $i^*(j,\epsilon'_p)$ by a woman who enters the marriage market solves the maximization problem as defined in 3.

2. *Markets clear, such that*

$$\forall i \in 1,...,I; j \in 1,...,I \;\; n_{i,j,t} = \int 1_{j^*(i,\epsilon_p)=j,p \text{ on the market}} dp = \int 1_{i^*(j,\epsilon'_p),p' \text{ on the market}} dp'$$

$$\forall i \in 1,...,I;\;\; n_{i,\emptyset,t} + \sum_{j=1}^{J} n_{i,j,t} = m_{i,t}$$

$$\forall j \in 1,...,J;\;\; n_{\emptyset,j,t} + \sum_{i=1}^{I} n_{i,j,t} = f_{j,t}$$

where $m_{1,t}=0$, $f_{1,t}=\lambda_f N_t$, $m_{2,t}=\lambda_m N_{t-1}$,
$m_{it}=N_{t-i+1}-W^m_{i,t}$, $f_{j,t}=\lambda N_{t-j+1}-W^f_{j,t}\forall i>2, j>1$.
with $W^m_{i,t}=\sum_{s=1}^{i-1} n_{i-s,I,t-s}$ and $W^f_{j,t}=\sum_{s=1}^{i-1} n_{I,j-s,t-s}$

Lastly, to describe the solution I need to put distributional assumptions on the idiosyncratic preferences.

Assumption 1. *Every person when born is endowed with a vector of preferences for being single versus being matched with all possible age-gaps, that are independent within and across individuals, and are distributed identically standard extreme value type I:*

$$(\epsilon_{\emptyset,p}, \epsilon_{a(1)-a(I),p}, ..., \epsilon_{0,p}, ..., \epsilon_{a(I)-a(1),p})$$

Every period these are used to define the relevant vector of preferences: $\epsilon_{ijp}=\epsilon_{a(i)-a(j),p}$ and $\epsilon_{i\emptyset p}=\epsilon_{\emptyset,p}$

[20]This assumption is fairly innocuous in steady state with the parametrization bellow. It is equivalent to assuming that that widows get a choice between enjoying the period benefit from their previous marriage (perhaps a warm glow from the correct match) or returning to the market. Since transfers are stable in the age gap, widows would choose to not join the market and rather enjoy the period benefit from their optimal choice.

Assumption 1 has two notable parts. First, I introduce persistence in preferences. Namely, the idiosyncratic preference for being single is constant for an individual over their life-cycle, and the idiosyncratic preferences defined over age-gaps, instead of a specific age, are constant. This is done to so that people who match with someone because they have a high idiosyncratic error are incentivized to stick with this match. Second I use a convenient distribution that will allow for an approximate closed form solution of period specific matches $n_{i,j,t}$.

Assumption 2. *Entry of men at age* $i = 2$ *is determined by* $F(\epsilon_{8,p}) \leq \alpha_m$*. Entry of women at age* $i = 1$ *is determined by* $F(\epsilon_{0,p'}) \leq \alpha_f$*.*

Assumption 2 states that for both men and women early entrants on the marriage market are selected not to have a high idiosyncratic preference for the kind of match that will be only possible to realize one period after their entry. This assumption, as well as the staggered timing of entry, is designed to minimize unnecessary divorces as much as possible (as with flexible rematching the biggest challenge in matching the initial marriage patterns is to match a low initial rate of divorce).[21]

The number of divorces each period is given by

$$D_t = \sum_{i \in \{1,...,I-1\}} \sum_{j \in \{1,...,I-1\}} \lambda \; n_{ij,t-1}(1 - P(j' = j+1|i+1)P(i' = i+1|j+1)) \tag{4}$$

The probability that an existing couple who got to try rematching is not divorced is equal to the chances that both the husband and the wife choose to stick with their age-gap $P(j' = j+1|i+1)P(i' = i+1|j+1)$. I determine these by numerically integrating the decisions of a simulated population, given equilibrium transfers τ_{ijt}.

Table 3 shows the calibration of the model.

Table 3: Calibrated parameters

ages	$\{21, 25, ..., 69\}$ (I=13)	C	1.25	$v_{ij} =$	$C_m - c(\lvert a(i) - a(j) - gap\rvert)^{1+\sigma}$
$u_{\emptyset,j} = v_{i,\emptyset} =$	0	c	0.3791	$u_{ij} =$	C_f
$\lambda_m = \lambda_f$	0.75	*(ideal) gap*	4	σ	0.1086

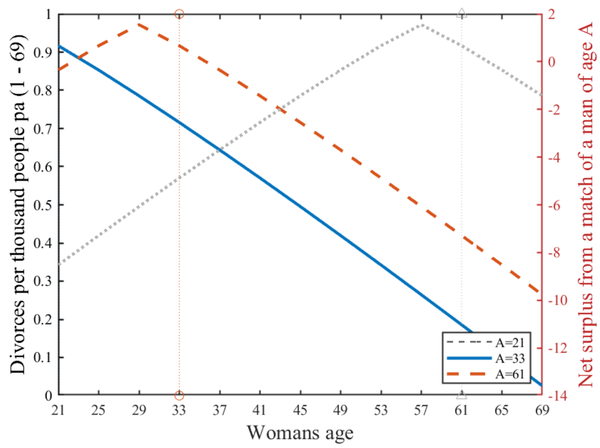
Value of single-hood is normalized to 0. I normalize v_{ji} to a constant (same as the peak value of u_{ij}).[22] Systematic gains to marriage for men u_{ij} are assumed to be constant in $a(i) - a(j)$ (maximizing stability of existing matches). Parameters c, gap and σ govern the nature and strength of the preference of men for

[21] Section

[22] u_{ij} and v_{ji} is not separately identified.

somewhat younger women. I select them to match as closely as possible the matching patterns by age in 1960.[23]. Figure 4 plots $\frac{u_{ij}+v_{ij}}{2}$ for selected ages of men and all ages of women.

Figure 4: $\frac{u_{ij}+v_{ij}}{2}$ for selected ages of men and all ages of women.



C is selected to match the share of people ever married in the 1960 Census while keeping share ever-divorced at bay. λ_m and λ_f is selected to minimize the share of people ever-divorced in the model.

Table 4a summarizes characteristics of the marriage market in the steady state. Overall marriage rates are around 90%, consistent with the data pre-divorce-boom. Women marry on average for the first time at 23 (very close to 21, the average age at first marriage for women who married around 1950 (Rotz, 2016)). Men marry later (as they do in the data), though the average age is slightly higher than in the data (before 1950 the average was approximately 25 (Rotz, 2016)). The share of men and women who ever divorced, and the age at first divorce, are slightly higher in the model than in the data before the boom while the aggregate divorce rate is slightly lower. This suggests that the model underestimates the share of the population that divorces many times.

Starting from a steady state (a constant cohort size), I study the behavior of the model when new cohorts are suddenly bigger. First, I study the reaction to a one time cohort size increase. Figure 5 shows the effects on divorcing behavior. Already in the period when the first part of the bigger cohort enters, divorces (and divorce rates) increase. This is purely an effect of rematching of existing couples, as the youngest cohort certainly could not have divorced yet. The spike in divorces continues in the second period, being a combination of the new entrants breaking up couples formed before period 0, and rematching of couples formed in period 0 (who actually experience a higher chance of divorcing then their steady state counterparts). Figure 5b shows this explicitly, decomposing the increase in divorces into an effect of a divorce risk and a pure compositional effect

[23] I use the number of matches in the 1960 Census in each cross-age group for men 36-55 and women 31-50 years old, where matching is likely to be settled and everybody is reasonably thought of as participating in the marriage market. I use the result from 1 (taken from Choo and Siow (2006)) and loop though parameters to minimaze the sum of squares between the right and

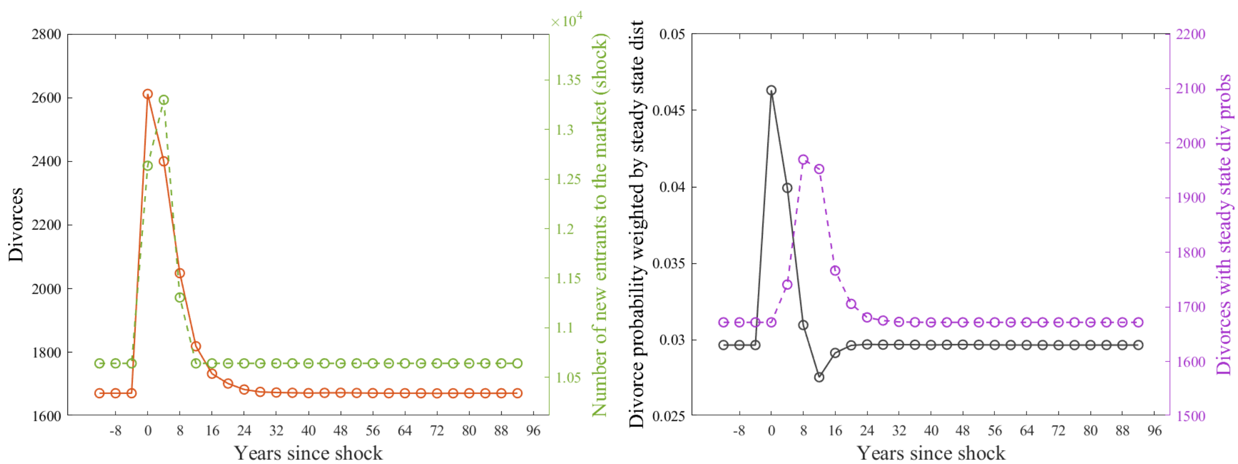
Table 4: Baseline characteristics of the model compared to the pre-baby-boom marriage market.

(a) Baseline (steady state) characteristics of the model

(b) Data

Statistic	Men (a)	Women (a)	Men (b)	Women (b)	Source
Share ever-married	91.7%	91.6%	93.2%	92.7%	*age 60-69 in 1950 Census*
Share ever-divorced	25.1%	25.1%	20.5%	18.4%	*age 60-69 in 1960 Census**
Mean age-gap	3.8		3.7		*all married people in 1950 Census*
Sd of age-gap	4.8		5.4		*all married people in 1950 Census*
Age at 1st marriage	26.2	22.2	25.1	22.1	*all ever-married people from 1960 Census**
Age at 1st divorce	33.2	29.1	34.4	29.2	*women age 40-45, cohort 1920-1934, NSFG*
				~ 34	*women age 50-74, SIPP, from Goldin and Katz (2016)*
Divorces per thousand people	2.2		2.3		*Vital statistics, average 1956-1960*

* Not available in 1950 Census.



(a) Divorces vs. New entrants

(b) Decomposition of the increase in divorces

Figure 5: Impulse response to a one time increase in cohort size by 50%.

of an increase in the number of couples at risk. First, the figure shows the evolution of a sum of probabilities of divorce for couples on the market, $(1 - P(j_t = j|i, j_{t-1} = j-1)P(i_t = i|j, i_{t-1} = i-1))$, weighted by the steady state distribution of couples who could have gotten divorced $\{N_{i-1,j-1}\}$. This weighted probability of rematching, summarizing the average risk of divorce for existing couples, increases in year 0 and remains high in year 4. Second, figure 5b plots what divorces would have been if in every period the number of divorces was calculated with the steady state probabilities $(1 - P(j' = j|i)P(i' = i|j)$. Overall, the increase in divorces caused by a larger number of couples at risk is negligible until 8 years after the initial shock, contributing very little to the overall spike in divorces.

Figures 6 and 7 show the implications of the model for divorces when the cohort size dynamics are calibrated to match the baby-boom, compared with the observed divorce boom in the data. The solid line in figure 6a shows the raw number of births per year in the data aggregated to three years, and replaced by a constant for cohorts before 1928 and after 2009 (to start the model in a steady state)[24]. This represents the exogenous cohort size series as fed into the model.

The model does generate a substantial boom in the divorce rate (calculated as the number of divorces in a period divided by the number of people, including children who have not reached marriageable age yet, and divided by the length of the period in years, to make it comparable with the empirical divorce rate calculated each year). Figure 6b shows that the divorce rate in the model rises from 1.8 in 1960 to a maximum of 2.8 in 1988.[25] This shows the model is roughly successful in matching the timing of the divorce boom. Quantitatively, actual divorce rate increased by approximately 150% (3 in absolute terms) in the data and by only 50% (1 in absolute terms) in the model, suggesting this mechanism is only partial responsible for the divorce boom.

Figure 7 plots the share of men ever divorced by cohort (both in the model and in the data). The model does predict that the divorce boom has a strong cohort component. The cohort measure in the model peaks for men born in 1939, which represents the cohorts responsible for the beginning of the divorce boom. Similar to the aggregate divorce rate, the bust in the cohort measure is faster in the model than in the data. Moreover, the share of men ever divorced starts increasing for cohorts who are in their 30s after 2000 (born in late 1960s). Quantitatively, the share of men ever divorced increased by 100% (0.2 in absolute terms) in the data and by only 15% (0.037 in absolute terms) in the model.

Overall, the model matches the timing of the start of the divorce boom very well. However, the size of the boom in the data is much larger and the divorce bust after 1980 is less rapid.

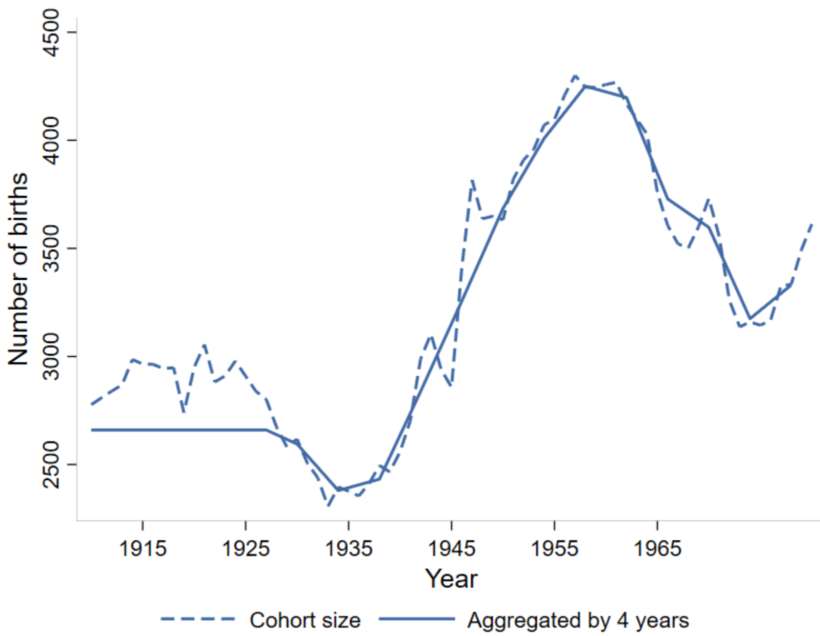
Quantitatively, the model explains between a seventh and a third of the divorce boom (depending on the

left hand sides, restricting *gap* to be an integer.

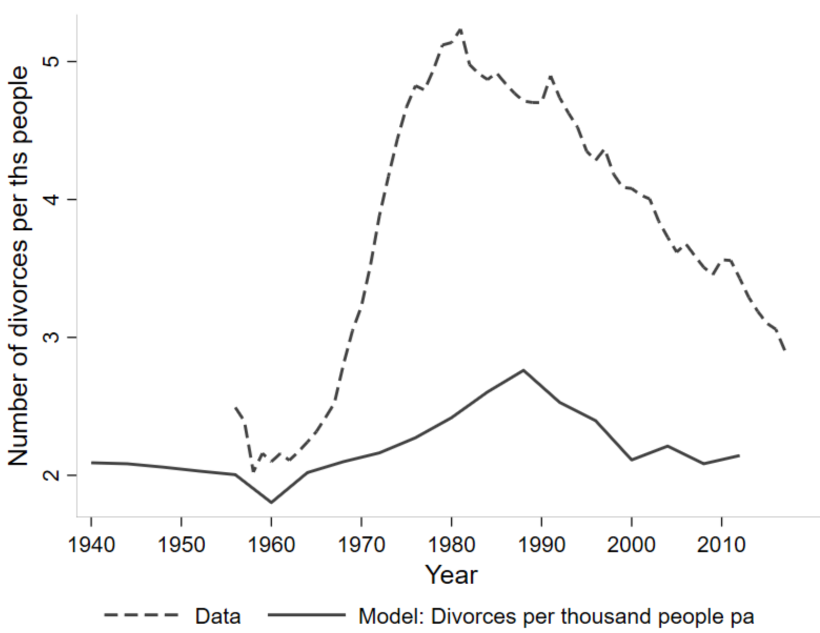
[24]Starting from a level of 2.66 million births per year and finishing at 4 million births per year (which roughly matches the average of a couple of years before 1928 and after 2009)

[25]Notice the divorce rate does not start in steady state, because the upcoming cohort size dynamics affect the denominator.

Figure 6: Divorce boom vs baby-boom 25 years ago in the model vs in the data

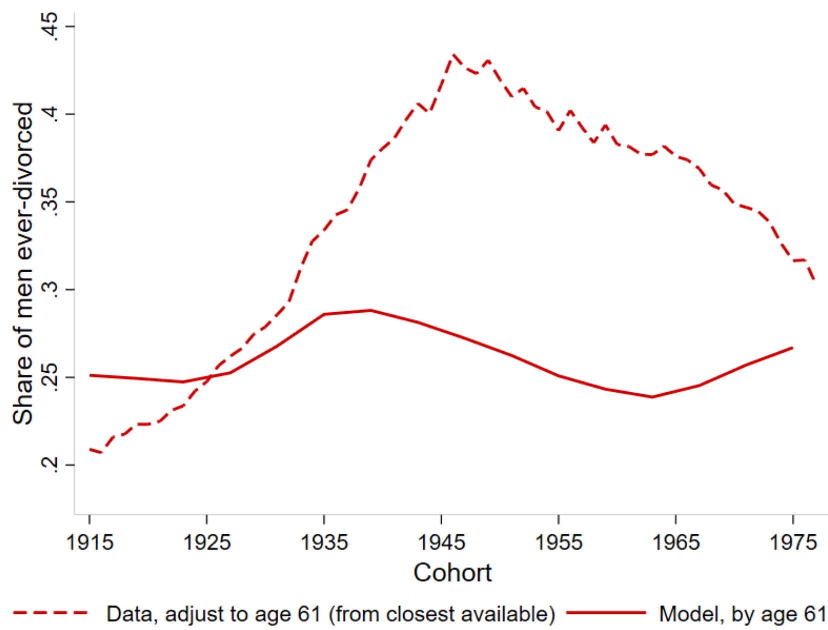


(a) Births (data vs model)



(b) Divorce rate (data vs model)

Figure 7: Divorce boom across cohorts in the model vs in the data.



(a) Model

metric chosen). Together, this is consistent with the hypothesis that the proposed mechanism has likely played a substantial role in the run up of the divorce boom in the 1970s. It also suggests that the model is possibly missing a strengthening and persistence generating mechanism. As suggested by Chiappori and Weiss (2003) and Chiappori and Weiss (2007), even a small spark in divorces can increase divorce rates substantially (and potentially permanently). If unmatched divorcées are disproportionately more likely to look for a partner, and if people can influence their access to new matches (for example through an endogenous search effort or divorce to single-hood), then divorce generates divorce. The simple model proposed in this paper is not capable of capturing such mechanisms. Yet it is likely that it could help explain the mismatch between model and data after 1990. Other potential explanations for the slow, but more persistent, bust can be a change in the structural reasons for marriage. Starting with the baby-boom cohorts, marriage rates decreased substantially. Lower rates of marriage also imply lower rates of divorce. Similarly, the age-gap in marriage has been trending downwards, suggesting that the nature of age-preferences has also been changing.

5 Conclusion

In this article, I propose a causal link between the two demographic 'booms' of the 20th century in the United States. I show that on the aggregate the divorce boom took off right when baby-boomers started to enter the marriage market, creating better remarriage prospects for men in the pre-baby-boom generations. Across cohorts, the divorce boom was more pronounced among men for whom the remarriage opportunities improved the most. Moreover, I provide cross-state evidence that men and women divorced more if born in states with a larger baby-boom, an effect that is robust to controlling for socio-demographic characteristics, and more importantly to instrumenting the size of the baby-boom with WWII mobilization rates. Lastly, I built a simple repeated matching model, which can generate a sizable divorce boom (both in aggregate divorce rates and across cohorts) from a realistically behaved steady state marriage market.

Overall, this paper adds to the literature on large and surprising consequences of cohort size swings. Future research should examine the micro implications of this aggregate dynamic. In particular, if divorce is motivated by remarriage, the threat of divorce is changing over the life-cycle and across cohorts with changes in the underlying age-structure. This can have predictable implications on labor supply, savings, investment in children, and other important decisions that people typically make within marriage. The importance of remarriage considerations for divorce should also motivate development of new structural marriage market models capable of capturing this mechanism, which are now strikingly missing from the literature. Such contributions could also help reconcile the dynamics of divorce after 1990 that this paper fails to explain.

References

Abramitzky, R., Delavande, A., and Vasconcelos, L. (2011). Marrying Up: The Role of Sex Ratio in Assortative Matching. *American Economic Journal: Applied Economics*, 3(3):124–157.

Acemoglu, D., Autor, D. H., and Lyle, D. (2004). Women, war and wages: The effect of female labor supply on the wage structure at midcentury. *Journal of Political Economy*, 112(3):497–551.

Ananat, E. O. and Michaels, G. (2008). The effect of marital breakup on the income distribution of women with children. *Journal of Human Resources*, 43(3):611–629.

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

Bergstrom, T. and Lam, D. (1991). The Two-Sex Problem and the Marriage Squeeze in an Equilibrium Model of Marriage Markets. *Center for Research on Economic and Social Theory CREST Working Paper*, (91):17.

Bergstrom, T. and Lam, D. (1994). The Effects of Cohort Size on Marriage Markets in Twentieth Century Sweden. *The Family, the Market, and the State in Ageing Societies*, (1987).

Bound, J., Jaeger, D. A., and Baker, R. M. (1995). Problems with instrumental variables estimation when the correlation between the instruments and the endogenous explanatory variable is weak. *Journal of the American Statistical Association*, 90(430):443–450.

Bozon, M. (1991). Women and the Age Gap Between Spouses: An Accepted Domination? *Population: An English Selection*, 3(1991):113–148.

Brainerd, E. (2017). Uncounted Costs of World War II: The Effect of Changing Sex Ratios on Marriage and Fertility of Russian Women. *Review of Economics and Statistics*, (March).

Bronson, M. A. and Mazzocco, M. (2018). Cohort Size and the Marriage Market: Explaining Nearly a Century of Changes in U.S. Marriage Rates. *Working Paper*, pages 1–55.

Browning, M., Chiappori, P.-A., and Weiss, Y. (2011). Family Economics. *Handbook of.*

Bruze, G., Svarer, M., and Weiss, Y. (2015). The dynamics of marriage and divorce. *Journal of Labor Economics*, 33(1):123–170.

Burdett, K., Imai, R., and Wright, R. (2004). Unstable relationships. *Frontiers of Macroeconomics 1.*

Cherlin, A. J. (2004). The deinstitutionalization of American marriage. In *Journal of Marriage and Family*, volume 66, pages 848–861. John Wiley & Sons, Ltd.

Chetty, R., Hendren, N., Kline, P., and Saez, E. (2014). Where is the land of Opportunity? The Geography of Intergenerational Mobility in the United States. *The Quarterly Journal of Economics*, 129(4):1553–1623.

Chiappori, P. A. and Weiss, Y. (2003). Marriage Contracts and Divorce: an Equilibrium Analysis. *Unpublished manuscript. Department of Economics, University of Chicago*, (January):1–35.

Chiappori, P. A. and Weiss, Y. (2007). Divorce, remarriage, and child support. *Journal of Labor Economics*, 25(1):37–74.

Chiappori, P. â., Radchenko, N., and Salanié, B. (2018). Divorce and the duality of marital payoff. *Review of Economics of the Household*, 16(3):833–858.

Choo, E. (2015). Dynamic Marriage Matching: An Empirical Framework. *Econometrica*, 83(4):1373–1423.

Choo, E. and Siow, A. (2006). Who Marries Whom and Why. *Journal of Political Economy*, 114(1):175–201.

Choo, E. and Siow, A. (2007). Lifecycle marriage matching : Theory and evidence. *Working Paper*.

Cornelius, T. J. (2003). A search model of marriage and divorce. *Review of Economic Dynamics*, 6(1):135–155.

De La Croix, D. and Mariani, F. (2015). From Polygyny to Serial Monogamy: A Unified Theory of Marriage Institutions. *The Review of Economic Studies*, 82(2):565–607.

Del Boca, D. (2003). Mothers, fathers and children after divorce: The role of institutions. *Journal of Population Economics*, 16(3):399–422.

Depew, B. and Price, J. (2018). Marriage and the economic status of women with children. *Review of Economics of the Household*, 16(4):1049–1061.

Díaz-Giménez, J. and Giolito, E. (2013). Accounting for the timing of first marriage. *International Economic Review*, 54(1):135–158.

Doepke, M., Hazan, M., and Maoz, Y. D. (2013). The baby boom and world war II: A macroeconomic analysis. *Review of Economic Studies*, 82(3):1031–1073.

England, P. and McClintock, E. A. (2009). The gendered double standard of aging in US marriage markets. *Population and Development Review*, 35(4):797–816.

Friedberg, L. (1998). Did Unilateral Divorce Raise Divorce Rates? Evidence from Panel Data. *American Economic Review*, 88(3).

Garcia-Moran, E. M. (2018). Differential fecundity and child custody. *Journal of Economic Dynamics and Control*, 90:156–170.

Giolito, E. (2004). A Search Model of Marriage with Differential Fecundity. *Working paper*.

Goldin, C. and Katz, L. F. (2016). Changes in Marriage and Divorce as Drivers of Employment and Retirement of Older Women. *NBER Working paper*, (2013):113–156.

Greenwood, J., Guner, N., Kocharkov, G., and Santos, C. (2016). Technology and the changing family: a unified model of marriage, divorce, educational attainment and married female labor-force participation. *American Economic Journal: Macroeconomics*, 8(1):1–41.

Groves, E. R. and Ogburn, W. F. (1928). *Marriage and Family Relations*. Henry Holt & Co, New York.

Johnson, W. R. and Skinner, J. (1986). Labor Supply and Marital Separation. *American Economic Review*, 76(3):455–69.

Kenrick, D. T. and Keefe, R. C. (1992). Age preferences in mates reflect sex differences in human reproductive strategies. *Behavioral and Brain Sciences*, 15(1):75–91.

Kenrick, D. T., Keefe, R. C., Gabrielidis, C., and Cornelius, J. S. (1996). Adolescents' Age Preferences for Dating Partners: Support for an Evolutionary Model of Life-History Strategies. *Child Development*, 67(4):1499–1511.

Lafortune, J. and Low, C. (2017). Tying the Double-Knot: The Role of Assets in Marriage Commitment. *American Economic Review: Papers & Proceedings*, 107(5):163–167.

Low, C. (2017). A 'Reproductive Capital' Model of Marriage Market Matching. *Working paper*, pages 1–82.

Low, H., Meghir, C., Pistaferri, L., and Voena, A. (2018). Marriage, Labor Supply and the Dynamics of the Social Safety Net. *SSRN Electronic Journal*.

McKinnish, T. G. (2007). Sexually Integrated Workplaces and Divorce. *Journal of Human Resources*, XLII(2):331–352.

Moffitt, R. (1992). Incentive effects of the US welfare system: A review. *Journal of Economic Literature*, 30(1):1–61.

Moffitt, R. (1997). The Effect of Welfare on Marriage and Fertility: What Do We Know and What Do We Need to Know? *Institute for Research on Poverty Discussion Papers, University of Wisconsin Institute for Research on Poverty*.

Mortensen, D. T. (1988). Matching: Finding a Partner for Life or Otherwise. *American Journal of Sociology*, 94:S215–S240.

Ni Bhrolcháin, M. (2001). Flexibility in the Marriage Market. *Population*, 13(2):9–47.

Rotz, D. (2016). Why have divorce rates fallen? The role of women's age at marriage. *Journal of Human Resources*, 51(4):961–1002.

Ruggles, S. (1997). The rise of divorce and separation in the United States, 1880-1990. *Demography*, 34(4):455–466.

Ruggles, S., Flood, S., Goeken, R., Grover, J., Meyer, E., Pacas, J., and Sobek, M. (2019). IPUMS USA: Version 9.0 [dataset]. *Minneapolis, MN: IPUMS.*

Shephard, A. (2019). Marriage Market Dynamics, Gender, and the Age Gap. *SSRN Electronic Journal.*

Smock, P., Granda, P., and Hoelter, L. (2015). Integrated Fertility Survey Series, Release 7, 1955-2002 [United States]. *Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor], 2015-06-18.*

Staiger, D. and Stock, J. H. (1997). Instrumental Variables Regression with Weak Instrument. *The Econometric Society*, 65(3):557–586.

Stevenson, B. and Wolfers, J. (2007). Marriage and Divorce: Changes and their Driving Forces. *Journal of Economic Perspectives*, 21(2):27–52.

Weiss, Y. and Willis, R. J. (1997). Match quality, new information, and marital dissolution. *Journal of Labor Economics*, 15(1).

Wolfers, J. (2006). Did Unilateral Divorce Laws Raise Divorce Rates? A Reconciliation and New Results. *American Economic Review*, 96(5):1802–1820.

A Appendix: Supplementary graphs and tables

A.1 Age correction

The data do not allow me to observe all relevant cohorts at the same age to compare on the aggregate which cohorts ultimately divorced the most. To get around that I make use of the fact that I do observe some of the cohorts multiple times and fit a quadratic age gradient in the probability of being ever-divorced. Specifically, I pool IPUMS 1960-1980 Census and ACS $\geq$ 2008 samples, men of $age \in [40, 80)$, and run the following regression:

$$y_{ict} = \alpha_c + \beta age_{ct} + \beta_2 age_{ct}^2 + \epsilon_{ict}$$

Table 5 presents the results. Using β and β_2 I adjust the share of men ever divorced ($s_{c,age=a}$) to age 50

Table 5: Age gradient

Variable	β	β_2
Ever-divorced (among men)	.0184299	-.0001324

Estimating an age gradient, using IPUMS 1960-1980 Census and $\geq$ 2008 ACS.

$$s_{c,50} = s_{c,age=a} + \beta(50 - a) + \beta_2(50^2 - a^2)$$

The identifying variation comes from cohorts (1890-1976) who are observed at least twice. Cohorts 1929-1939 are observed both in 1980 and in later ACS. Figure 8 is equivalent to figure 2a, but without the age correction.

A.2 What is a common age to first divorce?

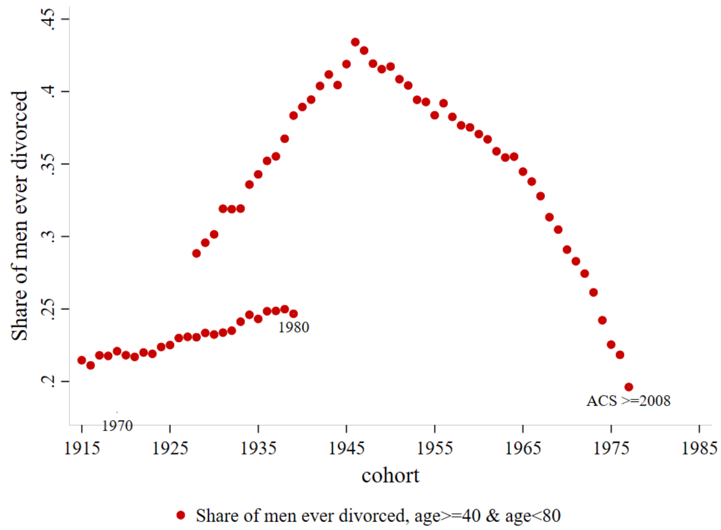
People commonly divorce in their 30s. Figure 9 shows evidence of this from the NSFG. Similarly, Goldin and Katz (2016) shows a mean age at divorce for women around 34, and decreasing to 30 with later cohorts, for women observed 50-74 in the SIPP.

A.3 Supplementary cohort evidence

Especially compared to the downward trend in the age-gap (husband minus wife), men in the pre-boom (and also in the 1930-35 cohorts) lived in marriages with bigger age-gaps on average (see figure 10). This is consistent with the idea that men in these generations (re)matched with younger (second) wifes.[26]

[26] Analysis by the type of marriage (first versus higher order) shows that this is a composition effect of more remarriage, rather than even first marriages happening more often with larger age-gaps. Since the age-gap in husband's second marriage tends to be higher than in their first marriage, an increase in the share of higher order marriages implies an increase in the age-gap.

Figure 8: Share ever divorced

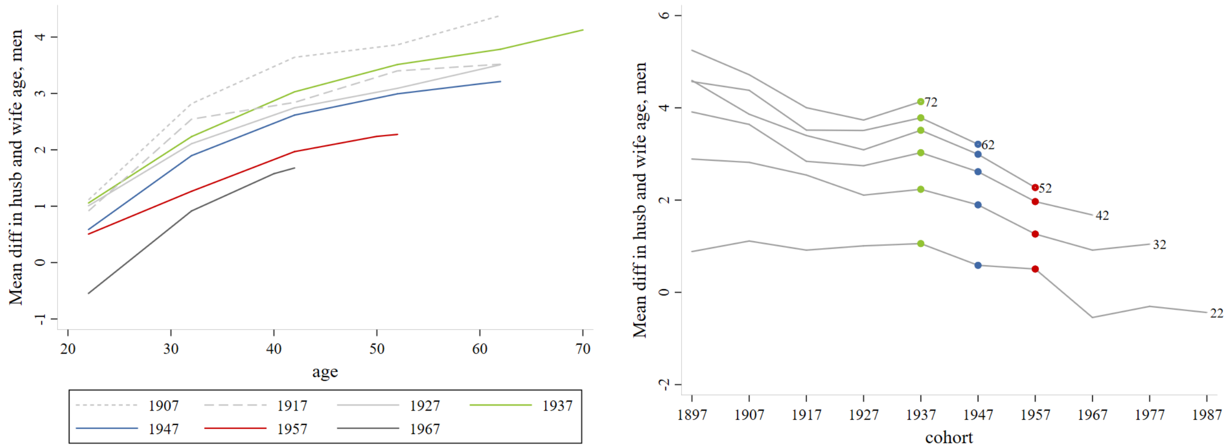


2a without an age correction.

Figure 9: Age at divorce

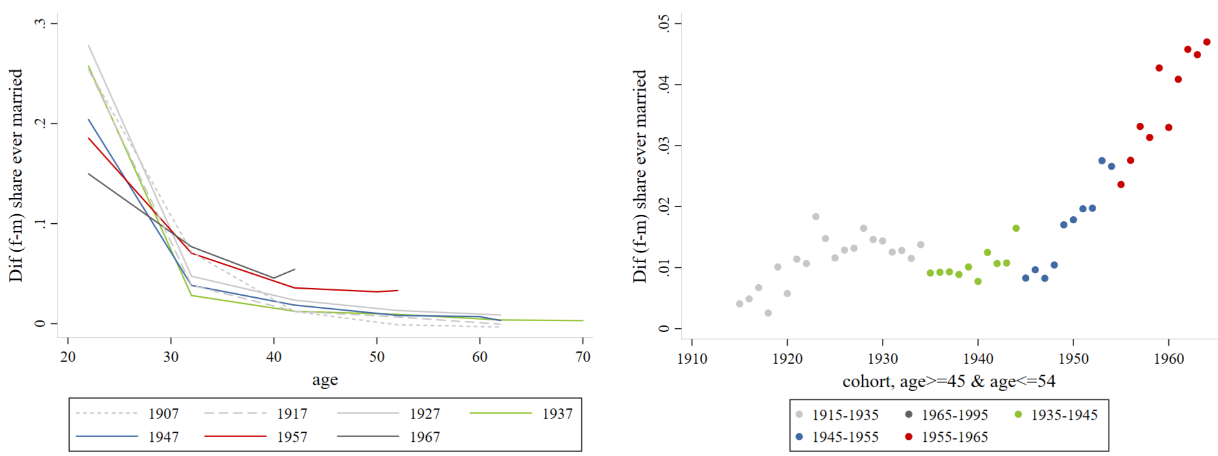
Source: NSFG, women 40-45, ever divorced, age at wife's first divorce.

Figure 10: Age-gap (husband minus wife) for men by age for selected cohorts/ by cohort for selected ages.



Among the 'treated' cohorts, men almost caught up with women in whether they ever got married by midage, especially compared to a more downward trend in marriage among men compared to women (see figure 11). This is consistent with the story, as this cohort of men faced an exceptionally large pool of eligible women and a low competition from older men.

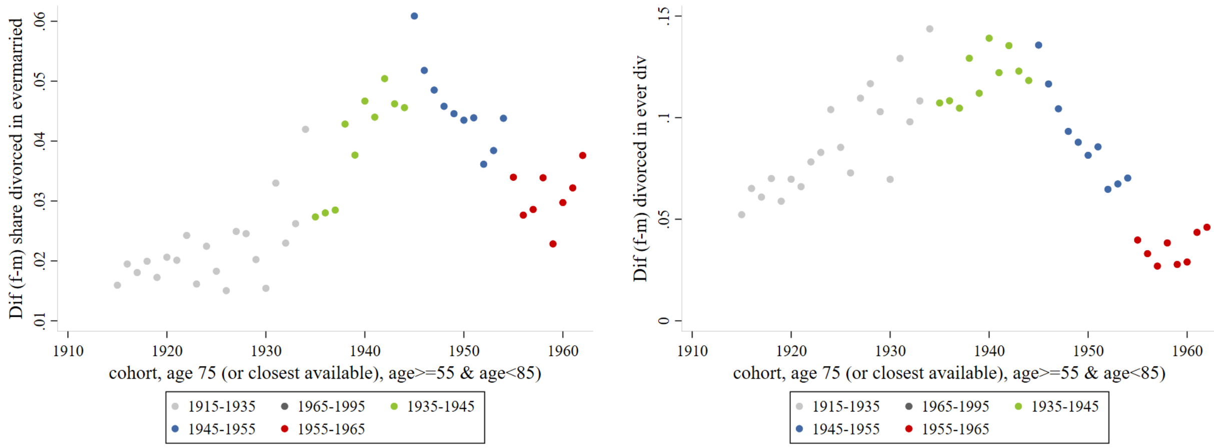
Figure 11: Gender difference in the share of ever-married.



Compared to other cohorts, women in the treated cohorts were especially likely to stay divorced (not remarry) compared to men (see figure 12).[27] This is consistent with the hypothesis that the primary cause for divorce was an increase in the remarriage options of the husband.

[27]Note - this is actually true for the early-boomers as well.

Figure 12: Gender difference in the share of ever-married/ ever-divorced who live as divorced by cohort for selected ages.

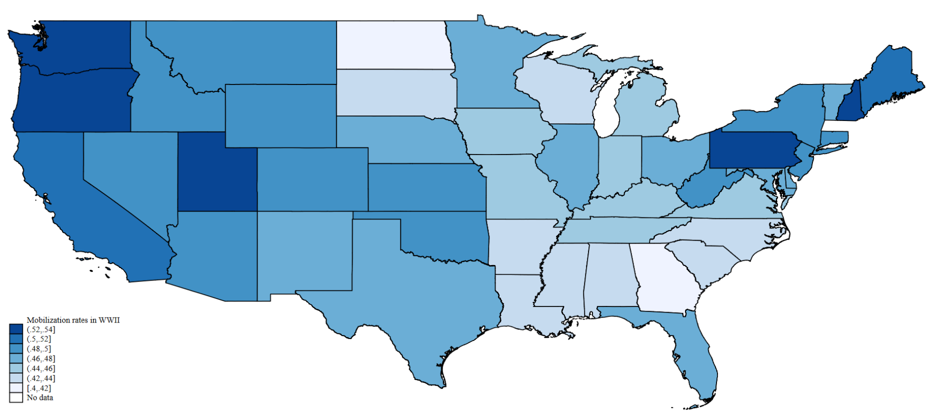


Especially among the 'treated' cohorts, women stayed divorced (single) more than men (even though both divorced a lot).

A.4 Supplementary analysis of variation in mobilization rates

Mobilization rate is defined as the fraction of registered men between the ages of 18 and 44 who were drafted or enlisted for WWII, as generously provided by Acemoglu et al. (2004).[28] Figure 13 shows the geographic distribution of mobilization rate. Table 6 shows descriptive characteristics of men and women aged 35-44 in high and low mobilization rate states.

Figure 13: Mobilization rate in WWII (Acemoglu et al., 2004).



There are baseline differences (even before the WWII) between these states in race, education and rate of

[28]Original source: published tables of the Selective Service System (1956). Since essentially all men in the relevant age range were registered, mobilization rate is effectively the fraction of men in this age range who have served (Acemoglu et al., 2004).

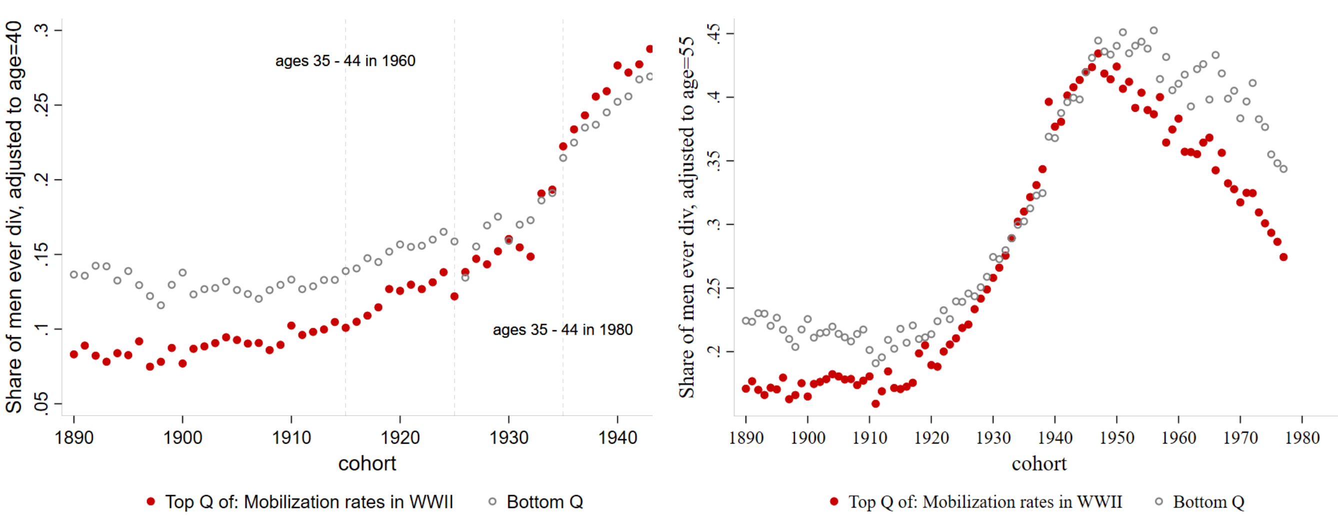
Table 6: Descriptive statistics of high and low mobilization states respectively, men and women 35-44 years old.

Cohorts	1940 (1895-1904) $> p50$	1940 (1895-1904) $\leq p50$	1960 (1915-1924) $> p50$	1960 (1915-1924) $\leq p50$	1980 (1935-1944) $> p50$	1980 (1935-1944) $\leq p50$
White	0.97	0.83	0.96	0.85	0.94	0.83
Graduated high-shool	0.31	0.24	0.59	0.47	0.83	0.75
Graduated college	0.07	0.05	0.10	0.08	0.22	0.17
Farm	0.11	0.28	0.05	0.12	0.02	0.04
Urban	0.65	0.50	.	.	.	.
In Metro area	0.69	0.44	0.68	0.54	0.74	0.64
Ever-married	0.88	0.90	0.93	0.94	0.93	0.94
Ever-divorced	.	.	0.14	0.17	0.26	0.27

urbanization, as already discussed in Acemoglu et al. (2004). At least race and farm explicitly linked to the mobilization rules.

Figure 14 shows that starting from a lower level high mobilization states had a bigger boom in divorces among 1925-1955 cohorts. This graph looks very similar when splitting by high and low $\Delta_{1960-1980}\frac{n_{30-34}}{n_{30-44}}$,

Figure 14: Share of ever-divorced, adjusted to age 40/50, top and bottom quartiles in terms of WWII mobilization rates. Mobilization rates assigned by state of birth.



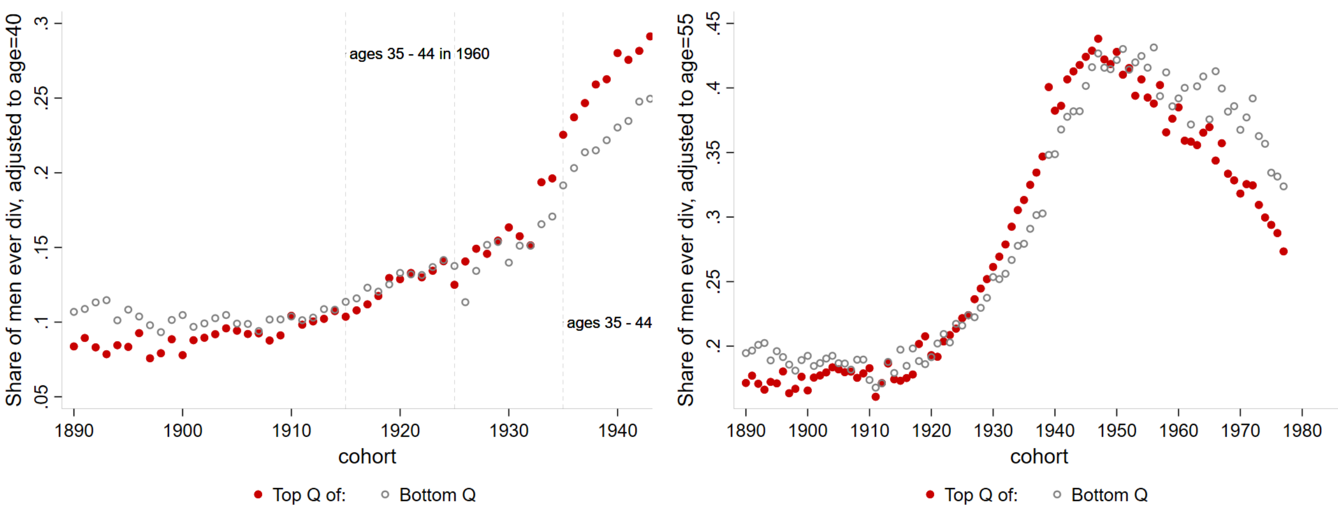
or when plotting share ever divorced in ever married.

The differences in the baseline can be largely accounted for by differences in race and education. Using the IPUMS 1960 Census data and ages 35-80 (cohorts 1890-1924, i.e. cohorts that should not be affected by the baby-boom yet), I fit the following regression:

$$y_{ist} = \alpha_s + \alpha_{1980} + \alpha_{age} + 0.11\ Black\ - 0.01\ Hispanic \\ - 0.01\ Graduated\ Highschool\ - 0.04\ Graduated\ college\ + \epsilon_{ist}$$

I then use the coefficients on race and education to residualize the (age adjusted) share of men ever-divorced in all cohorts 1890-1980. Figure 15 shows that even given this correction, men born in states with a higher mobilization rate had a bigger increase in the chance of getting divorces (now starting from a comparable baseline).

Figure 15: Share of ever-divorced, adjusted to age 40/50 and for baseline differences in race and education, top and bottom quartiles in terms of WWII mobilization rates. Mobilization rates assigned by state of birth.



Motivated by the baseline differences presented in table 6, I control for dummies for race, education and farm status (as detailed as available in IPUMS) in my regressions.

B Appendix: Does the effect of mobilization rates come from women's labor force participation or fertility?

The variation in mobilization rates has been shown to increase labor force participation of women during (and a bit after WWII) and decrease labor force participation and increase fertility of young women after the war by Acemoglu et al. (2004) and Doepke et al. (2013) respectively. This previous research motivates the use of mobilization rates as an instrument for the size of the baby-boom in this paper. However, a natural question is whether these previously documented mechanisms could not affect divorces directly, not just through shifting the future age distribution.

First, one might suspect that an increase in labor force participation of women might make couples more likely to divorce, as women gain work experience and their value outside of marriage increases (as suggested by Ruggles (1997), Weiss and Willis (1997)), thus implying a correlation between the divorce boom and mobilization rates in WWII. This alternative hypothesis is however not consistent with the incidence of WWII effects on labor force participation vs the divorce increase across cohorts. Figure 16 visualizes that it was mainly cohorts of women born before 1925 who were shown to have higher work attachment in high mobilization states by Acemoglu et al. (2004). These cohorts are not the ones responsible for the divorce boom. Moreover, in the baseline cross-sectional analysis these cohorts in fact serve as a control group to the treated cohorts of 1935-1944 (the two groups are indicated by dashed lines in the figure). As a result, if higher labor force participation by women causes divorce, the baseline results presented in table 2 should represent a lower bound.[29]

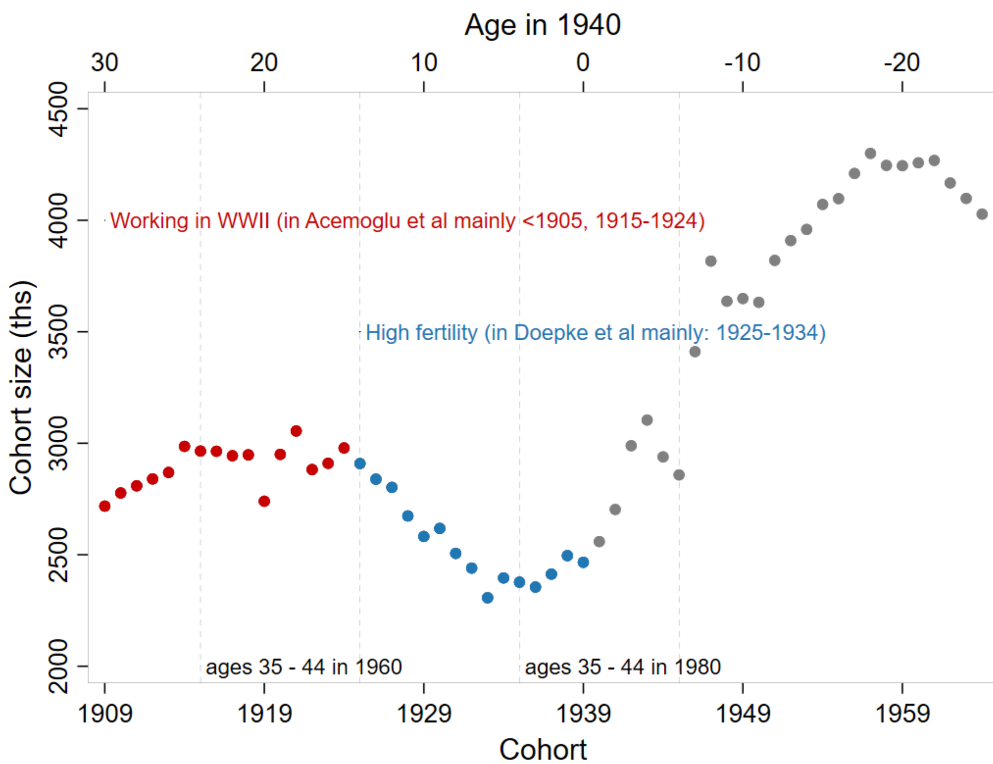
Figure 16 also shows that the cohorts having a higher fertility compared to previous cohorts in high mobilization states overlap partially with the treated cohorts who started the divorce boom. In other words, the cohorts of 1935-1939 were both mothers to the late baby-boomers and were themselves affected by the early baby-boomers entering the marriage market. A competing hypothesis could therefore be that couples in high mobilization states started divorcing more because women in these couples had more children and less work attachment, making them more vulnerable to divorce. To test this mechanism, I repeat the baseline analysis, but compare the treated cohorts of 1935-1944 to the cohorts 1925-1934, whose fertility was also high. Specifically, I repeat the analysis outlined in section 3 with a pooled sample of 1970 and 1980 IPUMS Census data,i.e. $t \in \{1970, 1980\}$ in

$$y_{ist} = \alpha_s + \alpha_{1980} + \alpha_{age} + \beta \frac{n_{30-34}}{n_{30-44}}_{st} + \gamma X_{ist} + \epsilon_{ist}$$

If the effect of mobilization rates on divorce is coming from a greater incidence of vulnerable stay-at-home women, comparing two cohorts of high fertility should lead null results. Table 7 shows that the results are

[29] As a further check, also rerun the analysis presented in table2 for women only and include dummies for labor force participation and (or) employment and confirm that the results are essentially unaffected.

Figure 16: Effects of mobilization rates across cohorts on female labor force participation and fertility, among cohorts used as treated and control in the baseline analysis.



in fact qualitatively the same as in table 2. Quantitatively, these results are smaller. This is likely caused by the fact that the cohorts of 1925-1934 and 1935-1944 are close enough so that the entry of baby-boomers affected both of them, yet the effect was bigger on cohorts 1935-1944.[30]

Table 7: Baseline with broader sample.

	Ever-divorced					
	OLS		2SLS			
$\frac{n_{30-34}}{n_{30-44}}{}_{st}$	0.466	0.477	0.884	0.951	1.133	0.916
	(2.99)	(2.92)	(3.83)	(3.28)	(3.15)	(3.08)
X_i:						
farm, metro	no	yes	no	yes	yes	yes
educ dummies	no	yes	no	yes	yes	yes
region-year fes	no	no	no	no	yes	no
In ever-married	no	no	no	no	no	yes
F-stat in 1st stage			56.6	56.6	40.3	56.1

N = 1222220 (1143381 in column 5), 48 clusters

t statistics in parentheses. SEs clustered at the state level.

All regressions include year and state fixed effects, and dummies for age, sex, race.

Same setup as table 2 except using pooled 1970 and 1980 IPUMS Census data, i.e. measuring growth in divorces between cohorts 1925-1934 and 1935-1944.

[30]Moreover, I confirm that including age at marriage/ number of children ever born/ number of young children in the household dummies as controls does not diminish the baseline results in table 2.

B.1 Supplementary tables for cross-sectional evidence

First stage This section presents evidence on the strength of the first stage in my IV strategy. Following Doepke et al. (2013) I show that states with higher mobilization rates during WWII has a sharper cohort size growth after WWII, materializing in a sharper increase in the remarriage opportunities in the 1970s for men in their 30s. Specifically, I regress the remarriage opportunity measure, $\frac{n_{30-34}}{n_{30-44}}_{st}$ or $\frac{n_{0-4}}{n_{0-14}}_{s,t-30}$ (notice t still stands for a year 1960 or 1980) with the WWII mobilization rates interacted with a dummy for year 1980, including state and year fixed effects as well as individual level controls used in the main analysis presented in table 2.[31] This specification is a panel fixed-effects version of a regression predicting the change in the remarriage opportunity measure with WWII mobilization rates. Table 8 shows that this strategy has a strong first stage (with F statistics on mobilization rates above 15), with perhaps the exception of the specification regressing $\frac{n_{0-4}}{n_{0-14}}_{s,t-30}$ and including region times year fixed effects, where the F statistic falls slightly short of the rule of thumb of 10 (as suggested by Staiger and Stock (1997) to avoid weak instruments issues as brought to light by Bound et al. (1995)).

Table 8: First stage results, as relevant for the 2SLS results in table 2 and 9.

$$u_{st} = \alpha_s + \alpha_{1980} + \alpha_{age} + \delta mob \cdot 1_{t=1980} + \gamma X_{ist} + \epsilon_{ist}$$

u_{st}	$\frac{n_{30-34}}{n_{30-44}}_{st}$			$\frac{n_{0-4}}{n_{0-14}}_{s,t-30}$		
δ	0.214	0.214	0.202	0.287	0.283	0.211
F statistic	25.60	25.70	31.13	15.36	15.05	8.76
X_i:						
sex, race	yes	yes	yes	yes	yes	yes
farm, metro	no	yes	yes	no	yes	yes
educ dummies	no	yes	yes	no	yes	yes
region-year fes	no	no	yes	no	no	yes

$N = 2032220$, 48 clusters

SEs clustered at the state level.

All regressions include year, age and state fixed effects.

[31] These controls take out variation in the remarriage opportunity measure correlated with state level compositional changes in sociodemographic variables

Proxy cohort size by 1980 with cohort size among children by 1950 To confirm that this results is not driven by selective migration in adulthood I reconstruct the measure $\frac{n_{30-34}}{n_{30-44}}_{st}$ in 1950 and 1930 Census as $\frac{n_{0-4}}{n_{0-14}}_{s,t-30}$ and run the following equivalent regression:

$$y_{ist} = \alpha_s + \alpha_{1980} + \beta \frac{n_{0-4}}{n_{0-14}}_{s,t-30} + \gamma X_{ist} + \epsilon_{ist}$$

Table 9 shows the results.

Table 9: Baseline with baby-boom size measured at ages 0-14.

			Ever-divorced			
	OLS		2SLS			
$\frac{n_{0-4}}{n_{0-14}}_{s,t-30}$	0.666	0.832	1.109	1.564	1.836	1.085
	(5.75)	(6.39)	(4.58)	(5.20)	(3.56)	(4.49)
X_i:						
sex, race	yes	yes	yes	yes	yes	yes
farm, metro	no	yes	no	yes	yes	yes
educ dummies	no	yes	no	yes	yes	yes
region-year fes	no	no	no	no	yes	no
In ever-married	no	no	no	no	no	yes

N = 2032220 (1902899 in column 5), 48 clusters

t statistics in parentheses. SEs clustered at the state level.

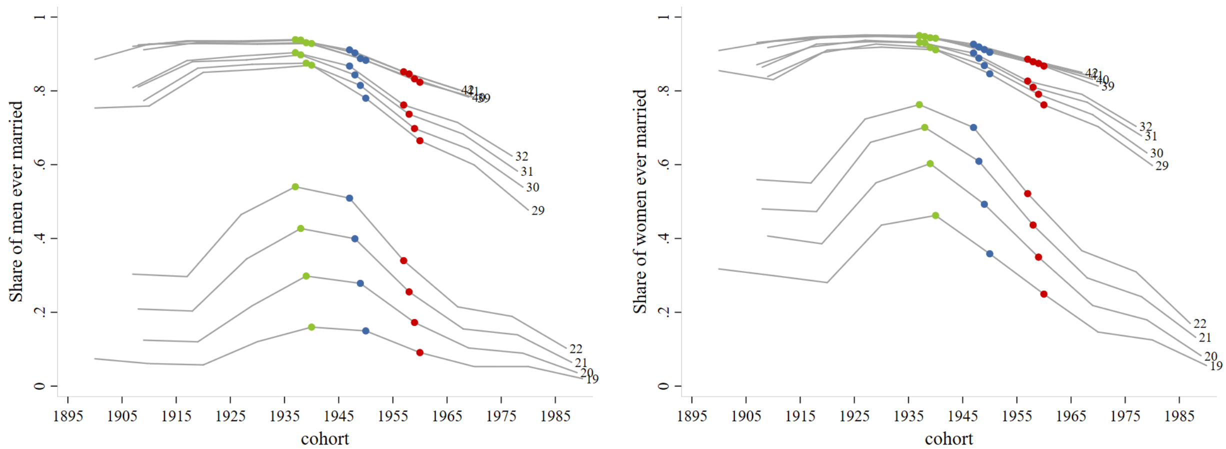
All regressions include year, age and state fixed effects.

Same as table 2 except $\frac{n_{30-34}}{n_{30-44}}_{st}$ is replaced with $\frac{n_{0-4}}{n_{0-14}}_{t-30}$ (constructed from 1930 and 1950 IPUMS Census data).

The results are broadly similar to table 2, though slightly smaller in magnitude.

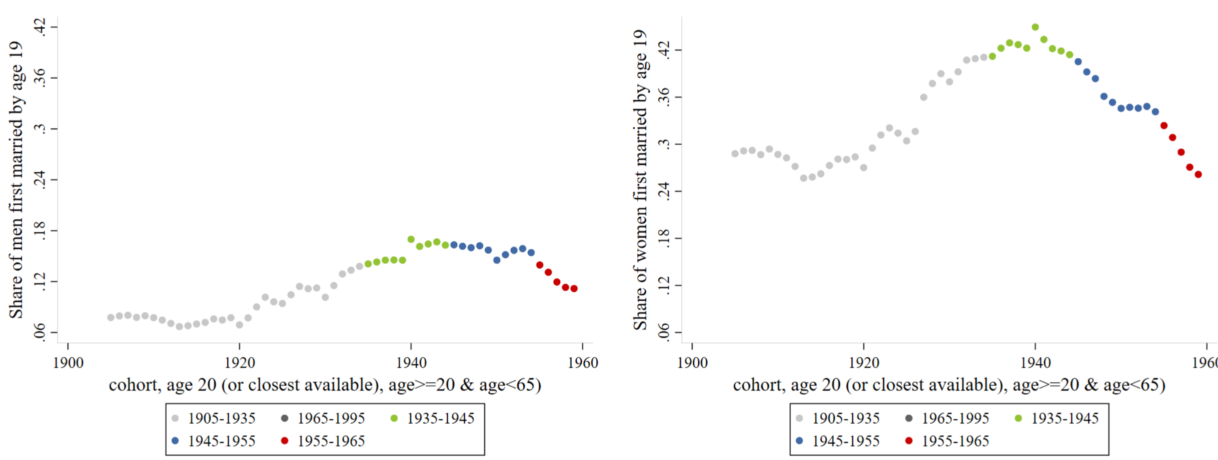
C Appendix: What is the role of age at marriage?

Figure 17: Share of men and women ever-married in selected by age plotted for selected available ages.



Data source: all IPUMS Census and ACS.

Figure 18: Share of men/ women ever married by age 19 (using the reported age at marriage) by cohort.

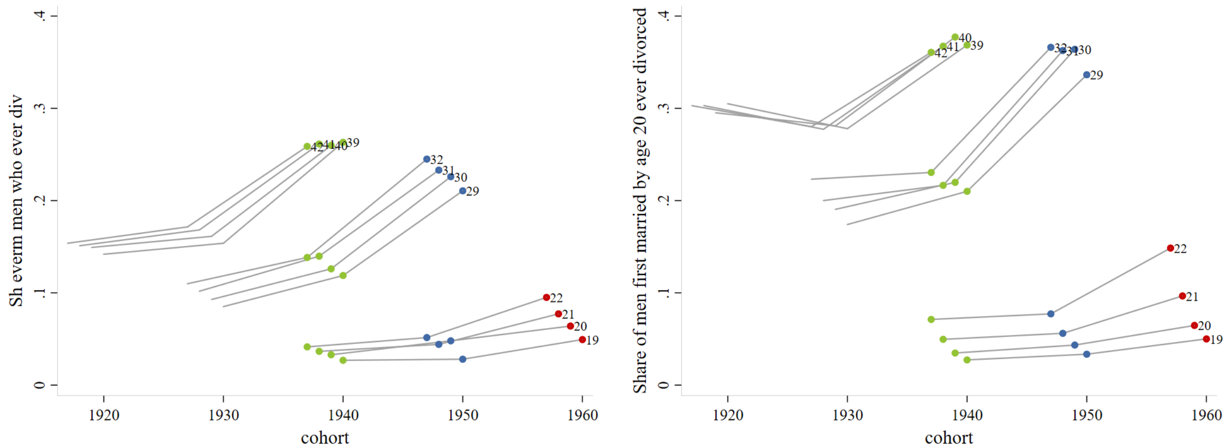


Data source: IPUMS Census, 1960-1980.

We know from Rotz (2016) (and others) that low age at marriage is strongly predictive of divorce. This is also visible in figure 19 (between 1960 and 1980 the Census asked both about age at first marriage and the order of current marriage, making it possible to analyze divorce probabilities by age at marriage).

This could suggest that the divorce boom might only have been a result of low age at marriage. However,

Figure 19: Share of men who ever divorced, among those who ever-married/ among those who married by age 20.



including age at marriage as a control in my cross-sectional regressions does not diminish the result that pre-boom generations divorced more when faced with a large baby-boom cohort. I suspect that low age at marriage among the pre-baby-boom (and somewhat extending to the early baby-boom) generations was endogenous to the cohort size dynamics. The pre-baby-boom cohorts (especially cohorts born during 1930s) were very small (there was a baby-bust during the great depression). This creates a shortage of young women in the marriage market, possibly pushing men not to postpone proposing to lock in a match. If this mechanism is persistent (as young women are matching at earlier ages, their peers also perceive a shortage of partners), it could also explain the low age at marriage extending to the cohorts born after 1935. However, including this mechanism is beyond the scope of this paper.

Table 10: Baseline, controlling for age at marriage.

	Ever-divorced				
	OLS		2SLS		
$\frac{n_{30-34}}{n_{30-44}}st$	0.569	0.664	1.145	1.647	1.366
	(2.89)	(2.64)	(3.45)	(4.38)	(3.27)
X_i:					
sex, race	yes	yes	yes	yes	yes
farm, metro	no	yes	no	yes	yes
educ dummies	no	yes	no	yes	yes
region-year fes	no	no	no	no	yes

N = 2032220 (1902899 in column 5), 48 clusters

t statistics in parentheses. SEs clustered at the state level.

All regressions include dummies for age at marriage, bottom-coded at 15

All regressions include year, age and state fixed effects.

Same setup as table 2 except including age at marriage dummies (bottom coded at 15).

D Appendix: Why did early baby-boom cohorts divorce even more than the 1935-1944 cohorts?

My hypothesis is that a combination of two factors, marrying to divorcees and establishing more marriages with a small age gap, caused the 1945-1954 cohort to divorce even more than the 1935-1944, even though their remarriage options were worse. Both of these are endogenous responses to the cohort size movement. As the 1945-1954 are much bigger than preceding cohorts, women naturally face a shortage of marriage partners (with a standard age-gap in marriage). Bergstrom and Lam (1991) and others show that a way the marriage market adjusts is by increasing the number of partnerships among peers. These marriages, however, are more at risk of divorce due to rematching. This is because a small age-gap implies a bigger pool of women that are younger than the existing wife. I propose that a second adjustment comes in form of more matches with divorcees (which have also been shown to be less stable). Together these imply that large cohorts enter more in inherently unstable marriages.

Table 11 and figure 22 confirms that controlling for age at first marriage[32], an age-gap that is lower (or

[32] Without controlling for age at first marriage a small gap is associated with delayed marriage, so the positive effect on

Figure 20: Age-gap in existing marriages by age.

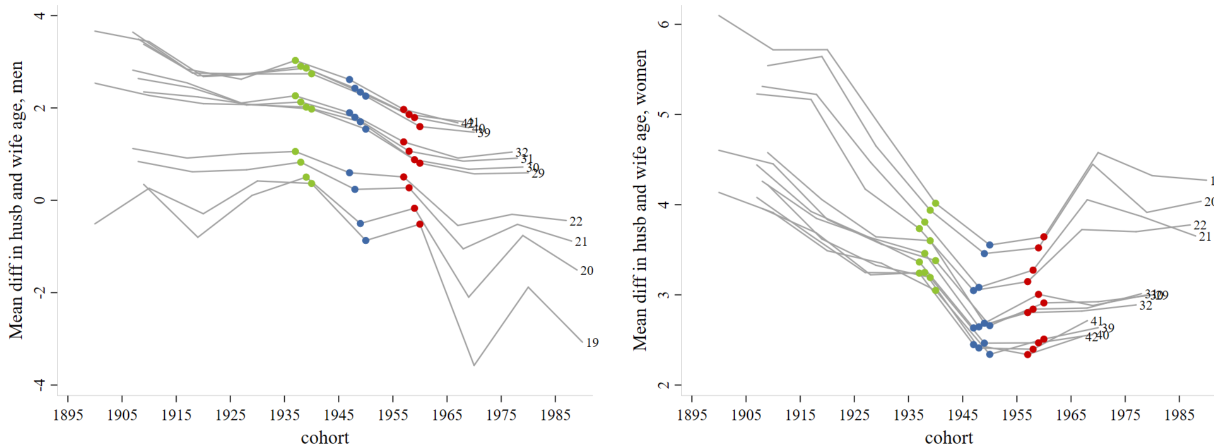
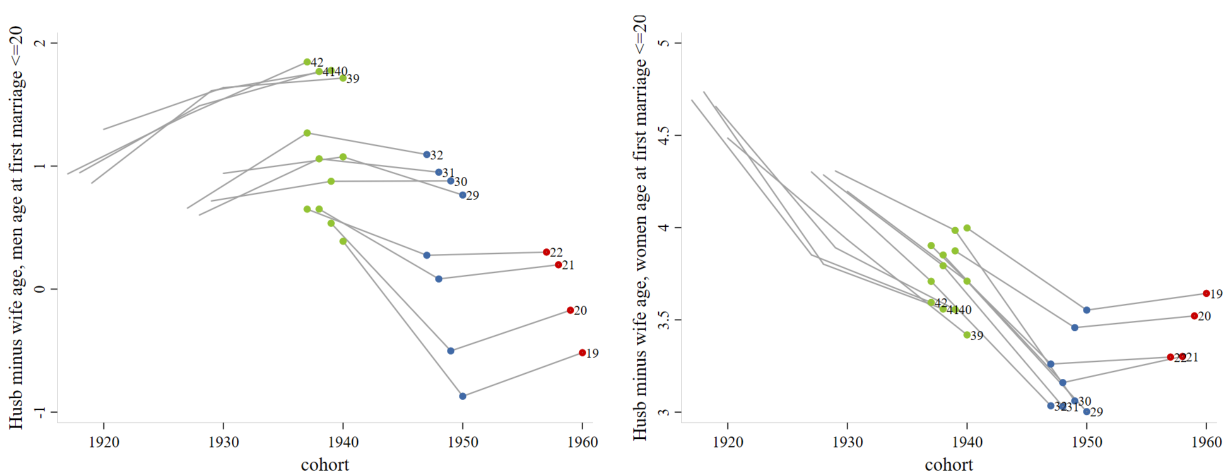


Figure 21: Age-gap in existing marriages by age, for a group of men/ women who were first married by age 20.



This is a consistently defined group across sample. The sharp increase in the age gap can only happen for two reasons. First, the marriages with small age-gaps broke. Second, some of the marriages at young age broke and rematched with a much younger partner. Both are consistent with the mechanisms studied in this paper. This shows that the increase in age-gap with age is not caused only by selection into first marriage.

much higher) than standard is correlated with a higher chance of divorce (using marital histories of cohorts 1920-1970, from the public use versions of the Growth of American Families, National Fertility Surveys and National Surveys of Family Growth as compiled by the Integrated Fertility Survey Serie, (Smock et al., 2015).

Table 11: Ever-divorced based on age at first marriage

		Ever-divorced
		Age at 1st marriage < 20
Age at first marriage	-0.0327***	-0.0623***
	(-25.18)	(-10.38)
Age-gap in 1st mar >= 10	0.165***	0.231***
	(8.75)	(7.48)
Age-gap in 1st mar $[5, 10)$	0.0354***	-0.00225
	(2.69)	(-0.11)
Age-gap in 1st mar $[1, 5)$	0	0
Age-gap in 1st mar < 1	0.0545***	0.0558**
	(4.46)	(2.25)
Observations	51164	23710

t statistics in parentheses. SEs clustered at the state level.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

IFSS (1955 and 1960 GAF, 1965 and 1970 NFS,

1973 1976, 1982, 1988, 1995 and 2002 NSFG waves)

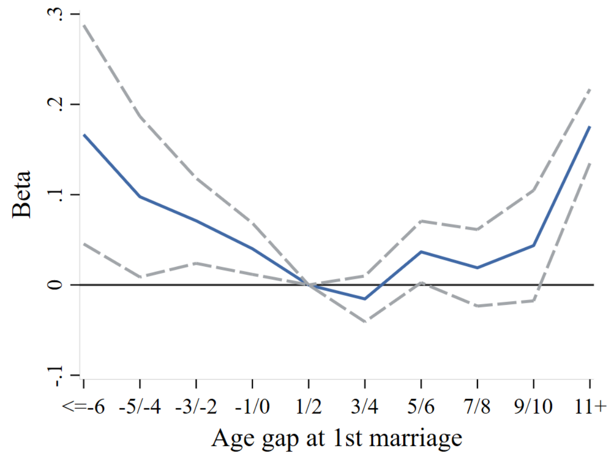
cohorts $1920 - 1970$, age 20-45, ever marrried women, X: cohort and age dummies

Source of data: GAF, NSF and NSFG refer to public use versions of the Growth of American Families, National Fertility Surveys and National Surveys of Family Growth as compiled by the Integrated Fertility Survey Serie, (Smock et al., 2015)

This suggests that 1980 was special for two reasons: first green cohorts divorced as blue cohorts snitched their husbands. Second, blue marriages that happened at a young age and among peers or with a reversed

divorce is not as apparent.

Figure 22: Ever-divorced by age-gap at first marriage



Regressing a dummy of ever being divorced on age at marriage and dummies for age-gap ar first marriage, plotting coefficients on age-gap dummies (1-2 year age-gap being the base). Source of data: GAF, NSF and NSFG refer to public use versions of the Growth of American Families, National Fertility Surveys and National Surveys of Family Growth as compiled by the Integrated Fertility Survey Serie, (Smock et al., 2015)

age gap broke (with a subsequent remarriage or just because value of singlehood of the men went up as the 1955-1964 cohorts were entering the marriageable age).

E Appendix: Model solution

Given the assumptions on idiosyncratic preferences, the within period matching problem is almost fully equivalent to the static model in Choo and Siow (2006), who show that given supplies of men and women of each type, the matches solve a simple system of equations.

Corollary 1. *Given m_{it} and f_{jt}, every period the matching outcomes $n_{ijt}, n_{i\emptyset t}, n_{\emptyset jt}$ satisfy*

$$e^{\frac{\alpha_{ij}-\alpha_{i\emptyset}+\gamma_{ij}-\gamma_{\emptyset j}}{2}} = \frac{n_{ijt}}{\sqrt{n_{i\emptyset t}n_{\emptyset jt}}}$$

and

$$ln\left(\frac{n_{ijt}}{n_{i\emptyset t}}\right) = \alpha_{ij} - \alpha_{i\emptyset} - \tau_{ijt},\ ln\left(\frac{n_{ijt}}{n_{\emptyset jt}}\right) = \gamma_{ij} - \gamma_{j\emptyset} + \tau_{ijt}$$

The assumptions of Choo and Siow (2006) are violated, because of the exit of widows and widowers from the marriage market. As a consequence, the distribution of errors in the market is not exactly iid extreme value type I, because widows and widowers are not a random sample of the population. However, numerically the transfers given by $\tau_{ijt} = -ln\left(\frac{n_{ijt}}{n_{i\emptyset t}}\right) + \alpha_{ij}$ with the solution for $\{n_{i,j,t}\}$ as specified in 1 clear the simulated market almost exactly.