(Changing) Marriage and Cohabitation Patterns in the US: do Divorce Laws Matter?*

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Abstract

What is the role of unilateral divorce in the rise of unmarried cohabitation? Exploiting the staggered introduction of unilateral divorce across the US states, we show that after the reform singles become more likely to cohabit than to marry, and that newly formed cohabitations last longer. To understand the mechanisms driving these outcomes, we build a life-cycle model with partnership choice, endogenous divorce/breakup, female labor force participation, and saving decisions. A structural estimation that matches the empirical findings suggests that unilateral divorce decreases the marriage gains that derive from cooperation and risk-sharing. This makes cohabitation preferred among couples that would have likely faced a divorce, which is more expensive than breaking up. As cohabiting couples formed after the reform are better matched, the average length of cohabitations increases by 27%. Consistent with data, the rise in cohabitation is larger in states that impose an equal division of property upon divorce. This is because men, who stand to lose more wealth in a divorce than in a breakup, convince women to cohabit in exchange for more household resources. A counterfactual experiment reveals that the time spent cohabiting would have been halved if the divorce laws had never changed.

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1 Introduction

Unmarried cohabitation is on the rise: the share of women that ever cohabited in the United States moved from 33% in 1987 to 60% in 2010 (Manning, 2013). This increase contributed to the overall changes in the structure and behavior of the American family. The higher instability of cohabitation contributes to the rise in the number of single mothers (Bumpass and Lu, 2000), which is associated with poor outcomes for children (Chetty and Hendren, 2018; McLanahan et al., 2013). Cohabiting people are less likely to engage in relationship-specific investments, such as intra-household specialization or having joint accounts. (Poortman and Mills, 2012). Finally, their children's well-being is worse even after controlling for parental resources (Brown, 2004). However, it is not clear to what extent cohabitants' outcomes are due to selection versus the direct effect of the form of partnership on the couple's behavior. Quantifying the relative importance of these two mechanisms is only possible if we understand the rationale for the partnerships choices. Why do people cohabit instead of marrying? Why is cohabitation on the rise?

Our paper addresses these questions by focusing on a major US policy change that took place mostly during the 1970s. During this period, most of the states made divorce easier by switching from the mutual consent divorce, requiring both spouses to agree to divorce, towards unilateral divorce, in which one spouse's decision was enough to initiate the procedure. The paper explores the role of unilateral divorce in the rise of cohabitation. Since marriage and cohabitation can be viewed as contracts whose attractiveness depends on their termination rights and costs, the switch from mutual consent to unilateral divorce offers a unique opportunity to learn about partnership choices. Understanding these choices is relevant for policy design: for example, protecting the weaker partner by increasing her/his rights within marriage can backfire if it causes the couple to choose a less protective partnership, such as cohabitation.

We answer the question with four contributions. First, we show that after the reform singles become more likely to cohabit than to marry, and newly formed cohabitations last longer. Second, we propose a theory of partnership choice and endogenous breakup/divorce to understand the mechanisms underlying these facts. Third, we quantify the importance of each mechanism by structurally estimating our model to match the empirical findings about the transition of the divorce regimes. Fourth, we perform several counterfactual experiments to understand the role of unilateral divorce in the rise of cohabitation and in changes in the pool of cohabiting couples. Thus, this paper consists of four parts, one for each contribution, which we now describe in more detail.

In the first part of the paper we document the effect of unilateral divorce on the choice between marriage and cohabitation and the duration of newly formed cohabitations. We use data from the National Survey of Family and the Household (NSFH) and from the National Survey of Family Growth (NSFG) to study the choice between marrying and cohabiting. Then, exploiting the exogenous variation coming from the staggered introduction of unilateral divorce over time across the US states, we estimate that couples formed after the policy change are 7-8% less likely to choose marriage over cohabitation than in the pre-reform period. Among the couples formed in the year the law changed, 30% chose cohabitation. Interestingly, the size of the effect depends on how property is divided upon divorce, being strongest in states where each spouse gets half of the wealth and where the judge decides the allocation of assets. This suggests that divorce settlements affect partnership choices when one spouse can divorce unilaterally. Moreover, we analyze how unilateral divorce affected the duration of cohabitation spells: our estimates show that cohabitations formed after the reform last longer because people both marry less and break up less.[1]

In the second part of the paper, we propose a theory to understand the mechanisms underlying the facts documented in the empirical part. We build a dynamic collective model of intra-household decision making and search in the mating market, where agents make decisions according to the realization of idiosyncratic permanent income shocks, their amount of wealth and couple-specific match quality. With some probability, single agents meet a potential partner drawn from an exogenous distribution of match quality, productivity, and wealth. After the draw, they decide whether to marry, cohabit, or stay single. Couples make decisions about consumption, savings, and female labor force participation. Women experience a productivity penalty for not working, and women's time can be used to produce a public good that captures utility gains from children, durable goods, and services.

In the model, cohabitation and marriage differ in their splitting costs and the way property is divided when the partnership dissolves. Moreover, there is a stigma affecting cohabitations, which is modeled as an exogenous disutility flow. It captures the negative judgment towards out-out-wedlock births and premarital sex. In the case of a breakup, assets are split according to individual property rights. In the case of divorce, we assume they are divided in half.[2] Additionally, we assume that unlike a breakup, divorce is financially costly. Breakup (unilateral divorce) can be initiated unilaterally, as opposed to mutual consent divorce, which requires both partners' agreement. Following Voena (2015), under mutual consent the couple always co-

[1] Hereafter we refer to the separation from cohabitation as a breakup to avoid confusion with legal separation.

[2] We estimate the model using community property states data to be consistent with this assumption.

operates while married and the allocation of resources corresponds to the Pareto-efficient intertemporal allocation. When just one spouse (cohabitant) can decide to terminate the relationship this causes a lack of commitment,[3] making the intra-household decision power responsive to shocks, as spouses (cohabitants) can credibly exercise the threat of divorce (breakup). Because utility is imperfectly transferable, in our model the Becker-Coase theorem does not hold. Hence, abandoning the mutual consent regime affects the risk of divorce and, in turn, the surplus of marriage.[4]

The barriers to divorce—represented by its costs and the right to veto it—affect the gains of marriage relative to cohabitation through three main channels. First, by acting as commitment technologies, they enforce a better risk sharing and a more efficient household specialization. Second, they increase the risk of being "trapped" in a bad marriage that provides low utility. Since this risk is larger for couples with a low match quality, these couples prefer to cohabit since a breakup is cheaper and easier to obtain. Third, they affect the expected value of marriage by modifying the risk of divorce, as the intra-household allocation during marriage differs from the allocation of resources in divorce, which depends on the mandated equal division of property. The effects of tightening or relaxing the barriers to divorce depend on which channel prevails. For example, the introduction of unilateral divorce has an uncertain impact on the share of couples that cohabit and marry. The outcomes of cohabitation and marriage depend not only on the rules underlying these contracts but also on the match quality through its effect on the couples' stability. In fact, making partnership-specific investments is more comfortable when the risk of splitting is low, which is when the match quality is high. Since a cheap breakup is most attractive to couples whose match quality is low, selection on match quality amplifies the differences in the behavior between married and cohabiting couples.

In the third part of the paper we do a structural estimation of the model to understand the quantitative relevance of the mechanisms that drive partnership choice. The model is estimated by indirect inference using as targets the regression results from our empirical analysis, mating market moments (NSFH), and female labor force participation moments (PSID). The introduction of unilateral divorce is modeled as an unexpected policy change. The estimated model

[3] Our modelling of the decision making in the couple builds on existing literature on limited commitment (Kocherlakota (1996), Ligon et al. (2002), Marcet and Marimon (2019) and Pavoni et al. (2018)), which has been applied to dynamic collective models in the household by Mazzocco (2007), Mazzocco et al. (2013), Bayot and Voena (2015), Oikonomou and Siegel (2015), Voena (2015), Ábrahám and Laczó (2018), Lise and Yamada (2018), Low et al. (2018), Foerster (2020) and Reynoso (2020) among others.

[4] According to Becker et al. (1977), divorce laws should not affect separation decisions *"if all compensations between spouses were feasible and costless"*. The assumptions underlying the Becker-Coase theorem are discussed by Chiappori et al. (2015) and Fella et al. (2004).

closely replicates the targeted moments. Our over-identification checks support the estimation results by highlighting that the model can match several non-targeted moments, among which the impact of unilateral divorce on cohabitation duration.

According to the estimates, a switch from mutual consent to unilateral divorce causes couples to start cohabiting more by reducing the married couple's ability to cooperate and by increasing the likelihood of a costly divorce.[5] Since cohabiting couples that would have married under the older regime are better matched and hence have a lower risk of dissolution than the average cohabiting couple, the reform increases the stability and the duration of newly formed cohabitations. The possibility of cohabitation has been crucial for translating institutional change (unilateral divorce) into social change (female empowerment within the couple). In fact, we find that the average Pareto weight of cohabiting women at the time the couple meets increases because men, fearing a larger loss of assets in a divorce than in a breakup, convince women to cohabit instead of marrying in exchange for more power in the couple. This mechanism is specific to the divorce regime where assets are split evenly. If spouses continue to own their assets separately, men would not need to choose cohabitation to insure their property. Consistent with the empirical evidence, the impact of unilateral divorce on cohabitation likelihood is lower in the model under separate ownership.

The fourth and last part of the paper conducts a series of counterfactual experiments to understand the quantitative importance of the forces that contributed to the rise of cohabitation. To assess the role of unilateral divorce, we perform a counterfactual experiment where unilateral divorce was never introduced. We find that people on average would have spent 1.24 years cohabiting instead of 2.19, while only 29.1% of people would have ever cohabited instead of 43.3%. In the second series of counterfactuals, we find that a decrease in the gender productivity gap and a drop in market prices of home goods increase the share of people that ever cohabited. Both effects are driven by a reduced scope for household specialization, which is better exploited within marriage. We also study various channels of how unilateral divorce affects welfare. The possibility of cohabiting limits the welfare losses for men who can secure their assets, while women suffer more because of couple-specific investments like children, that reduce their value of divorcing more than for men.

Literature. This paper adds to three strands of the literature. First, by documenting how divorce laws influence the choice between marriage and cohabitation, we add to the existing

[5] An increased likelihood of divorce can by itself reduce the ability of the couple to cooperate. Yet it also directly affects the marriage surplus by reducing the possibility of losing assets upon divorce. For example, if there was no wage uncertainty and women always participated in the labor market, unilateral divorce would affect marriage gains via the direct effect only.

literature that studies the effects of unilateral divorce. This policy change has been shown to affect the rate of divorce (Friedberg, 1998, Wolfers, 2006), female labor supply (Stevenson, 2008, Voena, 2015), savings (Voena, 2015), marriage rates (Rasul, 2003, 2006), children's well-being (Gruber, 2004), family violence (Stevenson and Wolfers, 2006), marriage-specific capital (Stevenson, 2007), assortative mating (Reynoso, 2020), the rise in serial monogamy (De La Croix and Mariani, 2015), and prostitution (Ciacci, 2017), among the others. We complement the findings of Rasul (2003, 2006) by showing that the decrease in marriage rates after the introduction of unilateral divorce is not only driven by more people staying single, but also by more people choosing to cohabit. This suggests that marriage and cohabitation are substitutes.[6] Our paper builds on Voena (2015), who studies how the interaction of unilateral divorce with property rights upon divorce affected married couples' household behavior. We extend her work both by considering cohabitation as an alternative relationship and by analyzing selection into partnership. This paper also extends the work of Fernández and Wong (2017) by showing that not considering cohabitation as an alternative to marriage biases upwards the negative impact of unilateral divorce on men's welfare. The intuition is that men can limit the losses stemming from the increased risk of divorce by cohabiting.

Second, our paper adds to the literature that studies the choice between marriage and cohabitation. A first subset thereof has focused on identifying the gains from marriage and cohabitation, highlighting the role of commitment (Matouschek and Rasul, 2008), specialization within the couple (Gemici and Laufer, 2014), learning about match quality (Brien et al., 2006), income dynamics (Blasutto, 2020), assets (Lafortune and Low, 2017, 2020) and investment in children (Lundberg and Pollak, 2015). We extend these works by showing how an increase in the ease of divorce decreases the couple's ability to cooperate and makes divorce allocations more relevant for partnership choices, since the likelihood of divorce increases. Consequently, the relative power of potential partners and the rules about the division of assets upon divorce become crucial. These results highlight a new role for partners' relative power and assets in partnership choices, which is analyzed within a framework that extends the theory of Blasutto (2020) and Gemici and Laufer (2014) by including saving decisions.[7]

Another subset of these papers studies the effect of changes in cohabitants' rights on part-

[6] Cohabitation can also be a substitute for being single or dating, as Rindfuss and VandenHeuvel (1990) point out. Moreover, Blasutto (2020) and Brien et al. (2006) claim that cohabitation can also be a complement for marriage, which allows the couple to learn about its match quality before making the binding decision of getting married.

[7] Lafortune and Low (2020) also highlight the role of assets: our model features their intuition that assets can act as a commitment technology, but our framework also allows assets to influence partnership choices via a direct effect of the risk associated with divorce. Thanks to this mechanism, we can explain why unilateral divorce caused cohabitation to increase more in community property states than in title-based ones.

nership choices and cohabiting couples' behavior, highlighting the role of alimony rights (Chiappori et al., 2017; Goussé and Leturcq, 2018), taxation (Leturcq, 2012) and division of assets at breakup (Chigavazira et al., 2019; Fisher, 2012; Goussé and Leturcq, 2018). We extend this literature by showing that the introduction of unilateral divorce impacts both the decision to cohabit and cohabitation's stability, even though cohabitants' rights are not directly affected.[8] Further, the effects on the intention to cohabit depends on property division rights, which indicates that partnership choices depend on divorce allocations. This evidence suggests that the design of changes in family law should treat marriage and cohabitation as substitutes.

Finally, this paper is tied to the extensive literature that studies the changes in character of the American household over the last decades. Various studies explored the role of health improvements, wage distribution and dynamics, norms and technology in the rise in female labor force participation (Albanesi and Olivetti, 2016; Fernández et al., 2004; Greenwood et al., 2016, 2005), the changes in household formation and dissolution (Ciscato, 2019; Greenwood et al., 2016), the rise in positive assortative mating (Ciscato, 2019; Fernandez et al., 2005; Greenwood et al., 2016) and the increase in the age at marriage (Santos and Weiss, 2016). We extend this literature by showing that the introduction of unilateral divorce was followed by a rise in cohabitation. Advances in the home production technology and the reduction in the gender wage gap also contributed to the rise.

The paper is organized as follows. Section 2 offers an overview of US divorce laws. Section 3 documents the effect of introducing unilateral divorce on partnership choices. Section 4 presents and develops the theoretical model. Section 5 describes the model's estimation, while section 6 discusses the main mechanisms of the model. Section 7 reports the results of the welfare analysis. Section 8 performs a series of counterfactual experiments, while Section 9 contains the conclusion.

2 US Divorce and Cohabitation Laws: an Overview

Divorce Laws. Between the late 1960s and early 1980s, most US states experienced fundamental changes in the divorce law. These changes affected both the right to initiate a divorce without the other spouse's consent and the division of assets upon divorce.

Before the 1960s the vast majority of US states had a mutual consent divorce regime.[9] Both spouses' agreement was needed to obtain a divorce for mundane reasons (i.e., without mis-

[8] Matouschek and Rasul (2008) study the effect of a decrease in the cost of divorce, proxied by unilateral divorce, on marriage and divorce rates using a framework where cohabitation is a choice. Since they abstract from intra-household bargaining, they cannot capture the effect of property rights upon divorce.

[9] All the states apart from New Mexico, Oklahoma and Alaska.

conduct by either spouse). However, divorce was still permitted for grounds showing guilt of misconduct by either of the two spouses: for those cases, the innocent party's agreement alone was enough to have a divorce granted. Examples of guilt or misconduct are adultery or abandonment.

From the late 1960s and early 1980s, most US states switched to a unilateral divorce regime. Under this regime divorce can be filed by one spouse without the consent of the other. More detailed chronology about the introduction of unilateral divorce in different states can be found in table A.1in the appendix.

Another dimension along which divorce laws differ across states and over time is property division. In the United States, there are three types of regime:

1. *Community Property*. Under this regime the couple jointly owns family wealth, both that obtained during the marriage and before. This implies that when divorce occurs, each spouse gets precisely half of the total family wealth.

2. *Equitable distribution*. Under this regime, the court decides how to split family wealth between the two spouses. This decision is driven by the principle of equity, which is ambiguous. In some cases, the wealth is divided exactly in half; in others, a larger share is allocated to the party that contributed the most to its accumulation.

3. *Title Based Regime*. Under this regime, wealth is split according to the title of ownership, as the spouses own their assets separately.

The option of signing prenuptial agreements gives the couple the power to agree to split the assets differently, avoiding the legal prescription,, but legal scholars believe that their effect is quite limited. In fact, these contracts could not be enforced by courts until the 1970s. After the introduction of the Uniform Premarital Agreements Act of 1983, it has been easier to enforce these contracts even though today prenuptial agreements are signed in a minority of marriages (5-10%) according to Rainer (2007), which might be due to social stigma or lack of information on their benefits (Mahar, 2003).

Breakup/Divorce laws compared. The regulation of cohabitation in the US is limited, and small changes have been made since the 1960s, when *"Cohabitation created no rights or obligations"* (Garrison, 2008). The research by Garrison (2008) also analyzes the effects of the *Marvin vs. Marvin* case (1976), where palimony — a compensation from one member of an unmarried couple to another after breakup — was awarded to the female partner: *"[the case has] not produced results markedly different from those permissible under pre-Marvin case law."* Finally, she argues that claims for financial relief have rarely reached the courts because 1) cohabitation is usually very short and not committed 2) cohabitants are younger and poorer than marrieds and 3) cohabitants do not usually adopt sharing behavior, unlike in the Marvin cases. Similarly, Bowman (2004) claims that remedies based on the contract had a limited application.

Hence, breakup resembles unilateral divorce because one partner can end cohabitation without the other partner's consent. Concerning property division rights, cohabitation de facto falls under the title-based property regime. One crucial difference between divorce and breakup is that the former requires the couple to undergo a legal process, which implies monetary and time costs, while the latter does not.[10] The lower costs of a breakup are consistent with the findings of Avellar and Smock (2005), who show that for women the drop in income following the couple's breakdown is larger for a divorce than for a breakup. To further support the claim that divorce is more costly than a breakup, in appendix B we select from the PSID a sample of couples that divorced/broke up to study how their net-worth changes after splitting. The point estimates of several event studies indicate that richer couples' divorce results in a loss of assets, while we could not observe the same pattern for the divorce and breakup among poorer couples.

3 Data and Empirical Evidence

Is the introduction of unilateral divorce related to the rise of cohabitation? Figure 1 suggests that the link between these two events merits investigation. The left panel shows a negative correlation between the share of couples that decide to marry as opposed to cohabit, and the share of couples that are formed under a unilateral divorce regime. Even more interestingly, the right panel shows that the decrease in the share of couples that decide to marry instead of cohabiting over time accelerates once unilateral divorce is introduced. Since these two graphs do not control for possible confounders and do not provide a credible counterfactual, in this section we tackle these issues to offer more convincing evidence on the effect of unilateral divorce on i) partnership choices and on ii) the pool of people who cohabit.

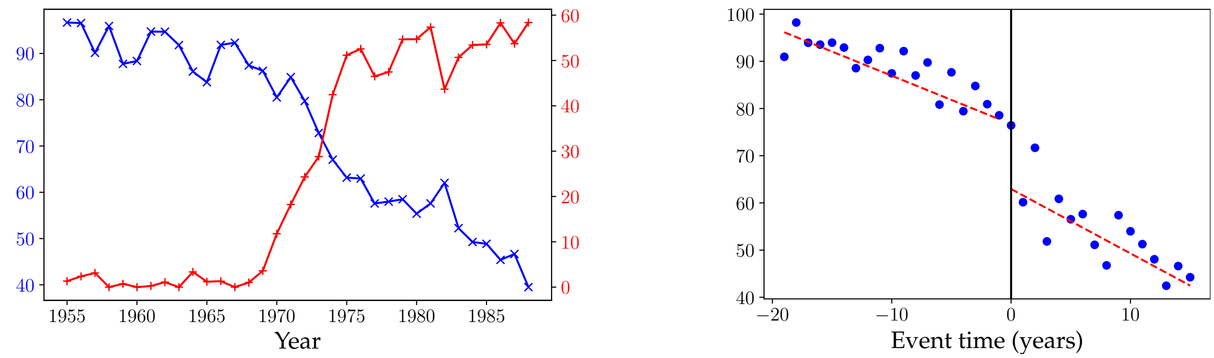
[10] While Garrison (2008) argues that claims for financial reliefs after a breakup are rare, the breakup is treated like a divorce under the doctrine of common law marriage, a legal framework under which a couple is considered as married without having formally registered their relationship. Lind (2008) explains that the existence of the implied contract is presumed once continuous cohabitation and reputation (holding out as husband and wife) are proven. However, it is still possible that the couple — even if cohabiting for many years—is not considered to be in a marriage agreement, with marital rights and obligations. These rules create uncertainty regarding recognizing common law marriage for some couples, especially those close to a breakup, where the two partners might disagree about the existence of an implied marital agreement.

FIGURE 1

Newly formed relationships (either married or cohabiting) and unilateral divorce (U.D.)

(A) % new couples choosing marriage instead of cohabitation and % relationships born under U.D.

(B) % new couples choosing marriage instead of cohabitation around U.D.



NOTES. The left panel shows in blue the evolution over time of the % of relationships formed in year t such that the couple chose marriage as opposed to unmarried cohabitation. The red line represents the % of relationships formed in t in a state that already adopted unilateral divorce by year t. The right panel shows the % of relationships where the couple chose to marry instead of cohabiting in a year whose distance in time from the introduction of unilateral divorce in that state is equal to d years. The red dotted lines are obtained by running linear regressions on a dummy for marriage, using the event time as the only regressor. All the variables presented in this figure are constructed with a sample of first and second relationships (which can be either marriage or cohabitation) from the 1988 wave of the NSFH: we provide further details about the survey and the sample construction in section 3. All the variables depicted in the two figures are constructed using sample weights.

3.1 Dataset

We begin by describing our data. We use the wave I (1987-1988) of the National Survey of Family and the Household (NSFH), and the National Survey of Family Growth (NSFG), 1988 wave. Both surveys were designed to study the causes and consequences of changes happening in families and households within the United States. This is reflected in detailed questions regarding the retrospective family history of respondents, including information about both marriage and cohabitation. Moreover, primary respondents are asked a large set of questions regarding their socio-economic background and the demographics of the household.[11] While the NSFH I is the first of three longitudinal waves, NSFG is made of several repeated cross sectional samples.[12] A drawback of using this data is that we know the state of residence of the respondents only at age 16 for the NSFH and at birth for the NSFG.[13] Since we also know

[11] One adult per household was randomly selected as the primary respondent, while in the NSFG respondents are all women of 15-44 years of age.

[12] We decided not to use the other two other waves of the NSFH because in the second wave all currently cohabiting households were dropped from the survey. Moreover, the 1988 wave of NSFG is the only one with publicly available information about the residence of the respondents, which is crucial for identifying the divorce regime that applies to the respondent.

[13] We do not have the choice of using other surveys for our analysis, since they either lack the state of residence variable, or they miss information about cohabitation history.

whether people lived all their life in the same state, we can overcome this and perform our empirical analysis both on the universe and on the subsample of never movers. We will show that point estimates turn out to be statistically indistinguishable between the two samples. Further details regarding those two surveys can be found in Bumpass et al. (2017) and Mosher and Bachrach (1996). We use this dataset to build two samples, the one of *first and second relationships* and the one of *first cohabitations*, that are described below.

First and Second Relationships Sample. We build a sample to analyze the type of relationship that respondents decided to have, which can be either marriage or cohabitation. The sample is made of first and second relationships.[14] One first relationship is defined observing the first time (if ever) a certain person started cohabiting or married. This observation is associated to the date at which the relationship starts, to the characteristics of the respondent member of the formed couple, and with a *type*, which can either marriage or cohabitation. Note that the type of relationship of couples that cohabited before marriage is “cohabitation": the transition from cohabitation to marriage is analyzed using the sample of *first cohabitations*. Second relationships are defined in a similar fashion, but they include only respondents that ended the relationship with their first partner and started a new one with a different person. The way this sample is built implies that for some respondents we will have zero corresponding observations in this sample, while for others we will have one, and for others we will have two. We did not consider third or higher order relationships for our analysis since these individuals would be further away from the age at which we knew their state of residence. Finally, we consider only relationships that lasted at least one month and started when the respondent was 20 years old or older. Relationships that started before 1955 are dropped to minimize the recall bias. In table 1 we report the descriptive statistics of this sample.

[14] Dating is not considered, since we cannot observe this state. Hence, people dating will fall under the category of singles.

TABLE 1
Descriptive statistics, relationship sample

Statistic	N	Mean	Median	St. Dev.
Unilateral Divorce Dummy	10,533	0.349	0	0.477
Age Relationship Starts	10,533	25.471	23	7.214
Married	10,533	0.650	1	0.477
College	10,533	0.252	0	0.434
Female	10,533	0.655	1	0.475
Birth year	10,533	1,950.016	1,952	10.630
NSFH Dummy	10,533	0.733	1	0.442

First Cohabitation Sample. This sample is built to analyze the decisions of cohabiting couples to breakup or to marry. It is composed of the first non-marital cohabitation experienced by respondents. This sample includes couples that cohabited before marriage, but it also includes cohabitations experienced by people with the following marital history: marriage without premarital cohabitation, divorce, cohabitation with a different person. Each observation of this sample is associated with a starting date, a possible ending date, and an outcome, which can be still cohabiting, married or breakup. In table 2 we report the descriptive statistics of this sample.

TABLE 2
Descriptive statistics, cohabitation sample

Statistic	N	Mean	Median	St. Dev.
Unilateral Divorce Dummy	5,675	0.454	0	0.498
Age Cohabitation Starts	5,675	23.701	22	6.976
Year Cohabitation Starts	5,675	1,978	1,980	7.160
College	5,675	0.162	0	0.368
Female	5,675	0.758	1	0.428
Cohabitation Duration (months)	5,675	24.170	13	29.513
Year of birth	5,675	1,954	1,956	13.790
NSFH Dummy	5,675	0.562	1	0.496
Censored	5,675	0.102	0	0.303
Married	5,675	0.490	0	0.500
Separated	5,675	0.408	0	0.491

3.2 Empirical Evidence

Does unilateral divorce affect the partnership choice of couples? We exploit the timing in the adoption of unilateral divorce as a source of exogenous variation in the right to divorce.[15] This strategy has already been used several times in the literature to study the non-neutrality of the rights to divorce on various economic and demographic outcomes.[16] According to Gruber (2004), who reviews the legal literature about the topic, the introduction of unilateral divorce was not intended as a tool of social policy, but rather a way to reduce the legal burden of divorce trials. This reasoning is consistent with the fact that this change was not initiated by the most liberal states: New York was the last state to introduce unilateral divorce in October 2010, almost 40 years later than Kentucky. Moreover, Reynoso (2020) shows that there is no geographic correlation in adoption.

Relationship Choice

What is the effect of unilateral divorce on the partnerships that couples choose? To answer this question, we estimate equation (1), where i is the newly formed couples, t is the calendar time, and s is the state:

$$\text{married}_{i,t,s} = \beta_0 + \beta_1 \cdot \text{Unilateral}_{t,s} + \gamma' \mathbf{X}_i + \delta_s + \nu_t + \epsilon_{i,t,s}. \tag{1}$$

The dependent variable is a dummy that takes value 1 if the couple i, established at time t in state s is a marriage, and 0 if it is cohabitation. The vector $\mathbf{X}_i$ includes a set of socio-demographic controls, while δ_s are the state fixed effects and ν_t are the time fixed effects. The variable $\text{Unilateral}_{t,s}$ is a dummy that takes value 1 if unilateral divorce was in place in state s at time t: β_1 is the coefficient that is informative about the effect of unilateral divorce on partnership choice. The results of the estimation are reported in table 3 for different samples. Column (1) reports the results for the full sample described in section 3.1, while column (2) is restricted to observations for which we know that the person lived their own life in the reported states, ensuring that they did not migrate. Finally, columns (3) and (4) restrict the sample to respectively the NSFH and NSFG surveys only.

[15] See table A.1 for the timing of adoption of unilateral divorce.

[16] Among the others, see Wolfers (2006), Stevenson (2008), Voena (2015), Reynoso (2020) and Ciacci (2017).

TABLE 3
OLS Regression. Observation: first and second relationships

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
Unilateral Divorce	−0.069***	−0.088***	−0.077***	−0.067*
	(0.020)	(0.021)	(0.025)	(0.037)
State Fixed effects	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	10,533	6,846	7,722	2,811
R^2	0.146	0.166	0.163	0.139

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

The results reported in table 3 suggest that unilateral divorce decreased the share of couples that are married by $7-8\%$ depending on the specification. These results are robust to an alternative specification that includes state specific linear trends (table F.1), to the use of a multinomial logit that takes into account the triple choice between staying single, cohabiting and marrying (table F.5), and to dropping California from the sample (table F.3). Moreover, table F.6 shows that the shift towards cohabitation holds both for households where the respondent has some children and where she/he is childless and does not want any children. The limitations of using two way fixed effects estimators is highlighted by a recent literature (Goodman-Bacon, 2018 and de Chaisemartin and D'Haultfœuille, 2020) that casts doubts on their validity when treatment effects are heterogeneous across time. We address these issues following the recommendation of Goodman-Bacon (2018) and use an event study design as a robustness check. The results, reported in figure F.1, show that the effect stays significant and it is even slightly larger in size.

We then examine the heterogeneity hidden behind the effect of unilateral divorce. While in some states assets are split in the same way in both breakup and divorce, which is the case of *title-based regime* states, in others this rule is different, which is the case of *community property* and *equitable distribution* states. Analyzing this heterogeneity is interesting for understanding how much the asset sharing rule is important for understanding relationship choices. We hence estimate equation (2)

$$\begin{aligned}\text{married}_{i,t,s} = \beta_0 + &\beta_1 \cdot \text{Unilateral}_{t,s} \cdot \text{No Title Based}_{t,s} \\ &+ \beta_2 \cdot \text{Unilateral}_{t,s} \cdot \text{Title Based}_{t,s} + \\ &\quad \beta_3 \cdot \text{Title Based}_{t,s} + \gamma' \mathbf{X_i} + \delta_s + \nu_t + \epsilon_{i,t,s},\end{aligned} \tag{2}$$

whose indexes and controls are the same as in equation (1), with the difference that now we capture the interaction of unilateral divorce with asset division regimes by interacting $\text{Unilateral}_{t,s}$ with $\text{Title Based}_{t,s}$ and $\text{No Title Based}_{t,s}$, which indicates whether state s at time t had or not a title-based regime. In table 4 we report the results of the estimation of equation 2. Similarly to table 3, column (1) reports the results for the full sample described in section 3.1, while column (2) is restricted to the observations for which we know that the person lived all their life in the reported states, which ensures that they did not migrate. Finally, columns (3) and (4) restrict the sample to respectively the NSFH and NSFG surveys only.

TABLE 4
OLS Regression. Observation: first and second relationships

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
UnDiv*NoTit	-0.074^{***}	-0.090^{***}	-0.084^{***}	-0.068^{*}
	(0.020)	(0.022)	(0.025)	(0.039)
UnDiv*Tit	-0.015	-0.053	-0.014	-0.046
	(0.031)	(0.037)	(0.040)	(0.048)
Tit	-0.014	-0.011	-0.011	-0.017
	(0.021)	(0.026)	(0.027)	(0.037)
State Fixed effects	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	10,533	6,846	7,722	2,811
R^2	0.147	0.166	0.164	0.139

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

The results show that the effect of unilateral divorce on the likelihood that a couples chooses marriage over cohabitation in non-title-based states is significant with a magnitude between -7% and -9% depending on specification, while it is not significant and smaller in title-based states. These results suggest that having a sharing rule decided by the law is not enough to replace the mutual consent regime as an alternative commitment technology. These results are consistent with the view that the richest partner is less inclined to marry when divorce becomes unilateral, since they stand to lose more than their partner would upon divorce. This does not happen in a mutual consent regime, since they could exercise their right to veto the divorce. In a title-based state this threat to the richer member of the couple does not exist, hence marriage surplus with respect to cohabitation does not vary significantly. These results are robust to an alternative specification that includes state specific linear trends (F.2), to the use of a multinomial logit that takes into account the triple choice between staying single, cohabiting and marrying (F.5), and to dropping California from the sample (F.4).

Cohabitation Duration

What is the effect of unilateral divorce on cohabitation duration? How much of the change is due to a variation in the risk of breakup versus the risk of marriage? To answer this question, we construct a model of cohabitation duration with multiple risks, namely breakup and marriage. Our model builds on Jenkins (1995), who shows that a logistic regression can be used for studying the duration of events by reshaping the dataset to obtain unit of time per spells observations, where the dependent variable takes the value 1 whenever the event of interest occurs. The natural extension of this model to a multiple risk environment would be to use a multinomial logit. However, the problem with this model is that it assumes independence of irrelevant alternatives, which is particularly unappealing for our problem, since it would imply that the relative probability of choosing marriage over breakup stays the same after cohabitation is no longer an option. Hence, we chose to model cohabitation duration with a multinomial probit, where the independence of irrelevant alternatives does not need to be satisfied. We then study the choice of cohabiting couple i, at calendar time t in state s and at duration d estimating the following model:

$$\begin{aligned}
Y_{i,s,t,d}^{\text{Marry}} &= \beta^{\text{Marry}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Marry}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon_{i,s,t,d}^{\text{Marry}}, \\
Y_{i,s,t,d}^{\text{Cohabit}} &= \beta^{\text{Cohabit}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Cohabit}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon_{i,s,t,d}^{\text{Cohabit}}, \\
Y_{i,s,t,d}^{\text{Breakup}} &= \beta^{\text{Breakup}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Breakup}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon_{i,s,t,d}^{\text{Breakup}},
\end{aligned} \tag{3}$$

where

$$\begin{pmatrix} \epsilon_{i,s,t,d}^{\text{Marry}} \\ \epsilon_{i,s,t,d}^{\text{Cohabit}} \\ \epsilon_{i,s,t,d}^{\text{Breakup}} \end{pmatrix} \sim \mathcal{N}(\mathbf{0}, \mathbf{\Sigma}), \tag{4}$$

and

$$Y_{i,s,t,d} = \begin{cases} \text{Marry} & \text{if } \; Y^{\text{Marry}}_{i,s,t,d} > Y^{\text{Cohabit}}_{i,s,t,d} \; \text{and} \; Y^{\text{Marry}}_{i,s,t,d} > Y^{\text{Breakup}}_{i,s,t,d} \\ \text{Cohabit} & \text{if } \; Y^{\text{Cohabit}}_{i,s,t,d} > Y^{\text{Marry}}_{i,s,t,d} \; \text{and} \; Y^{\text{Cohabit}}_{i,s,t,d} > Y^{\text{Breakup}}_{i,s,t,d} \\ \text{Breakup} & \text{otherwise.} \end{cases} \tag{5}$$

The model described above is estimated with Bayesian techniques via Markov chain Monte Carlo following the procedure of Imai and Van Dyk (2005), which is implemented using the standard options provided by the *R* package *MNP* developed by Imai et al. (2005). In table 5 we report results from the full sample in column (1), from the resident only sample in column (2) and from the observations coming from only the NSFH and NSFG surveys respectively in columns (3) and (4). Table 5 reports the parameters of the multinomial probit and the average relative risk that the event of interest (marriage of breakup) is realized.[17] The results show that unilateral divorce caused an increase in the duration of cohabitation, which comes from a reduced hazard both of marriage and of breakup. While the result about the risk of marriage is not unexpected in light of the estimation results described above, the reduced risk of breakup brings new insights about the possible mechanisms underlying partnership choices. In fact, the decrease in the risk of breakup is consistent with a selection effect: some cohabiting couples would have married if mutual consent divorce was still in place. If the match quality of cohabitations is lower than that of marriages,[18] unilateral divorce drives down the risk of breakup because of a selection effect.

[17] These risks are computed relatively to the probability to continue cohabiting.

[18] This seems plausible because the risk of divorce is much lower than the risk of breakup.

TABLE 5
Multinomial Probit. Observation: person-month of cohabitation

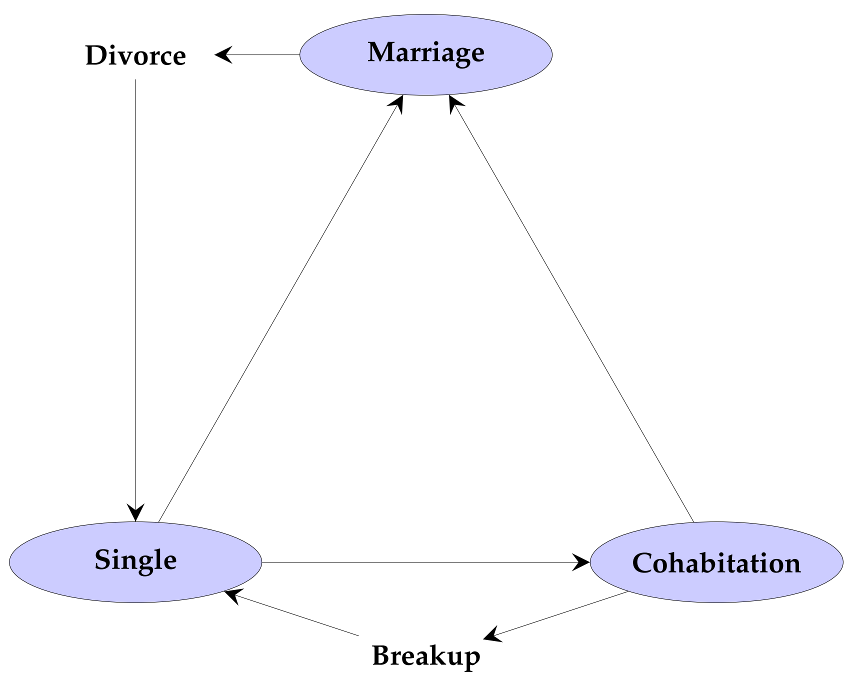
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
	Risk of Marriage relative to Cohabitation			
Unilateral Divorce	-0.24^{***}	-0.25^{***}	-0.28^{***}	-0.28^{***}
	(0.06)	(0.08)	(0.09)	(0.09)
Average Relative Risk	0.64	0.63	0.59	0.6
	Risk of Breakup relative to Cohabitation			
Unilateral Divorce	-0.19^{***}	-0.16^{***}	-0.08	-0.24^{*}
	(0.07)	(0.06)	(0.05)	(0.14)
Average Relative Risk	0.67	0.71	0.83	0.62
State Fixed effects	Yes	Yes	Yes	Yes
Year Fixed effects	Yes	Yes	Yes	Yes
Age Polynomial	Yes	Yes	Yes	Yes
Pice-wise Duration	Yes	Yes	Yes	Yes
Observations	138012	81920	77826	60186
Censored spells(%)	10.18	10.98	11.6	8.38

NOTES: the values reported in the table are the mean and the standard deviation (in parenthesis) of the posterior distribution of parameters obtained using the Markov chain Monte Carlo estimation described by Imai and Van Dyk (2005). Coefficients' distributions whose interpercentile range do not contain 0 are denoted by the following system: *90%, **95% and ***99%.

4 Theory

To identify the channels through which unilateral divorce impact partnership choice, we develop a dynamic life-cycle model of partnership formation and dissolution, savings, female labor force participation and home production. Couples act cooperatively, and according to the divorce regime they can be subject to limited commitment, which means that there might be renegotiations in response to changes in the outside options, which are assumed to be divorce or breakup. Time is discrete and in each period men and women draw their productivities. If single, with some probability they meet a potential partner: after drawing a match quality shock they decide whether to marry, cohabit or to stay single. Couples observe the match quality shock, their productivity and assets, and decide whether to stay together or to split. Cohabiting couples can also decide whether to marry. Both singles and couples make consumption and saving decisions, using their money for private or public good expenditure. Couples also make female labor participation decisions and women's time can be used to produce public goods, but this comes at the cost of a loss in productivity.

FIGURE 2



4.1 Preferences

Women f and men m derive utility from consuming a private good c and a household public good Q. The public good can be interpreted in terms of both the quantity and quality of children, as well as the goods and services produced within the household, such as washing clothes or preparing meals. Preferences are separable in the two goods and across time. Agents derives utility from a couple specific love shock ψ, which evolves over time and can be interpreted as the value of love and companionship in a couple. The intra-period utility of a single agent $s \in (f, m)$ is:

$$u(c_t^s, Q_t^s) = \frac{c_t^{s\,1-\sigma}}{1-\sigma} + \alpha \frac{Q_t^{s\,1-\xi}}{1-\xi},$$

where the superscript s on Q accounts for the fact that there is no partner to share the public good. The utility for an agent $s \in (f, m)$ in a couple is:

$$u^C(c_t^s, Q_t) = \frac{c_t^{s\,1-\sigma}}{1-\sigma} + \alpha \frac{Q_t^{1-\xi}}{1-\xi} + \psi_t,$$

where the match quality ψ evolves according to the following law of motion:

$$\psi_t = \psi_{t-1} + \epsilon_t, \text{ where } \epsilon_t \overset{\text{i.i.d.}}{\sim} \mathcal{N}(0, \sigma_\psi^2).$$

The love shock at first meeting can have a different variance, denoted by $\sigma^2_{\psi,I}$. Note that if the couple is cohabiting, the utility of the two partners is decreased by γ, which captures the stigma associated with premarital sex, premarital cohabitation and out-of-wedlock births. This assumption fits the fact that for people born in 1940-1955 (whose behavior will be used to build the target moments for the structural estimation) conservative attitudes towards premarital sex were common.[19]

4.2 Wages

The labor income for agents $s \in \{f, m\}$ depends on their age t and on a permanent income component z_t^s:

$$\ln(w_t^s) = f_t^s + z_t^s,$$

where f_t^s is a gender specific function that captures the evolution of productivity over age. The permanent income component z_t^s evolves over time as:

$$z_t^s = z_{t-1}^s - (1 - P_t^s)\mu + \zeta_t^s, \text{ where } \zeta_t^s \overset{\text{i.i.d.}}{\sim} \mathcal{N}(0, \sigma_\zeta^{2_s}), \text{ and } \zeta_1^s = z_1^s. \tag{6}$$

[19] The shame associated with an out-of-wedlock birth, whose interaction with technology is studied by Fernández-Villaverde et al. (2014), can be a factor leading young women to prefer marriage over cohabitation even if the rules governing these two partnerships were identical. Blasutto (2020) can match closely marriage and cohabitation choices using a theoretical framework close to ours, without needing to introduce a stigma towards cohabitation. This is possible because he analyzes the behavior of people born in 1980-1984, for whom the stigma towards premarital sex and premarital cohabitation is arguably lower than for those born in 1940-1955.

where P_t^s is a dummy of labor force participation. Men and single women are always assumed to participate in the labor market, hence $P_t^m = 1$.[20] Parameter μ is the loss in productivity that affects women that are not participating in the labor market. It can be interpreted as a reduced form way of capturing both the missed opportunity to accumulate human capital while working and the skill atrophy from interruptions (Adda et al., 2017). Modeling the loss in productivity for not working is an important feature of our model as it creates an incentive to join the labor force for women that expect to divorce or breakup soon.

4.3 Home Production

In our model each agent has one unit of time. Singles and men in a couple supply inelastically a fraction $1-\phi$ of their time to the labor market, while women in a couple can be out of the labor force, devoting their time producing the home good Q. The public good can also be produced buying d goods in the market. Following Greenwood et al. (2016) we define the production function of home goods for couples as:

$$Q_t = [d_t^{\nu} + \kappa(2\phi + (1-P_t^f)(1-\phi))^{\nu}]^{\frac{1}{\nu}}, \text{ where } 0 < \nu < 1, \tag{7}$$

while for singles of gender $s \in \{f, m\}$

$$Q_t^s = [(d_t^s)^{\nu} + \kappa\phi^{\nu}]^{\frac{1}{\nu}}. \tag{8}$$

The parameter ν captures the degree of substitutability between women's time and the use of durables in the production of home goods. This structure implies that when the relative price of d_t decreases and when wages go up,[21] women spend less time producing household goods and their employment outside the home increases.

4.4 Budget Constraints

The budget constraint of a single agent of gender $s \in \{f, m\}$ is:

$$a_{t+1}^s = Ra_t^s + w_t^s(1-\phi) - c_t^s - d_t^s, \text{ with } a_{t+1}^s \geq 0, \tag{9}$$

[20] The assumption that men, as opposed to women, always participate in the labor market is rather common in the literature (Ciscato, 2019; Low et al., 2018; Voena, 2015) and it is in line with the gender roles typically observed in the period under analysis. In our PSID sample only 5% of men between 20 and 60 do not supply working hours in the market.

[21] The relative price of d_t is normalized to 1 in equations (7) and (8)

where a^s are agent's savings and w^s is the wage. c^s and d^s are the private good consumption and the expenditure used to produce the public good. The budget constraint for a couple is:

$$a_{t+1}^f + a_{t+1}^m = Ra_t + w_t^m(1-\phi) + P_t^f w_t^f(1-\phi) - c_t^f - c_t^m - d_t, \text{ with } a_{t+1} \geq 0, \tag{10}$$

When a couple divorces in t, we assume

$$a_t^m + a_t^f = \delta a_t,$$

where δ is the fraction of total assets a_t left after divorce. We assume $\delta = 1$ for breakup.[22] An important feature of our model is the role of property rights, which defines how assets are divided upon divorce/breakup. Since we use data from community property states to estimate the model, this regime applies to divorce. Accordingly, upon divorce each spouse keeps half of the assets, while the division of assets upon breakup is a couple's decision. We describe the details of this choice in this section, where the problem of the cohabiting couple is presented.

4.5 Problem of the Singles

We start by describing the problem for a single agent $i \in \{f, m\}$ in t. The agent makes consumption, saving and expenditure decisions. In $t+1$ she meets a potential partner j of the opposite sex with probability λ_{t+1} and she can decide to enter a partnership, which also depends on whether the potential partner will agree. If the two decide to marry, the variable M_{t+1} will take value 1, while $C_{t+1} = 1$ if the couple decides to cohabit. Otherwise, M_{t+1} and C_{t+1} will be equal to 0. The state variable of a single is $\omega_t^i = \{a_t^i, z_t^i\}$, while her choices are represented by the vector $\mathbf{q}_t^i = \{a_{t+1}^i, c_t^i, d_t^i\}$. We denote by $V_t^{i,S}(\omega_t^i)$ the value function of agent i, which we define

[22] The assumption that divorce erodes a fraction of wealth is common to Cubeddu and Ríos-Rull (2003). In appendix B we provide evidence that divorce results in a loss of net worth for rich but not for poor households. Moreover, we do not find evidenceof a loss of net worth following breakup for rich and poor households. In practice the cost of breakup is positive because of psychological distress associated with a separation and because looking for new accommodation takes time. However, these costs are common with divorce and hence they do not help explaining why couples should choose one partnership over the others. This reasoning is confirmed by the fact that when we tried estimating the model allowing for a positive cost of breakup, this parameter was not identified.

as

$$\begin{aligned}V_t^{iS}(\omega_t^i) = &\max_{\mathbf{q}_t^i} u(c_t^i, Q_t^i) + \beta E_t \Big\{ (1-\lambda_{t+1}) V_{t+1}^{iS}(\omega_{t+1}^i) + \\ &\lambda_{t+1}\big\{(1-M_{t+1})(1-C_{t+1})V_{t+1}^{i,S}(\omega_{t+1}) + \\ &M_{t+1}V_{t+1}^{i,M}(\Omega_{t+1}) + C_{t+1}V_{t+1}^{i,C}(\Omega_{t+1})\big\}\Big\}, \\ &\text{s.t. (9) and (8),}\end{aligned} \tag{11}$$

where $V^{i,M}$ and $V^{i,C}$ are the individual values of being married and cohabiting.

4.6 Household Planning Problem

The problem faced by the couple depends both on the type of relationship—cohabitation or marriage—and on the divorce regime, which can be either *mutual consent* or *unilateral divorce*. Breakup is always unilateral. Under the unilateral regime, one partner can initiate the breakup/divorce process alone, while under mutual consent the agreement of both partners is needed.

Mutual Consent Regime

Under a mutual consent regime, marriage is denoted by $\hat{M}$. Couples solve a Pareto problem where the weight of the wife is θ^f and that of the husband is $1-\theta^f$.[23] The state vector is $\Omega_t^{\hat{M}} = \{a_t^m, a_t^f, z_t^f, z_t^m, \psi_t, \theta^f\}$, while the variables over which the couple maximizes are summarized by the vector $\mathbf{q}^{\hat{M}_t = \{a_{t+1}^f, a_{t+1}^m, d_t, c_t^m, c_t^f, P_t^f, D_t\}}$, where D_t is a dummy variable that takes value 1 if divorce occurs and 0 otherwise. The formal problem solved by a couple who enters period t as married is:

$$\begin{aligned}V_t^{\hat{M}}(\Omega_t^{\hat{M}}) = &\max_{\mathbf{q}_t^M}(1-D_t)\{\theta^f u(c_t^f, Q_t) + (1-\theta^f)u(c_t^m, Q_t) + \psi_t + \beta E_t V_{t+1}^{\hat{M}}(\Omega_{t+1}^{\hat{M}})\} \\ &+ D_t\{\theta^f V_t^{fS}(\omega_t^f) + (1-\theta^f)V_t^{mS}(\omega_t^m))\} \\ \text{if } D_t = 0: \quad & \text{s.t. (10) and (7)} \\ \text{if } D_t = 1: \quad & \text{s.t. (9), (8) for } i \in \{f, m\}, \\ & a_t^m + a_t^f = \delta a_t, \\ & V_t^{fS}(\omega_t^f) \geq W_t^{f\hat{M}}(\Omega_t^{\hat{M}}), \\ & V_t^{mS}(\omega_t^m) \geq W_t^{m\hat{M}}(\Omega_t^{\hat{M}}).\end{aligned} \tag{12}$$

[23] Later in this section we describe how initial Pareto weights are set.

The individual value of marriage conditional on $D_t = 0$ is $W_t^{i\hat{M}}$ for $i \in \{F, M\}$, and it is defined as

$$W_t^{i\hat{M}} = u(\tilde{c}_t^i, \tilde{Q}_t) + \psi_t + \beta E_t V_{t+1}^{i\hat{M}}(\Omega_{t+1}^{\hat{M}}), \tag{13}$$

where $\tilde{\mathbf{q}}_t^{\hat{M}} = \{\tilde{a}_{t+1}^m, \tilde{a}_{t+1}^f, \tilde{d}_t, \tilde{c}_t^m, \tilde{c}_t^f, \tilde{P}_t^f\}$ is the $\arg\max$ of problem (12) conditionally on having chosen $D_t = 0$. $V_{t+1}^{i\hat{M}}(\Omega_{t+1}^{\hat{M}})$ can be obtained by the expectation of the sum of the time utilities that the agent gets from $t+1$ to T, where the variables entering the utility function derive from the Pareto problem if the agent is in a relationship, otherwise they are the solution of (11), which represents the singles' problem.

Under the mutual consent regime, the allocation corresponds to the Pareto efficient solution if the couple is intact. Intuitively, the fact that Pareto weights stay constant allows for functioning risk-sharing and the female labor force participation decisions are taken cooperatively, ruling out the possibility that women over-supply labor to increase their bargaining power. In this framework, the conditions for divorce are particularly stringent: the couple splits only if both partners are better-off divorcing than staying together for a feasible allocation. Moreover, if only one spouse wishes to divorce under the divorce allocation dictated by the law where assets are split equally, they will "bribe" the other by offering a larger share of assets to make them indifferent between staying married and divorcing.[24]

Unilateral Divorce Regime

Under the unilateral divorce regime marriage is denoted by $\overline{M}$. Couples solve a Pareto problem where the weight of the wife is θ_t^f and that of the husband is θ_t^m. Note that, unlike in the mutual consent regime, Pareto weights can vary over time. The state vector of this problem is $\Omega_t^{\overline{M}} = \{a_t, z_t^f, z_t^m, \psi_t, \theta_t^f, \theta_t^m\}$, while the variables over which the couple maximize are summarized by

[24] Note that if both partners are better off divorcing under the sharing rule dictated by the law, which corresponds to an equal division in community property states, no bribing happens.

the vector $\mathbf{q}_t^{\overline{M}} = \{\tilde{a}_{t+1}, \tilde{d}_t, \tilde{c}_t^m, \tilde{c}_t^f, \tilde{P}_t^f\}$. The formal problem of a couple entering t as married is:

$$\begin{aligned}
V_t^{\overline{M}}(\Omega_t^{\overline{M}}) = & \max_{\mathbf{q}_t^M}(1-D_t)\{\theta_t^f u(c_t^f, Q_t) + \theta_t^m u(c_t^m, Q_t) + \psi_t + \beta E_t V_{t+1}^{\overline{M}}(\Omega_{t+1}^{\overline{M}})\} \\
& + D_t\{\theta_t^f V_t^{fS}(\omega_{t+1}^f) + \theta_t^m V_t^{mS}(\omega_t^m))\} \\
\text{if } D_t = 0: \quad & \text{s.t. (10) and (7),} \\
& \theta_{t+1}^f = \theta_t^f + \mu_t^f, \\
& \theta_{t+1}^m = \theta_t^m + \mu_t^m, \\
\text{if } D_t = 1: \quad & \text{s.t. (9), (7) for } i \in \{f, m\}, \\
& a_t^m + a_t^f = \delta a_t, \\
& a_t^m = a_t^f,
\end{aligned} \tag{14}$$

where θ_{t+1}^f and θ_{t+1}^m adjust such that the following participation constraints are satisfied:

$$\begin{aligned}
W_t^{f\overline{M}}(\Omega_t^{\overline{M}}) \geq V_t^{fS}(\omega_t^f), \\
W_t^{m\overline{M}}(\Omega_t^{\overline{M}}) \geq V_t^{mS}(\omega_t^m).
\end{aligned} \tag{15}$$

Note that μ_t^i are the Lagrange multipliers associated with spouses' participation constraints. The individual value of marriage conditional on $D_t = 0$ is denoted by $W_t^{i\overline{M}}$ and it can be obtained following the procedure described in the mutual consent regime section.

Under the unilateral divorce regime Pareto weights vary every time one participation constraint is binding. Whenever a spouse is better off divorcing, the other member will try to convince them not to split by offering them more bargaining power, such that she is indifferent between divorcing and staying married. In this framework risk-sharing is less functional than under the mutual consent regime, since variations in the Pareto weight imply less smooth consumption patterns over time. Labor market specialization is also less functioning, since conditionally on having the same state variables, the risk of divorce is higher, which makes women willing to insure against this event through labor market participation. While cooperation is more effective under mutual consent than unilateral divorce, it is still possible that the individual value of being married under the latter regime is larger. This is possible because of the possibility of exiting the marriage without the consent of the other spouse.

Cohabitation

The problem of cohabiting couples is like that of marriage under the unilateral divorce regime, but it differs in three crucial ways. First, there is no loss of assets upon breakup. Second, the choice set of the cohabiting couple $\mathbf{q}_t^C = \{a_{t+1}, d_t, c_t^m, c_t^f, P_t^f, D_t, M_t, \chi_{t+1}\}$ and the state variables $\Omega_t^C = \{a_t, , z_t^f, z_t^m, \psi_t, \theta_t^f, \theta_t^m, \chi_t\}$ are different. Note that M_t is a dummy that indicates the choice of marrying and χ_t is the share of assets going to the woman in case of breakup.[25] Third, the time utility of the cohabiting couple is decreased by γ. The complete problem of the cohabiting couple can be found in section D of the appendix.

The fact that there is not breakup cost makes risk-sharing and cooperation less functional compared to marriage. This happens because the couple is left without a commitment-enhancing technology, which would have allowed the couple to improve its ability to commit.[26] On the other hand, assuming no cost of breakup makes cohabitation more appealing to couples whose risk of splitting is high. For example, this is the case of couples with a low match quality.

Property rights upon divorce/breakup differ between marriage and cohabitation. In the former, assets are divided equally when the couple splits, while in the latter assets are divided according to individual property rights. We model property rights at breakup following Bayot and Voena (2015), where upon divorce assets are split following the sharing rule decided by the couple in the previous period.[27] They show that this regime is always preferred to community property if outside options are invariant to property right regimes. In our framework this result implies that if the cost of breaking up was the same as that of divorcing and there was no stigma towards cohabitation, the value of cohabitation would always be higher than that of marriage. The benefits of having a positive cost of divorce and the stigma linked to cohabitation allows us to generate a positive number of marriages and to match the data.

4.7 Partnership Choice and the Mating Market

In each period t singles have a probability λ_t of meeting a potential partner. The productivity and the assets of the potential partner depend on the single agent's characteristics. Formally, the assets of the potential partner p are defined as:

$$\ln(a_t^p) = \ln(a_t^s) + \overline{a}^s + \epsilon^a, \tag{16}$$

[25] The way we model the title-based regime follows Bayot and Voena (2015).

[26] Under the limit case of an infinite cost of splitting, as long as the couple stays intact the allocation under the mutual consent and unilateral divorce regimes are the same and correspond to the inter-temporal Pareto-efficient allocation.

[27] Bayot and Voena (2015) study the choice between community and separation of property in Italy. Separation of property resembles the American title-based regime, which also applies to cohabitation.

where a_t^s are the assets of the individual, $\overline{a}_t^s$ is a number that depends on gender and ϵ^a is a normally distributed shock. The productivity of the potential partner is defined as:

$$\ln(z_t^p) = \alpha(\ln(\overline{z}_t^{s^*,i^*}) + \epsilon_t^z) + (1-\alpha)\ln(z_t^r), \tag{17}$$

where $\overline{z}_t^{s^*,i^*}$ represents the average productivity of singles of gender s^*, z_t^r is the productivity of the agent net of the gender and education-specific trend, while ϵ_t^z is a normally distributed shock. These assumptions capture in a reduced form fashion that people are mating assortatively both within marriage and cohabitation. Once the meeting occurs, agents must decide whether to stay in a couple and eventually decide which partnership contract to choose, and they must pick a Pareto weight. We now describe how these decisions are taken. Note that for the rest of this section we will refer to marriage as M, where $M \in \{\hat{M}, \overline{M}\}$ depending on the divorce regime. The decisions follow a three-steps procedure.

1. The couple considers marriage M (cohabitation C) as a viable alternative if the set of Pareto weights θ^f such that the couple prefers to marry (cohabit) is non-empty.[28] Formally, for relationship $J \in \{M, C\}$ the set is

$$\Theta_t^J(\Omega_t^J, \omega_t^f, \omega_t^m) = \left\{\theta_t^f : V_t^{fJ}(\Omega_t^J) \geq V_t^{fS}(\omega_t^f), V_t^{mJ}(\Omega_t^J) \geq V_t^{mS}(\omega_t^m)\right\}. \tag{18}$$

2. If the set for marriage (cohabitation) is non-empty, the Pareto weight for the potential marriage $\theta_t^{M,f}$ (cohabitation $\theta_t^{C,f}$) is set through symmetric Nash Bargaining.[29] Formally $\theta_t^{J,f}$ is set to :

$$\theta_t^{J,f} = \underset{\theta_t^f \in \Theta_t^J}{\arg\max}\, \Upsilon^J(\theta_t^f, \Omega_t^{J-1}, \omega_t^f, \omega_t^m), \tag{19}$$

where Ω_t^{J-1} is the state vector of the couple excluding Pareto weights and

$$\Upsilon^J(\theta_t^f, \Omega_t^{J-1}, \omega_t^f, \omega_t^m) = \left[V_t^{fJ}(\Omega_t^{J-1}, \theta_t^f) - V_t^{fS}(\omega_t^f)\right] \times \left[V_t^{mJ}(\Omega_t^{J-1}, 1-\theta_t^f) - V_t^{mS}(\omega_t^m)\right]. \tag{20}$$

3. Four possible situations can arise:

   - $\Theta_t^M =$ and $\Theta_t^C = \;\Rightarrow$ stay single.
   - $\Theta_t^M \neq$ and $\Theta_t^C = \;\Rightarrow$ marry.

[28] Without loss of generality, we impose $\theta_t^f + \theta_t^m = 1$ at first meeting.

[29] The assumption that the initial Pareto weight is pinned down by Nash Bargaining can be found in Mazzocco (2007) and Low et al. (2018).

- $\Theta_t^M =$ and $\Theta_t^C \neq \Rightarrow$ cohabit.
- $\Theta_t^M \neq$ and $\Theta_t^C \neq \Rightarrow$ the couple chooses the partnership that gives the largest Nash product. Formally, if $\Upsilon^M(\theta_t^{M,f}, \Omega_t^M, \omega_t^f, \omega_t^m) \geq \Upsilon^C(\theta_t^{C,f}, \Omega_t^C, \omega_t^f, \omega_t^m)$ the couple chooses marriage, otherwise cohabitation.

5 Estimation

We estimate the structural model following a two-step procedure. The first step is to set some parameters following the literature or by matching some features of the data without the need to simulate the model. In particular, we estimate the labor income processes of men and women outside the model: this procedure is common in the literature because it reduces the burden on structural estimation.[30] The second step is to estimate by indirect inference the remaining parameters of the model. In this section we detail the steps of the estimation, we discuss the identification of the structural parameters and we present the results.

5.1 Income Processes

The income processes of men and women are estimated using the 1968-1993 waves of the PSID, including people between age 20 and 65. We further restrict our sample by retaining men who are household heads or men who are married/cohabiting with the household head or who are household heads themselves. Similarly to Low et al. (2018), we drop observations where the hourly wage is less than half the minimum wage and where the hourly wage changes by more than 125% in two consecutive years. We compute the hourly wage rate of men and women, dividing the annual labor income by the number of yearly working hours supplied. This procedure avoids treating a variation in working hours as a productivity shock. This correction is particularly relevant for the estimation of the income process of women, because their hours worked vary significantly over the life-cycle. The income process of men is estimated by fitting the following linear model:

$$\ln(w_{i,t,s,sur}^m) = \iota_0^m + \iota_1^m \cdot t + \iota_2^m \cdot t^2 + \delta_s + \nu_{sur} + u_{i,t,s,sur}^m, \tag{21}$$

where i stands for individual, t for age, s for state and sur for survey year. Moreover, $u_{i,t,s,sur}^m = z_t^m + e_{i,t,s,sur}^m$, where z_t^m follows equation 6, while $e_{i,t,s,sur}^m$ is the measurement error. δ_s are state fixed effects and ν_{sur} are year of the survey fixed effects. The results are reported in table E.1. Then, using the residuals $\hat{u}_t^m$, we estimate through GMM 1) the variance of the permanent

[30] See for example Voena (2015), Reynoso (2020) and Gourinchas and Parker (2002).

component of income σ_ζ^{2m}, 2) the variance of the measurement error σ_e^{2m} using the following conditions:

$$\begin{aligned} E((\Delta\hat{u}_t^m)^2) &= \sigma_\zeta^{2m} + 2\sigma_e^{2m} \\ E(\Delta\hat{u}_t^m \Delta\hat{u}_{t-1}^m) &= -\sigma_e^{2m} \end{aligned} \tag{22}$$

Results are reported in table 6.

The estimation of women's income process differs from the men's one since we need to consider the endogeneity of female labor force participation. We do so by using a two-step Heckman selection correction procedure. The first step consists in estimating a probit model where the dependent variable is female labor force participation and the independent variables includes all the regressors in equation (21) plus the interaction of a dummy variable for unilateral divorce with the dummy variables for the property rights regimes upon divorce. These variables are used as an exclusion restriction following the work of Voena (2015), who finds that these affect female labor force participation by influencing intra-household bargaining.[31] Women participate in the labor market if

$$\gamma' \mathbf{Z}_{i,t,s,sur} + \pi_{i,t,s,sur} > 0, \tag{23}$$

where $\pi_{i,t,s,sur}$ is the sum of the measurement error and the permanent component of income and $\mathbf{Z}_{i,t,s,sur}$ contains the regressors. The second setup is estimating the following linear model:

$$\ln(w_{i,t,s,sur}^f) = \iota_0^f + \iota_1^f \cdot t + \iota_2^f \cdot t^2 + \delta_s + \nu_{sur} + \varphi_{i,t,s,sur} + u_{i,t,s,sur}^f, \tag{24}$$

where i stands for individual, t for age, s for state and sur for survey year. Moreover, $u_{i,t,s,sur}^f = z_t^f + e_{i,t,s,sur}^f$. z_t^f follows equation 6, while $e_{i,t,s,sur}^f$ is the measurement error. δ_s and ν_{sur} are respectively state and year of the survey fixed effects. The endogeneity of female labor force participation is considered by controlling for $\varphi_{i,t,s,sur}$, the inverse of the Mills ratio of the prediction obtained in the first step. The estimation results of the two steps are reported in tables E.3 and E.2. We then use the regression residuals from the second step $\hat{u}_t^m$ to estimate through GMM 1) the variance of the permanent component of income σ_ζ^{2f}, 2) the variance of the mea-

[31] Voena (2015) and Reynoso (2020) already used the interaction between grounds of divorce and division of property as an exclusion restriction for female labor force participation.

surement error $\sigma_e^{2^f}$ using the following conditions:[32]

$$\begin{aligned}E(\Delta\hat{u}_t^f|P_t^f=1,P_{t-1}^f=1)&=\sigma_\pi^f\frac{\phi(\tau_t)}{1-\Phi(\tau_t)},\\E((\Delta\hat{u}_t^f)^2|P_t^f=1,P_{t-1}^f=1)&=\sigma_\zeta^{2^f}+\sigma_\pi^{2^f}+2\sigma_e^{2^f}+\tau_t\frac{\phi(\tau_t)}{1-\Phi(\tau_t)},\\E(\Delta\hat{u}_t^f\Delta\hat{u}_{t-1}^f|P_t^f=1,P_{t-1}^f=1,P_{t-2}^f=1))&=-\sigma_e^{2^f}.\end{aligned}\tag{25}$$

where $\phi()$ and $\Phi()$ are respectively the density and the distribution function of a standardized normal, while $\tau_t=-\gamma'\mathbf{Z}_{i,t,s,sur}$. Results are displayed in table 6.

TABLE 6
Parameters of the income processes

Parameter	Symbol	Value
f's age return (constant)	ι_0^f	-0.383
f's age return (linear component)	ι_1^f	0.0244
f's age return (squared component)	ι_2^f	-0.0005
Variance of f's permanent income shock	$\sigma_\zeta^{2^f}$	0.0399
m's age return (constant)	ι_0^m	-0.342
m's age return (linear component)	ι_1^m	0.0495
m's age return (squared component)	ι_2^m	-0.0009
Variance of m's permanent income shock	$\sigma_\zeta^{2^m}$	0.0417

NOTES: The parameters are estimated using nonlinear least squares using single, cohabiting and married males and females from the PSID.

5.2 Preset Parameters

In this section we describe how we fix the set of preset parameters. Each period in the model lasts 1 year: we chose this length balancing the benefits of having a short period, which fits the fact that cohabitation spells are particularly short, and the computational burden associated with having too many periods. We assume that men (women) start making decisions at age 20 (18). Couples are always formed by men who are 2 years older than women. Agents retire at the age of 62 and the number of periods in the model is $T = 62$. The discount factor β and the relative risk aversion σ of private goods match those in Attanasio et al. (2008). The annual interest rate is set to 2%. The parameters relevant to the production of public goods, ν and κ match those in McGrattan et al. (1997). As far as the pensions are concerned, I follow Heathcote et al. (2010): they consider the progressive nature of the US system but they simplify

[32] The conditions are those used by Low et al. (2018).

it, assuming that only the last period before retirement is relevant for the amount of the pension that a person receives. Parameter ϕ is set to 0.189 to reflect the relative time that singles spend on house works relative to the time spent on the labor market.[33] Wages are normalized such that average log wages of male at age 30 is 0. The variance of male (female) earnings at age 20 (18), $\sigma^{2_m}_{\zeta,1}(\sigma^{2_f}_{\zeta,1})$ is taken directly from the PSID data. The parameters regarding the mating market, contained in equations 16 and 17, are pinned down to obtain a realistic degree of assortative mating with respect to assets and wages. In particular, we target the correlation in log wages in the PSID and the share of households with family income above the median whose wealth is also above the median in the Survey of Consumer Finances (SCF). The parameters of the mating market are pinned down to respect a second condition, which is *symmetry*. For example, married men at age t should have on average the same wage and wealth regardless of being simulated for their life cycle, or being partners of women who are simulated for their whole life cycle.[34] Since we set these parameters before the structural estimation takes place, we cannot perfectly match the mating market moment that we targeted. The correlation in log wages of couples in the PISD is 0.58 versus 0.62 in the simulated sample,[35] while the share of people that have a wealth above the median, conditionally on having a family income above the median, is 0.76 in the Survey of Consumer Finances and 0.82 in the model.

[33] In the PSID the average yearly time spent on house work by singles is 465.5 hours. Assuming that the yearly hours of full-time work in the labor market is 2000, we get $\phi = 465.5/(465.5 + 2000) = 0.189$. The median number of yearly hours spent in the labor market for single men is 1976, while for single women is 1848. We considered the 1940-1955 birth cohorts of the PSID for these computations because the moments that we will use in the structural estimation are based on the behavior of people born in those years.

[34] The agents who belong to our fictional sample are simulated for their whole life cycle and they marry/cohabit with partners that they meet randomly. The behavior of these partners is followed only while they are in a relationship with the person in our fictional sample. Figure H.1 shows the mean and variance of productivity and wealth by age, both for agents belonging to the "fictional sample" and to their "partners". The variables of interest are similar for the two groups, which means that the two groups are symmetric with respect to these variables.

[35] We obtain this value by simulating the behavior of agents under the parametrization of deep parameters described later in this section.

TABLE 7
Preset parameters

Estimated Parameters	Symbol	Value	Source
Initial age		18-20	
Retirement age		62	
Number of time periods	T	62	
Years per period		1	
m's average earnings at 30		1	Normalization
Mating market—productivities			PSID
Mating market—assets			SCF
Pensions			Heathcote et al. (2010)
Var. f's productivity in $t=1$	$\sigma^{2_f}_{\zeta,1}$	0.54	PSID
Var. m's productivity $t=1$	$\sigma^{2_m}_{\zeta,1}$	0.54	PSID
Interest rate	$R-1$	2%	
Relative Risk Aversion private good	γ	1.5	Attanasio et al. (2008)
Discount factor	β	0.98	Attanasio et al. (2008)
Function	**Symbol**	**Value**	**Source**
$Q_t = [d_t^{\nu} + \kappa(1-P_t^f)^{\nu}]^{\frac{1}{\nu}}$	κ	3.76	McGrattan et al. (1997)
	ν	0.19	McGrattan et al. (1997)

5.3 Indirect Inference

We use the method of indirect inference (Gourieroux et al., 1993) to pin down the vector $\vartheta = (\alpha, \lambda, \sigma_{\psi}, \sigma_{\psi,I}, \delta, \mu, \xi, \gamma)$ of the 8 remaining parameters of the model. We use 31 moments and regression coefficients for the structural estimation, which capture the process of marriage and cohabitation creation and dissolution, as well as female labor supply. More precisely, we include as targets the coefficient of unilateral divorce estimated through equation 1,[36] the hazard of divorce (6), the hazard of breakup (3), the hazard of marriage (3), the share of people ever married over time (7), the share of people that ever cohabited over time (7), female labor supply (1),[37] differences in female labor force participation between marriage and cohabitation (2) and

[36] Note that the sample used for estimating equation 1 in the empirical section and in the structural estimation is different. We will describe within this section how the sample used for structural estimation is constructed.

[37] Female labor supply in the model is constructed by multiplying the indicator of female labor force participation by 2000 hours. The assumption that working full-time corresponds to 2000 hours of work in a year was also used for calibrating ϕ. Alternatively, we could have targeted female labor force participation, picking a number of hours for full-time work such that female labor supply is also matched. Since the amount of part-time work is very different according to the status (married, cohabiting or single) of the women, the number of hours for participating women should have been differed by status. The problem with this approach is that women would have chosen their partnership according to the artificially fixed working schedule that partnerships offer, and not only according to the mechanisms that our model generates.

differences in log wages between married and cohabiting men (1). We use the retrospective marital history data from the NSFH wave III to construct the moments linked to partnership choice, while we all the others are computed using the PSID.[38] The data moments are constructed selecting men and women born in 1940-1955 in community property states.

The first step for the estimation is to solve the model for a vector of parameters ϑ, then simulating income, love shocks and unexpected divorce policy changes to obtain the simulated behavior for the given parametrization. The next step is to perform stratified sampling on the simulated population in order to obtain the same distribution over gender/age/regime of divorce as in the data used to construct the moments. This allows us to compare the simulated and data moments: the objective is to obtain ϑ such that this difference is the smallest possible. Formally, the problem that we solve is

$$\hat{\vartheta} = \arg\min_{\vartheta} \qquad (\mathbf{m} - \mathbf{m}_{\vartheta})' \mathbf{W} (\mathbf{m} - \mathbf{m}_{\vartheta}), \tag{26}$$

where $\mathbf{m}$ is the vector of empirical moments, as described in the section about target moments, while $\mathbf{m}_{\vartheta}$ is the vector of the moments simulated by the model parametrized with ϑ. $\mathbf{W}$ is a matrix where the diagonal contains the inverse of the variance of the data moments, while all the other entries are zeros. The minimization of this object function is performed using the global optimization algorithm TikTak, which according to Arnoud et al. (2019) outperforms an array of global and local optimizers when the target is a difficult objective function. In appendix C we describe in detail how the algorithm TikTak works and how me modify it to allow for the possibility of running it in parallel.

5.4 Identification

This section provides a description of how the structural parameters of the model are identified heuristically. The parameter α is identified by total female labor supply: when this parameter is large, the household wants to produce more public goods which requires women's time. Parameter μ affects the gap in female labor supply for married and cohabiting couples. When μ is large the gap increases, because household specialization within cohabitation becomes relatively harder, as this relationship lacks a commitment technology. Parameter λ is intuitively identified by the share of people in a relationship. The parameter σ_{ψ} has a role in identifying

[38] NSFH wave III is conducted in 2001/2003 following the original respondents of wave 1. This sample does not include respondents under age 45 as of January 2000 unless some particular conditions are met, but this is not an issue for us since the youngest person in our estimating sample was 44 in 2000. One possible issue with this data is that by mistake during NSFH wave II all cohabiting couples were dropped by the sample. We overcome this problem by simulating the same "mistake" on the sample drawn from the simulated data.

the stability of marriage and cohabitation by modifying the likelihood that marriage surplus becomes negative, but it is mostly identified by the share of people that are choosing marriage over cohabitation. In fact, as this parameter grows larger, money becomes less important than love for total utility. This means that agents care less about insuring against income shocks and labor specialization starts binding less, while the risk of breakup and divorce increases. The parameter σ_ψ alone is not able to generate a large enough marriage surplus, such that the number of ever married people is matched. For this reason we introduced parameter γ, thanks to which we can match the share of people ever married and that ever cohabited. Also parameter δ influences the gains of marriage with respect to cohabitation, but it does so in a non-monotonic fashion. On the one hand increasing the cost of divorce enhances commitment, while on the other hand it makes more costly to end the relationship. Hence, the effect of increasing or decreasing δ depends on its initial value. Since the introduction of unilateral divorce is to a first approximation like a decrease in the cost of divorce, the parameter δ is mostly identified by the coefficient of unilateral divorce of regression 1. Parameter $\sigma_{\psi,I}$ is identified by the hazard of breakup and marriage: when this parameter is small compared to the variance of the transitory shocks, agents are not *picky* about sorting into cohabitation, but they move fast to a marriage or they separate within the first periods of the relationship, according to the evolution of the love and productivity shocks. Finally, the parameter ξ influences the surplus of marriage and cohabitation by wealth. In fact, when ξ is small, wealthier agents find the consumption of the public good Q relatively more attractive. Since marriage makes it possible to consume a larger quantity of Q because it protects women that devote time to its production, marriage becomes a relatively more interesting option for wealthier families. Hence, ξ is identified by the difference in log wages of married and cohabiting men.

5.5 Model Fit

Table 8 reports the results of the structural estimation. The estimated standard deviation σ_ψ of the transitory match quality shock is 0.76, while standard deviation $\sigma_{\psi,I}$ of the love shock at first meeting is higher with a value of 1.67. The probability of meeting a partner λ is 0.38, while the share of assets left after divorce is 0.80. The weight on the public good α is 1.20, while the loss in productivity parameter μ is 0.07. Finally, the penalty for cohabiting γ is 0.15, while the coefficient of relative risk aversion for the public good ξ is 1.14.

The fit of the model is reported in table 9. The model matches generally well the hazard of marriage, breakup and divorce over time, even though it lies outside the 95% confidence interval of data moments. One exception is that the hazard of divorce and breakup are not hump shaped over the duration because our model abstracts from learning, which is necessary to match this pattern in the data (Blasutto, 2020). The share of people that ever cohabited and married over time is well matched. The data about female labor supply is well matched. The differences in log wages for married and cohabiting men are lower than in the data. Finally, the coefficient of unilateral divorce estimated through equation (1) is slightly larger than in the data, but it lies within the 95% confidence interval.

The model is validated according to its ability to reproduce the effects of unilateral divorce on cohabitation duration, the share of income in the couple earned by married women, the average wage earned by women over their working life and the ratio of hazard rates of richer over poorer men.[39] We use two Cox duration models (Cox, 1972) to study how the risk of breakup and marriage for cohabiting couples are affected by the introduction of unilateral divorce.[40] The results, reported in table 9, show that using both the empirical and the simulated sample the policy decreases the risk of marriage and breakup. Overall, the length of cohabitation increases by 27% in the simulated sample. Women's wages over their life-cycle and the average share of income provided by the wife in the household match the data, which validates the selection of women into the labor force. The fact that the model matches that divorce rates are lower for richer men supports our assumptions regarding the cost of divorce, which influences both the allocation within divorce and the surplus of marriage.

A further test for our model is to check whether the effect of unilateral divorce on the propensity to cohabit is lower under a title-based regime than under a community property regime, as it is in the data. We solve the model assuming a title-based regime and we obtain that the coefficient of unilateral divorce of equation (1) is -0.09, while it was -0.16 under community property.[41] This result is consistent with the idea that under community property regime the shift towards cohabitation is larger because men, who are those with the most decision power, start finding cohabitation attractive when the risk of divorce increases. This is because upon divorce they would lose most of their assets, leaving a part of them to their ex-wife. This mechanism bites less under a title-based regime, because men would keep the assets of their property upon divorce.

[39] Richer men are those whose income is above the median, while poorer men are those whose income is below the median.

[40] Since the aim of this exercise is to study how the pool of cohabiting couples changes after the reform, we exclude cohabitation spells that already started when the law changed. When the event of interest is marriage (breakup), breakup (marriage) is treated as if the spell was censored.

[41] Note that we do not expect to match exactly the empirical coefficient (1) under the title-based regime because we did not re-estimate the model using a sample of residents in title-based states. The parameters used for this exercise are those in table 9.

TABLE 8
Estimated structural parameters

Estimated Parameters		Value
Standard deviation of match quality shock	σ_ψ	0.76
Standard deviation of initial match quality shock	$\sigma_{\psi,I}$	1.67
Probability of meeting a partner	λ	0.38
Assets left upon divorce	δ	0.80
Weight of public good	α	1.20
Loss in productivity while not working	μ	0.07
Relative Risk Aversion public good	ξ	1.14
Penalty of Cohabiting	γ	0.15

TABLE 9
Model fit and validation

Estimated Moments	Model	Data	95% CI
Hazards over Time	fig. G.1	fig. G.1	fig. G.1
Share Ever Cohabited and Married	fig. G.2	fig. G.2	fig. G.2
FLS in a Couple (hours)	1007	1016	[1002,1029]
FLS if Married/ FLS if Cohabiting (<35 yrs.)	1.02	0.86	[0.78,0.95]
FLS if Married/ FLS if Cohabiting (≥35 yrs.)	0.97	1.00	[0.89,1.13]
Log wages Marriage-Log wages Cohabitation	-0.08	0.12	[0.04,0.12]
Unilateral Divorce coefficient equation (1)	-0.16	-0.11	[-0.21,-0.02]
External Moments	**Model**	**Data**	
Unilateral Divorce on the relative Risk of Marriage	0.72	0.42	[0.20,0.85]
Unilateral Divorce on the relative Risk of Breakup	0.92	0.26	[0.10,0.67]
Women wages over their life-cycle	fig. G.3	fig. G.3	fig. G.3
Divorce Rate Rich/Divorce Rate Poor	0.74	0.79	[0.75,0.84]
Share household income earned by women	0.34%	0.35%	[0.36-0.38]

NOTES: The coefficients and the relative hazard ratios in the table differs from those obtained with the same econometric model in section 3.2. The reason is that the sample used for the empirical part is different from that used for structural estimation as explained in the section.

6 Mechanisms

The aim of this section is 1) to better understand the quantitative relevance of the mechanisms underlying the introduction of unilateral divorce and the subsequent rise of cohabitation and 2) to quantify the gains of marriage with respect to cohabitation.

We start by analyzing how selection and intra-household bargaining change as a result of the reform. The estimated structural model allows us to study the evolution of the match quality

ψ and women's Pareto weight θ_t using a standard event study. Specifically, we estimate the following regression model on simulated data

$$\text{Variable of Interest}_{i,a,t} = \sum_{j=-5}^{5} \beta_j^{Uni} \cdot \mathcal{I}(t = j) + \alpha_0 + \alpha_a + \epsilon_{i,t} \tag{27}$$

where a is age, t is the year relative to switching to unilateral divorce ($t = -1$ is omitted) and i is a couple. We estimate the model for ψ and θ using as samples 1) cohabiting couples that just met 2) married couples that just met. Figure 3 reports the results. We normalize the coefficient estimates β_j^{Uni} by adding the average of the variable of interest in the year before unilateral divorce is introduced $E[\text{Variable of Interest}|t = -1]$.

Match quality ψ. We start by analyzing panel (a). First, note that the average match quality of married couples is higher than for cohabitants.[42] This fact is consistent with a strong selection on marriage and cohabitation with respect to match quality. Marriage guarantees a better commitment and cooperation, but when the match quality is low the best option is to choose cohabitation because breaking up is cheaper than divorcing. The results of the event study show that upon the introduction of unilateral divorce the match quality of newly formed cohabitations increases by a value that is around 35% percent of its structural standard deviation.[43] This result is consistent with selection of relatively high match quality couples into cohabitation after the policy change. This happens because unilateral divorce increases the risk of dissolution of marriage and affects the spouses'incentive to cooperate.

Women's bargaining power θ. Panel (b) depicts the evolution of women's bargaining power θ at meeting for cohabitation and marriage around the introduction of unilateral divorce. The average initial Pareto weight θ increases with respect to baseline for cohabitation after the policy change, while it decreases for marriage. Under mutual consent, marriage protected women against ending up divorced and poor, while cohabitation was chosen only by couples where the man was not able to commit to a long term relationship and the woman had little say about the decision. After the reform, men prefer cohabitation over marriage because the former avoids splitting up assets equally upon breakup, but in exchange women obtains a higher initial Pareto weight θ.[44] Similarly, the Pareto weight of women that marry goes down because men are willing to marry instead of cohabiting only if they can control more resources

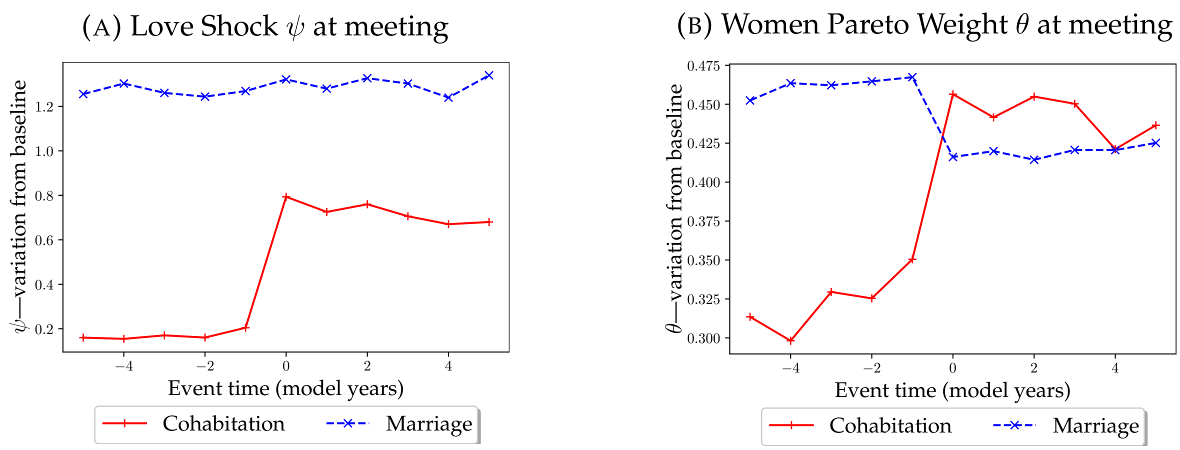
[42] A more in depth analysis reveals that the distribution of match quality at meeting of cohabiting couples dominates that of married couples. See figure H.2.

[43] Note that the observed and structural distributions of the initial match quality are different because couples are not formed when the match quality at meeting is too low.

[44] Note that upon breakup men receive on average around 65% of the couple's wealth.

within the household.

FIGURE 3
Event Studies Around the Introduction of Unilateral Divorce–Simulated Data



NOTES The figures display the evolution of the love shock ψ and the female Pareto weight θ around the introduction on Unilateral Divorce. The displayed patterns are normalized coefficients from event studies around divorce. The graphs are relative to couples that started a relationship.

Risk-sharing and consumption insurance. The veto over divorce in the mutual consent regime and the higher cost of divorce compared to breakup are commitment technologies that enforce cooperation within marriage. What is the quantitative relevance of these mechanisms on the ability of married and cohabiting couples to share risk? To answer this question we study how income shocks affect the changes in consumption over time for men in a couple.[45] First, we obtain paths of consumption and labor earnings of men who are in their first relationship by simulating their choices under different divorce regimes.[46] Then, for each of these samples we estimate the equation below

$$\Delta \log c_{it} = \alpha + \mu \Delta \log(w_{it}) + \nu_t + \epsilon_{it},$$

where c_{it} and w_{it} are the consumption and the labor earnings of men i at age t. ν_t are age fixed effects. Coefficient μ is informative about how much of the change in income translates

[45] Results for women in a couple are displayed in table H.1. We chose to show the results for men in this section since changes in their productivity always translate into changes in disposable income. This does not happen for women who do not participate in the labor market: for this reason results in table H.1 are largely driven by the different degree of female labor force participation under different divorce regimes and partnership types.

[46] First relationships correspond to the periods in which an agent spends time in a couple with his or her first partner without changing partnership type. This means that the first relationship of a man than first cohabited and then married his first partner is a cohabitation spell that stops when the couple gets married. We restrict our attention to first relationships only because this allows us to perform an exercise where we impose marriage on a couple that decided to cohabit. This exercise allows us to compare marriage and cohabitation controlling for selection.

into changes in consumption. Consequently, a low value of μ corresponds to a high degree of consumption insurance. The results displayed in table 10 provide important pieces of information about partnership types and consumption insurance. First, coefficient μ is the largest for cohabitation and the smallest for marriage under mutual consent divorce, as reported in the first row.[47] This means that consumption insurance is the most effective within marriage with mutual consent divorce and the least effective within cohabitation. Does this happen because of a selection effect or because partnership rules directly affect the ability to share risk? Rows 2 and 3 of table 10 suggest that selection matters: marriages that were not preceded by cohabitation display a stronger degree of consumption insurance than those that were preceded by cohabitation. This is because the initial match quality of couples that married directly is higher compared to those who first cohabited. The fact that match quality is higher implies that participation constraints bind more rarely, allowing for a smoother consumption path. Row 4 of table 10 shows that if we “force" men who had chosen to cohabit to marry, we find that the amount of consumption insurance lies in between that of marriage and cohabitation. Since this experiment controls for selection, we conclude that partnership rules have a direct effect on couples' ability to share risk. The results relative to consumption insurance and partnership types are driven by the frequency at which Pareto weights are renegotiated. Consistently with these results, figure H.4 shows that the share of periods in which Pareto weights are renegotiated is higher for cohabitation than marriage, and that only some of this difference can be explained by selection. This suggests that consumption insurance is tightly linked to the frequency with which the couples renegotiate the way resources are shared.

[47] We only consider individuals who spend their whole life-cycle under the same divorce regime.

TABLE 10
Partnership type and consumption insurance against income shocks

	Married and Cohabiting Men		
	Married		Cohabiting
	M.C.	U.D.	
Baseline	0.450	0.484	0.498
Only marriages preceded by cohabitation	0.454	0.486	-
Only marriages not preceded by cohabitation	0.449	0.483	-
Marriages with cohabitation selection	0.468	0.496	-

NOTES: the table reports the estimates of coefficients μ obtained from regression

$$\Delta \log c_{it} = \alpha + \mu \Delta \log(w_{it}) + \nu_t + \epsilon_{it}.$$

The sample includes the whole duration of the first relationship of simulated men i. The last row is run on a sample of men who decided to cohabit but we imposed marriage on them instead. This allow us to analyze the insurance within marriage controlling for selection into a relationship.

Consumption smoothing upon divorce/breakup. Is there a link between partnership types and consumption smoothing upon divorce/breakup? To answer this question we analyze the evolution of log consumption around divorce/breakup using a standard event study on the simulated sample. Note that, after the relationship breakdown, we report the log consumption of the household of the partner that is simulated for her/his whole life-cycle. Specifically, we estimate the following regression model

$$\log(c)_{i,a,t} = \sum_{j=-3}^{3} \beta_j^{Split} \cdot \mathcal{I}(t = j) + \alpha_0 + \alpha_a + \epsilon_{i,a,t}, \tag{28}$$

where a is age of the person observed after the relationship dissolves, t is the year relative to breakup/divorce ($t = -1$ is omitted) and i is the household. Note that we included age fixed effects. We estimate this model separately for formerly married and cohabiting households under different divorce regimes, following either men or women after the divorce/breakup. Figure H.3 report the results and shows two main facts. First, consumption drops more after divorce than breakup because the former is costly in terms of assets. Second, women lose more because they are less productive than men, especially if they devoted their time to producing home goods while they were in a couple.

7 Welfare

Previous research already studied the welfare effects of the introduction of unilateral divorce: both Reynoso (2020) and Fernández and Wong (2017) find that this policy change decreases welfare for both genders and that the loss for women is larger than for men. While we find a similar effect, in this section we claim that accounting for cohabitation results in an even stronger difference by gender in states where assets are split evenly upon divorce. To study well-being under the two divorce regimes we perform an *ex-ante* welfare comparison, where for each gender we compute the expected value of spending the whole life cycle under a certain regime, before the realization of productivity and love shocks. Table 11 reports the results, which show that welfare under a unilateral divorce regime is lower than under mutual consent for both genders. The difference is larger for women, who would need to receive almost $ 13,000 in assets in $t = 0$ to be indifferent between the two regimes, while men would need only $3,244 to be indifferent between the two. To understand the role of cohabitation for the changes in well-being, we repeat the welfare analysis assuming that cohabitation in no longer a choice.[48] For ease of exposition, we refer to the model with cohabitation as model A, while model B is the one without cohabitation. The results in table 11 show that the loss of welfare related to unilateral divorce is similar under models A and B for women, while men lose more under model B. This result suggests that not accounting for cohabitation overestimates the welfare losses for men when unilateral divorce is introduced. The intuition is that cohabitation is valuable for men under the unilateral divorce regime when assets are split evenly upon divorce. This is because they stand to lose more upon divorce than upon breakup. In fact upon breakup they can keep the assets of their property.

[48] In practice, we increase the stigma parameter towards cohabitation γ such that cohabitation is never chosen.

Table 11
Welfare by gender and divorce regime

Female		**Male**	
Mutual Consent	Unilateral Divorce	Mutual Consent	Unilateral Divorce
Life-Time utilities in $t = 0$			
-364.62	-368.13	-351.53	-351.88
Welfare Losses with Unilateral Divorce			
12933.66 $		3244.04 $	
Life-Time utilities in $t = 0$ when cohabitation is not in the choice set			
-362.49	-365.86	-353.11	-354.13
Welfare Losses with Unilateral Divorce			
13783.62 $		10426.25 $	

Welfare losses are obtained by computing the amount of wealth that must be transferred to men and women in $t = 0$ such that their lifetime utility under the unilateral divorce regime equals that under mutual consent. The wealth is measured in 1990 dollars.

8 Counterfactual Experiments

The aim of this section is to understand the quantitative importance of the economic mechanisms that contributed to the rise of cohabitation during the last decades. To do so, we examine the results from a series of counterfactual experiments.

Unilateral Divorce. The qualitative impact of unilateral divorce on the choice between marriage and cohabitation has been largely discussed throughout this paper. Here we assess its quantitative relevance by performing an experiment where unilateral divorce is never introduced. Table 12 reports the share of people that cohabited at 39 and the average years spent cohabiting under the baseline scenario and the counterfactual.[49] The results show that under the counterfactual only 67% of the people would have cohabited by the age of 39, while the years spent cohabiting would have fallen from 2.19 to 1.24. The latter effect is the strongest because it captures changes in both partnership choices of singles and in the duration of partnerships.

Shrinking gender wage gap. Table 12 reports the results of another scenario where the gender productivity gap is reduced by increasing women's potential wages by 10% and men's wages are reduced by 10%:[50] the share of people that ever cohabited increases from 43.3% to 47.3%, while the number of years spent cohabiting increases from 2.19 to 2.65. In the counterfactual there is less room for specialization in the couple when the two partners' wages are more similar and the opportunity cost of not working for women rises. Hence, in the counterfactual the couple's need for commitment decreases: cohabitation becomes relatively more attractive as it implies no cost of splitting. This result is consistent with the work of Anelli et al. (2019), who find that exposure to robots causes both a decline in market opportunities of men with respect to women and a decrease (increase) in the likelihood of being married (cohabiting).

Decreasing the price of home appliances. In table 12 we report one last counterfactual experiment that explores the effects of reducing by 10% the relative price of goods d, used to produce public goods Q. This change is to be interpreted as a result of the improvement in home production technologies, such as the dish washer or the washing machine, which freed up women's time. Previous research already showed the impact of those changes on female labor supply (Greenwood et al., 2005), the decline in marriage, the rise in divorce and assortative mating (Greenwood et al., 2016). The counterfactual experiment shows that the share of people that ever cohabited increases from 43.3% to 44.8%, while the years spent cohabiting increase

[49] We consider the number of years spent cohabiting between the age of 20 and the age of 55.

[50] The increase in women's potential wages might not be realized if they decide not to participate in the labor market.

from 2.19 to 2.27. Similarly to a reduction in the gender wage gap, improvements in the technology of home production decrease the need for labor specialization within the household and for a commitment technology to enforce it. Hence, improvements in the technology of home production not only caused a decline of marriage with respect to singleness, as Greenwood et al. (2016) claim, but also a change in the relative convenience of partnership contracts.

No stigma on cohabitation. Table 12 reports the results of one last counterfactual scenario where the stigma towards cohabitation γ is set to zero. In the counterfactual, over 80% of people have ever cohabited and agents spend on average more than 11 years cohabiting. These results suggest that norms have an important role for the rise in cohabitation over time. Finally, note that many people continue marrying: in this scenario 44% of people have ever married, which suggests that the economic incentives alone are able to generate a positive surplus of marriage with respect to cohabitation for certain individuals.

TABLE 12
Counterfactual experiments

Scenario	% people ever cohabited	Years spent cohabiting
Baseline	43.3	2.19
No Unilateral Divorce	29.1	1.24
$\downarrow$ gender productivity gap	47.3	2.65
$\downarrow$ 10% Price of good d	44.8	2.27
No stigma on Cohabitation ($\gamma = 0$)	82.4	11.40

NOTES. The Baseline scenario reports the model output with the parameters reported in the previous section. The scenario “No Unilateral divorce" assumes that all the agents live under a Mutual consent regime during all their life, while in the lower productivity gender gap scenario women’s productivity is increased by 10%, while men’s productivity is decreased by 10%. The share of people that ever cohabited is measured at the simulated age of 39, while years spent cohabiting are computed between ages 20 and 55.

9 Conclusion

In this paper, we show that partnership choices depend on the rights to divorce: the introduction of unilateral divorce in most US states influenced selection into marriage and cohabitation as well as the duration of these relationships and women’s bargaining power. Using NSFH and NSFG data, we show that the introduction of unilateral divorce is responsible for a 7-8% increase in the likelihood that singles choose cohabitation over marriage, and that newly formed cohabitations last longer. To understand the mechanisms that underlie those changes, as well as the welfare effect for the two genders, we build a dynamic structural model where agents can choose to marry, cohabit and when to end these relationships. We use regression results from survey data as well as moments that describe the mating market and female labor supply to estimate our model by indirect inference. The structural estimation reveals that couples choosing cohabitation instead of marriage are those that would have had the highest risk of divorce. Since cohabiting couples had on average a lower match quality than married ones, this selection effect increases the duration of newly formed cohabitations. Moreover, in the US states where assets are split equally, it is men who wish to cohabit after the policy reform, since they would lose more assets in a divorce than in a breakup. Women are convinced to enter this relationship in exchange for higher bargaining power, even though this makes them worse off if the couple subsequently breaks up. Finally, we show that the magnitude of the overall effect of unilateral divorce on cohabitation is large: a counterfactual experiment reveals that if the law never changed, time spent cohabitating for the birth cohorts used in our estimation would have been 1.24 years instead of 2.19, while the share of people that ever cohabited would have shifted from 43.3% to 29.1%.

Beyond what is studied in this paper it would be interesting to introduce explicitly fertility in our framework to understand why children born within cohabitation do not perform well later in life. A promising approach would be to follow Kozlov (2020), who distinguishes between fertility as a choice and as an unplanned event. In fact, it might be that children raised by single mothers are outcomes of unwanted births that happen within cohabitation. This situation might happen less frequently within marriage, since it is a more stable and committed relationship than cohabitation.

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Appendix

A History of US divorce Laws

TABLE A.1
Year Unilateral Divorce was Introduced

State	Year	State	Year
Alabama	1971	Montana	1973
Alaska	1935	Nebraska	1972
Arkansas	No	Nevada	1967
Arizona	1973	New Hampshire	1971
California	1970	New Jersey	2007
Colorado	1972	New Mexico	1933
Connecticut	1973	New York	2010
District of Columbia	No	North Carolina	No
Delaware	1968	North Dakota	1971
Florida	1971	Ohio	No
Georgia	1973	Oklahoma	1953
Hawaii	1972	Oregon	1971
Idaho	1971	Pennsylvania	No
Illinois	No	Rhode Island	1975
Indiana	1973	South Carolina	No
Iowa	1970	South Dakota	1985
Kansas	1969	Tennessee	No
Kentucky	1972	Texas	1970
Louisiana	No	Utah	1987
Maine	1973	Vermont	No
Maryland	No	Virginia	No
Massachusetts	1975	Washington	1973
Michigan	1972	West Virginia	2001
Minnesota	1974	Wisconsin	1978
Mississippi	No	Wyoming	1977
Missouri	2009		

NOTES: The data of this table is taken from table 1, column (1) in Ciacci (2017)

B Net worth around divorce/breakup

In this section we provide evidence about the evolution of household's net worth around the event of divorce/breakup. Using the 1997-2017 waves of the PSID, we build a sample of 1087 divorces and 1187 breakups that respect the following characteristics: 1) household wealth is observed before and after the relationship breakdown 2) the number of adults in the household move from two to one after the relationship breakdown 3) the net worth of the household is below the 96% of the relative distribution 4) we exclude households where the head is older than 65 years old.[51] Net worth is constructed using the PSID variables employed by Blundell et al. (2016). We analyze the evolution of net worth using a standard event study on our sample. Note that, after the relationship breakdown, we report the net worth of the household of the partner that the PSID kept interviewing. Specifically, we estimate the following regression model

$$\text{Net worth}_{i,a,t,y,ma} = \sum_{j=-6}^{4} \beta_j^{Split} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_a + \alpha_y + \alpha_{ma} + \epsilon_{i,t}, \tag{29}$$

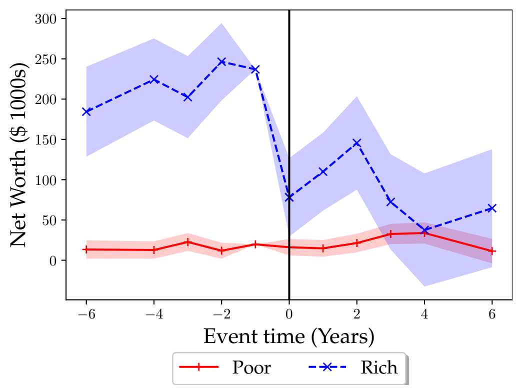
where a is age of the person observed after the couple dissolves, t is the year relative to divorce/breakup ($t = -1$ is omitted), and i is the household, y is the year and ma is the number of years since the start of the marriage/cohabitation. Note that we included year, years since marriage/cohabitation and age fixed effects. We estimate this model separately for formerly married and cohabiting households and we further subdivide our sample considering wealthier/poorer households and men/women.[52] Figure B.1 reports the results. We normalize the coefficient estimates β_j^{Split} by adding the average of net worth at divorce $E[\text{Net worth}|t = -1]$. In panel (a) we can see a decrease in net worth for richer households: the estimates indicate the year after the divorce the household is left with significantly less than half its original net worth, even though the large standard errors do not allow us to identify clearly the amount of net worth lost because of the divorce. No clear decrease in net worth can be observed for poorer households. Panel (c) shows that there is not clear loss in net worth for poor and rich cohabiting households. Finally, panels (b) and (d) show that no gender-related difference regarding the evolution of net worth can be detected.

[51] We could not distinguish the net worth of the couple/individuals against the other member if we considered households with more adults.

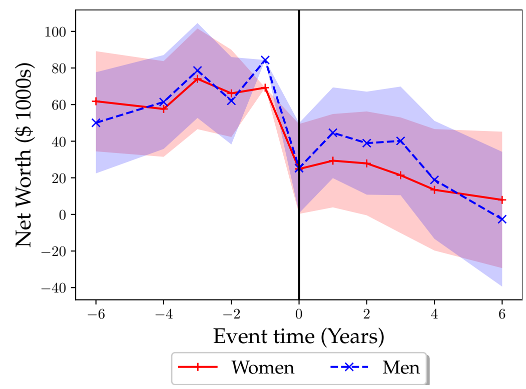
[52] A household is considered wealthy if its net worth before couple disruption is above the 75^{th} percentile of the distribution and poor otherwise.

FIGURE B.1
Event studies of net worth around divorce

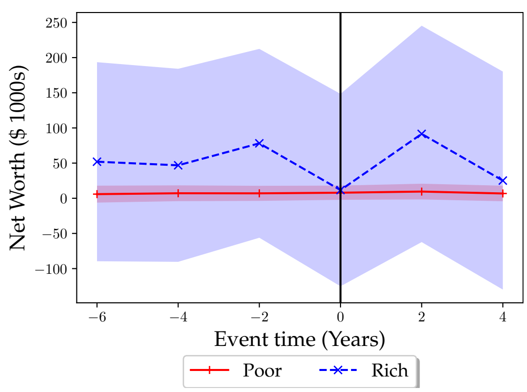
(A) Net worth—rich and poor households



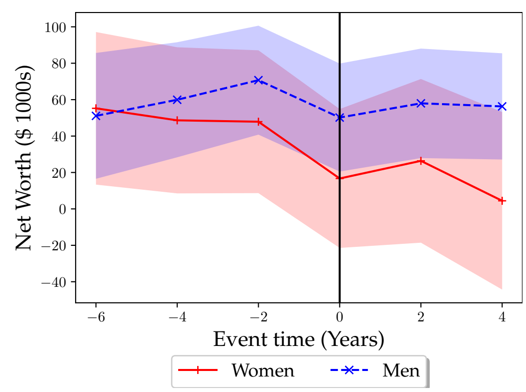
(B) Net worth—men and women



(C) Net worth—rich and poor households



(D) Net worth—men and women



NOTES. The figures display the evolution of net worth (measured in 1997$). The displayed patterns are normalized coefficients from event studies around divorce. Rich households are defined as those whose net worth is above the median in the first period they were observed. Poor households are those whose net-worth is below the 75^{th} percentile of the distribution. Net worth is constructed using the same PSID variables that Blundell et al. (2016) use.)

C Computational Appendix

Arnoud et al. (2019) compares an array of local and global optimizers, which are given the task of finding the global optimum of difficult objective functions. They find that the multi-start algorithm that they propose, called TikTak, outperforms the others in terms of time required to reach the solution and the probability that the algorithm finds the optimum. In light of these findings, we decided to use TikTak for solving problem (26). A description of the TikTak algorithm follows:

1. Determine the bounds for each parameter and generate a sequence of Sobol points with length N. Then evaluate the function value at each Sobol point.

2. Sort the N Sobol points (s_1 ,..., s_N), with $f(s_1) \leq \cdots \leq f(s_N)$ and keep the first N^* with $N^* < N$. Note that $f()$ is the objective function. We set N^* such that $N^*/N = 0.15$. Set the

global iteration number j to 1, then run a local minimizer starting from s_1. Call z_j^* the fit resulting from the local minimization,[53] and define the set $Z_1^* = \min\{z_1^*\} = z_1^*$.

3. Define a new starting point $\hat{s}_{j+1}$ defined as

$$\hat{s}_{j+1} = (1 - \theta_j)s_{j+1} + \theta_j Z_j^*,$$

where

$$\theta_j = \min\left[\max[0, 1, (j/N^*)^{\frac{1}{2}}], 0.995\right].$$

Run a local minimizer starting from $\hat{s}_{j+1}$ and call the local minimum found z_{j+1}^*. Then, define $Z_{j+1}^* = \min\{z_1^*, ..., z_{j+1}^*\}$. Update the global iteration number: $j = j + 1$. Repeat step 3 until $j = N^*$.

4. Return $Z_{N^*}^*$.

We adapt the original algorithm such that it can be run in parallel using M nodes. Other than evaluating more points at the same time on different nodes, the only difference is in step 3. In the parallel version of TikTak, Z_j^* is defined as the minimum among the outcomes of the local minimizers that already converged, while at the end of step 3 the global iteration number is updated to j^*, which stands for the number of global minimizations that already started, without necessarily having converged already.

D Problem of the cohabiting couple

Cohabiting couples, denoted by C, solve a Pareto problem where the weight of the wife is θ_t^f and that of the husband is θ_t^m.The state vector is $\Omega_t^C = \{a_t, z_t^f, z_t^m, \psi_t, \theta_t^f, \theta_t^m, \chi_t\}$, where χ_t is the share of assets going to the woman in the event of breakup. The variables over which the couple maximize are summarized by the vector $\mathbf{q}_t^C = \{a_{t+1}, d_t, c_t^m, c_t^f, P_t^f, S_t, M_t, \chi_{t+1}\}$. S_t and M_t are dummy variables that take value 1 if the couple respectively breakup or marry and 0

[53] We use the local minimization algorithm provided by Cartis et al. (2019), which is a derivative-free optimization (DFO) for nonlinear Least-Squares (LS) problems. This algorithm is robust to noise, which might arise because of the errors coming from the approximation of continuous problems on a discrete grid.

otherwise.[54] The formal problem that a cohabiting couple at t solves is:

$$
\begin{aligned}
V_t^C(\Omega_t^C) = &\max_{\mathbf{q}_t^C}(1-S_t)\{\theta_t^f u(c_t^f, Q_t) + \theta_t^m u(c_t^m, Q_t) + \psi_t - \gamma + \beta E_t V_{t+1}^C(\Omega_{t+1}^C)\} \\
&\text{if } S_t = 0: \qquad \text{s.t. (10) and (7)}, \\
&\qquad\qquad\qquad\quad \theta_{t+1}^f = \theta_t^f + \mu_t^f, \\
&\qquad\qquad\qquad\quad \theta_{t+1}^m = \theta_t^m + \mu_t^m, \\
&\text{if } S_t = 1: \qquad \text{s.t. (9), (7) for } i \in \{f, m\}, \\
&\qquad\qquad\qquad\quad a_t^m + a_t^f = a_t, \\
&\qquad\qquad\qquad\quad a_t^f = \chi_t a_t,
\end{aligned}
\tag{30}
$$

where θ_{t+1}^f and θ_{t+1}^m adjust such that the following participation constraints are satisfied:

$$
\begin{aligned}
W_t^{fC}(\Omega_t^C) \geq V_t^{fS}(\omega_t^f), \\
W_t^{mC}(\Omega_t^C) \geq V_t^{mS}(\omega_t^m).
\end{aligned}
\tag{31}
$$

Note that μ_t^i are the Lagrange multipliers associated with spouses' participation constraints. The individual value of cohabitation conditional on $S_t =$ is W_t^{iC} for $i \in \{f, m\}$, and it is defined as

$$
W_t^{iC} = u(\tilde{c}_t^i, \tilde{Q}_t^i) + \psi_t - \gamma + \beta E_t V_{t+1}^{iC}(\Omega_{t+1}^C), \tag{32}
$$

where $\tilde{\mathbf{q}}_t^C = \{\tilde{a}_{t+1}, \tilde{\chi_{t+1}}, \tilde{d}_t, \tilde{c}_t^m, \tilde{c}_t^f, \tilde{P}_t^f\}$ is the $\arg\max$ of problem (30) conditional on having chosen $S_t = 0$. $V_{t+1}^{iC}(\Omega_{t+1}^C)$ can be obtained by the expectation of the sum of the time utilities that the agent gets from $t+1$ to T, where the variables entering the utility function derives from the Pareto problem if the agent is in a relationship, otherwise they are the solution of (11). Similarly to the unilateral divorce regime, we assume that the planner evaluates the welfare of the two members of the couple if a breakup happens with the current Pareto weights.

[54] We denote marriage by M, which might fall under unilateral divorce regime $\overline{M}$ or mutual consent $\hat{M}$.

E Estimation of Income Processes

TABLE E.1
OLS Regression. Observation: males in year t.

	(1)
DEP. VARIABLE: MALE LOG EARNINGS	
ι_1^m	0.05
ι_2^m	-0.00
ι_0^m	-0.34
Survey Year Fixed Effects	✓
State Fixed Effects	✓
Observations	98118
R^2	0.152

NOTES: Standard errors are obtained through bootstrapping and they are reported in summary table 6.

TABLE E.2
OLS Regression. Observation: Females in Year t.

	(1)
DEP. VARIABLE: FEMALE LABOR EARNINGS	
ι_1^f	0.02
ι_2^f	-0.00
ι_0^f	-0.38
Survey Year Fixed Effects	✓
State Fixed Effects	✓
Observations	86891
R^2	0.085

NOTES: Standard errors are obtained through bootstrapping and they are reported in summary table 6.

TABLE E.3
Probit Regression. Observation: Females in Year t.

	(1)
DEP. VARIABLE: FEMALE LABOR FORCE PARTICIPATION	
Unilateral Divorce∗Community Property	-0.18***
Unilateral Divorce∗Title Based	-0.08
Unilateral Divorce∗Equitable Distribution	-0.06
Equitable Distribution	-0.00
ι_1^f	0.01***
ι_2^f	-0.00***
ι_0^f	1.95
Survey Year Fixed Effects	✓
State Fixed Effects	✓
Observations	127728

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

F More evidence on the impact of unilateral divorce on partnership choices

Relationship Choice - Linear State Trends

TABLE F.1
OLS Regression. Observation: first and second relationships

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
Unilateral Divorce	-0.054^{**}	-0.074^{***}	-0.064^{**}	-0.019
	(0.025)	(0.023)	(0.030)	(0.053)
State Fixed effects	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Linear trend by State	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	10,533	6,846	7,722	2,811
R^2	0.151	0.173	0.172	0.153

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

Relationship Choice - Heterogeneity by property regime and linear state trends

TABLE F.2
OLS Regression. Observation: first and second relationships

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
UnDiv*NoTit	-0.071^{***}	-0.082^{***}	-0.076^{***}	-0.033
	(0.021)	(0.017)	(0.027)	(0.053)
UnDiv*Tit	-0.024	-0.062	-0.039	0.003
	(0.037)	(0.038)	(0.045)	(0.047)
Tit	-0.039	-0.038	-0.033	-0.054^{*}
	(0.027)	(0.032)	(0.033)	(0.032)
State Fixed effects	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Linear trend by State	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	10,533	6,846	7,722	2,811
R^2	0.150	0.167	0.170	0.142

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

Relationship Choice - California left out of sample

TABLE F.3
OLS regression. Observation: first and second relationships. California dropped from initial sample

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
Unilateral Divorce	-0.062^{***}	-0.082^{***}	-0.068^{***}	-0.067^{*}
	(0.021)	(0.022)	(0.025)	(0.039)
State Fixed effects	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	9,699	6,206	7,070	2,629
R^2	0.142	0.162	0.156	0.143

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

Relationship Choice - Heterogeneity by property regime and California left out of sample

TABLE F.4
OLS regression. Observation: first and second relationships. California dropped from initial sample

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
UnDiv*NoTit	-0.068^{***}	-0.083^{***}	-0.077^{***}	-0.068^{*}
	(0.021)	(0.022)	(0.025)	(0.041)
UnDiv*Tit	-0.017	-0.057	-0.019	-0.040
	(0.032)	(0.038)	(0.040)	(0.048)
Tit	-0.010	-0.011	-0.003	-0.024
	(0.021)	(0.027)	(0.025)	(0.036)
State Fixed effects	Yes	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	9,699	6,206	7,070	2,629
R^2	0.142	0.162	0.156	0.143

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

Relationship Choice - Multinomial Logit

The empirical analysis on relationship choice that we conducted so far relied on a sample of newly formed partnerships. This implies that we studied the choice between marriage and cohabitation *conditionally* on starting a partnership. Here we provide a more complete analysis by studying the choice of singles, who can decide every month to stay single, cohabit or to marry. We do so by estimating a multinomial logit model on person month data, constructed using the first relationships of respondents of the NSFH and NSFG surveys. We construct the singleness spells by reporting individual choices from age 15 until the moment the first relationship (if it exists) begins. The results are reported in table F.5 show how the introduction of unilateral divorce impacts the relative risk of cohabiting with respect to marrying. The results confirm that unilateral divorce increased the likelihood that individuals cohabit as opposed to marrying, and that the effect is larger in non title based states. Moreover, the size of the effect is also similar to the results presented in the main text. In fact, an increase in the relative risk of cohabitation with respect to marriage of 30% (equivalent to a relative risk of 1.3) starting from an average number of first relationships that are cohabitations when the law changes of 30%, is equivalent to a decrease in the number of marriages of 9%. Interestingly, the estimated multinomial logit implies a smaller (2%-6%, not reported in the table) and non significant increase in the risk of staying single with respect to marrying.

TABLE F.5
Multinomial Logit. Observation: person month, the choices are: staying single, marry or cohabit

	Dependent variable: Relative risk of cohabiting wrt to marrying			
	Full Sample	Resident	Full Sample	Resident
	(1)	(2)	(3)	(4)
UnDiv	1.236**	1.310**		
	(0.106)	(0.141)		
UnDiv*NoTit			1.250**	1.325**
			(0.115)	(0.153)
UnDiv*Tit			1.130	1.313
			(0.179)	(0.273)
Tit			1.074	1.090
			(0.090)	(0.115)
State Fixed effects	Yes	Yes	Yes	Yes
Duration Polynomials	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes
Observations	112,697	70,882	112,697	70,882

NOTES: standard errors are clustered at the state level. The numbers displayed in the table are the exp of the coefficient of the multinomial logit and they indicate the relative risk of cohabiting with respect to marrying. When these numbers are larger than 1 the regressor of interest is associated with a relative increase in the risk of cohabiting. Relative risks that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

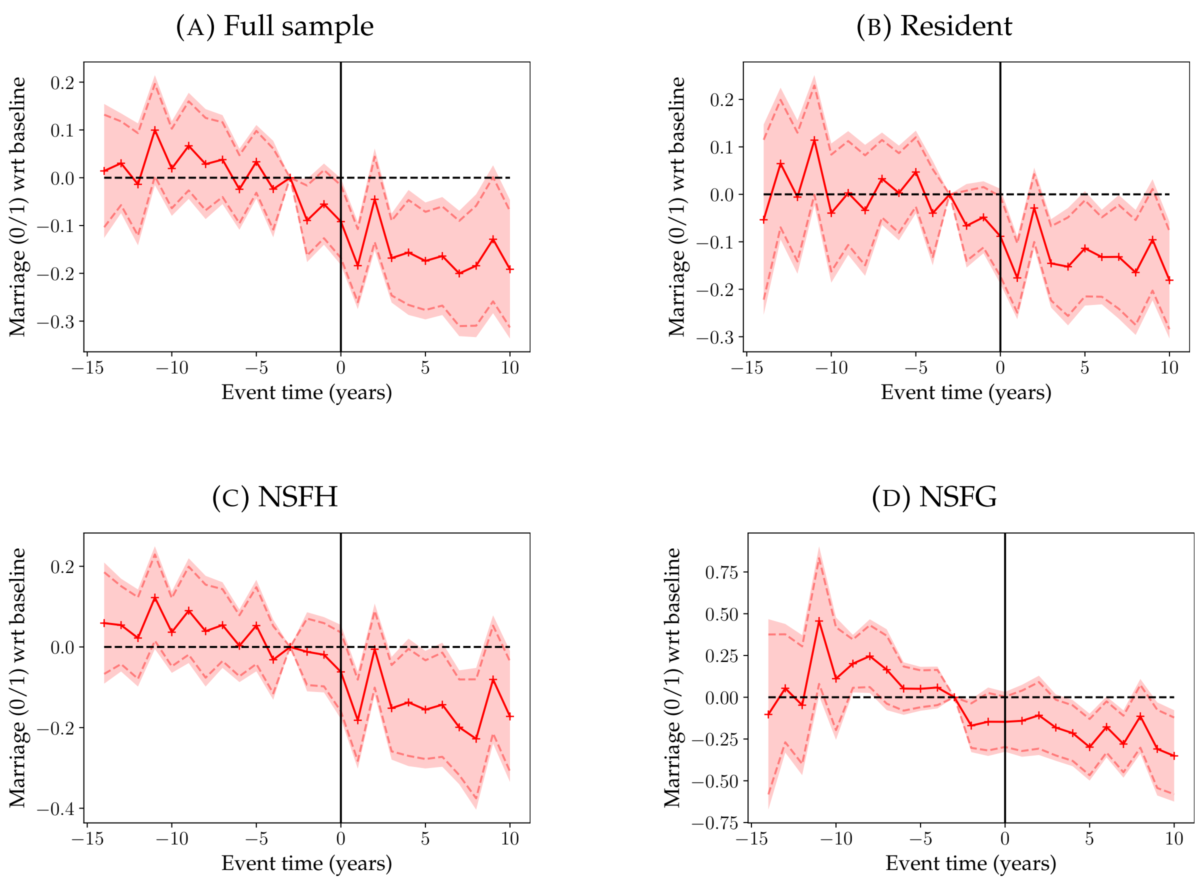
Relationship Choice - Event Study

Two-way fixed effect regressions have been widely used in the literature on the effects of unilateral divorce on economic outcomes. Yet, a recent work by Goodman-Bacon (2018) shows that this technique is problematic for recovering causal effects when treatment effects are heterogeneous across time. We follow his recommendation to use an event study design as a robustness check for the results in the main text. Specifically, we restrict the sample to states that passed the unilateral divorce law before 1988 to estimate the following equation

$$\text{married}_{i,s,y,b} = \sum_{j=-32}^{55} \beta_j^{Unid} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_b + \alpha_y + \alpha_s + \gamma' \mathbf{Z_i} + \epsilon_{i,s,y,b}, \tag{33}$$

where i stands for the respondent/household, y for the month the relationship starts, s is the state related to the household, b is the age at birth of the respondent and $\mathbf{Z}_i$ contains some characteristics of the respondent. Figure F.1 plots β_j^{Unid} for the different samples used for the estimation, where $j = -3$ is used as the reference. The results show no pretends and a significant reduction in the share of couples that cohabit after unilateral divorce is introduced: the size of the effect is larger than the one that comes from the two way fixed effect estimates. This might be due to the fact that the two weight fixed effect estimates over-weights observations that are closer to the policy change, as Goodman-Bacon (2018) notice. This matter for our case as individuals might not recall the exact date at which the couple started living together. This is confirmed by the fact that the results reported in figure F.1 show that β_j^{Unid} are negative (even though not significant) 1/2 years before the policy changes.

FIGURE F.1
Event studies on share of couples choosing marriage instead of cohabitation, around the introduction of unilateral divorce



NOTES. The figure plots the coefficients β_j^{Unid} obtained from equation 33. Each panel shows the estimates obtained using one of our four samples. The red area around the lines indicates the 95% confidence interval, while the dotted line indicates the 90% confidence interval.

Relationship Choice and Children

Is there a shift towards cohabitation for both childless couples and couples with children? Using the NSFH sample, in table F.6 below we show that unilateral divorce is associated with a shift towards cohabitation both for couples whose respondent has some children, is childless or is childless and does not want to have children. The shift towards cohabitation is lower in absolute value for couples whose respondent has children. This can be due to the fact that marriage is the best partnership for enhancing cooperation, which is a desirable feature for couples with children since these require large investments in terms of time and money. This heterogeneity is captured by our model: couples with a high relationship quality will be less sensitive to the changes in the law and are also more likely to produce large quantities of the public good, of which children are a part.

TABLE F.6
OLS regression. Observation: first and second relationships.

	Dependent variable: Married (0/1)		
	Some children	Childless	Childless+Do not want children
	(1)	(2)	(3)
Unilateral Divorce	-0.077^{***}	-0.108^{**}	-0.117^{**}
	(0.025)	(0.053)	(0.053)
State Fixed effects	Yes	Yes	Yes
Year established Fixed Effect	Yes	Yes	Yes
Birth Year dummies	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes
Observations	7,722	1,868	1,623
R^2	0.163	0.202	0.220

NOTES: standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by the following system: *10%, **5% and ***1%.

G Model Fit

FIGURE G.1
Hazards by duration of spells: data and simulations

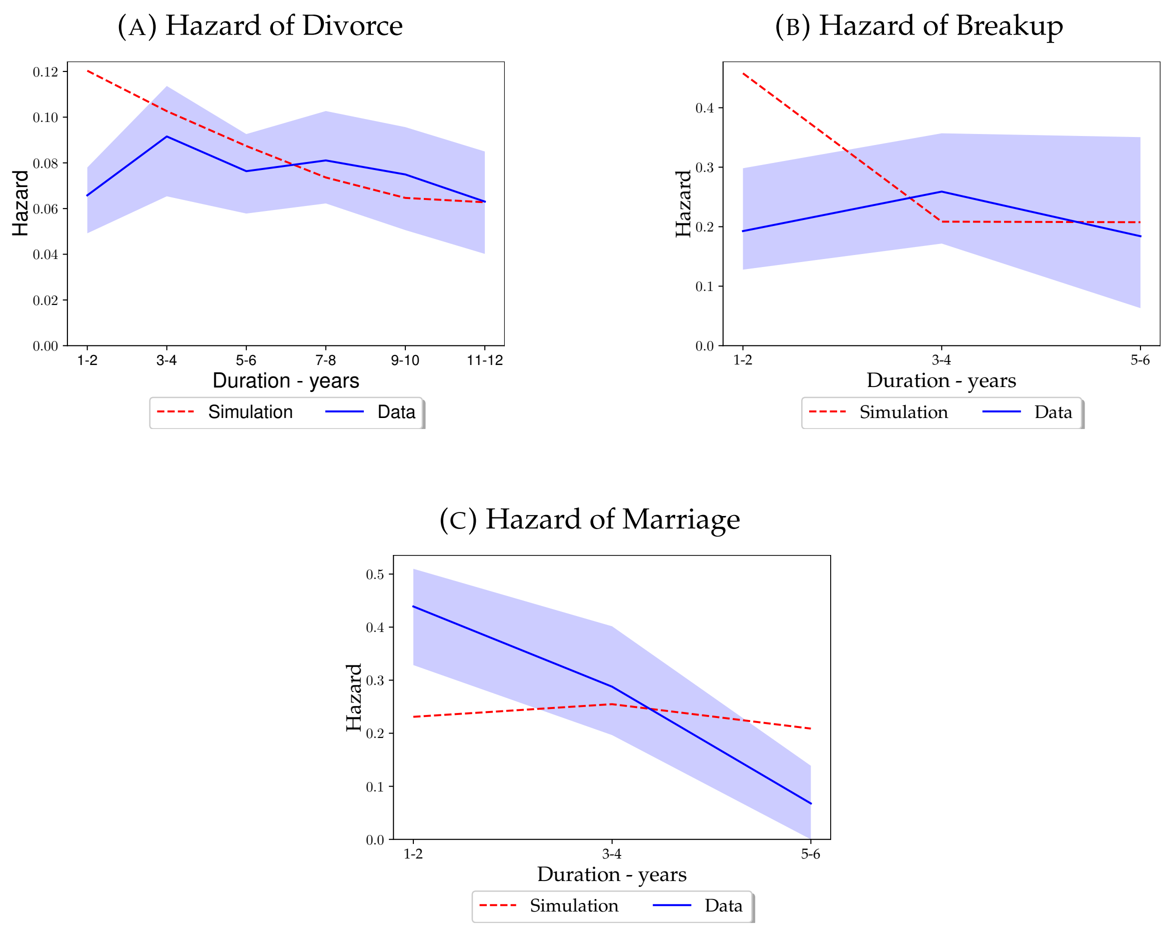


FIGURE G.2
Share ever cohabited and married: data and simulations

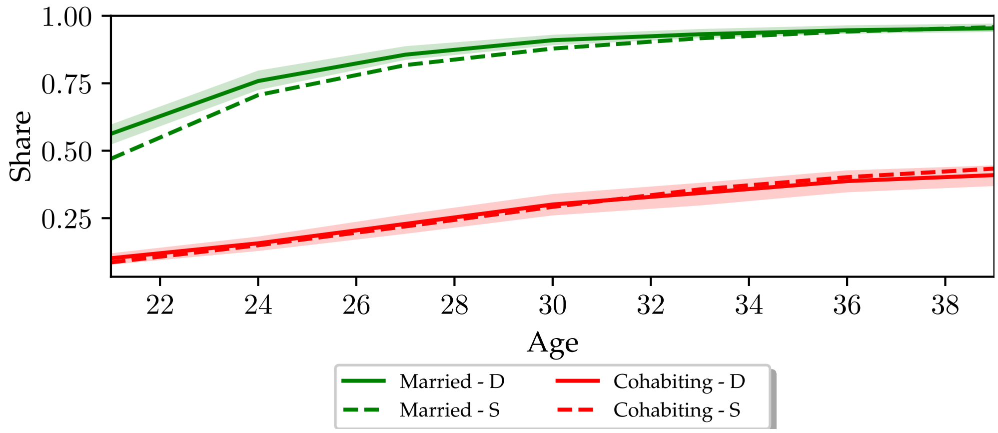
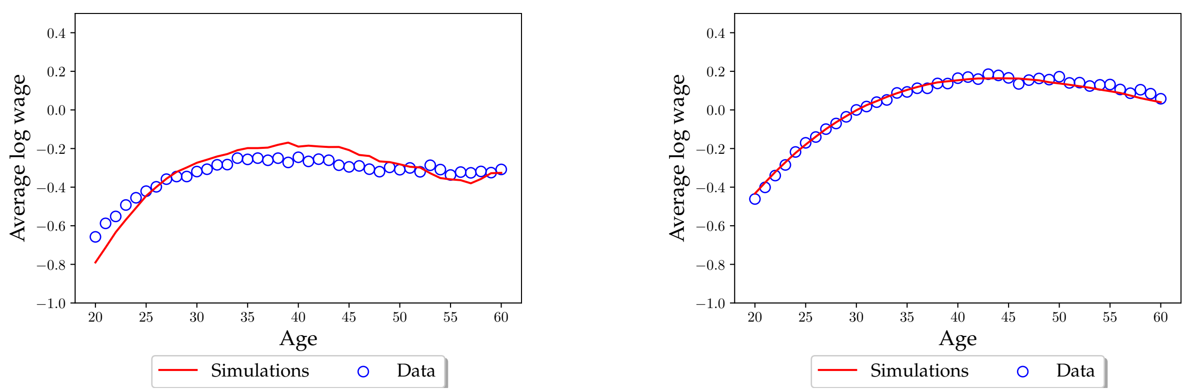


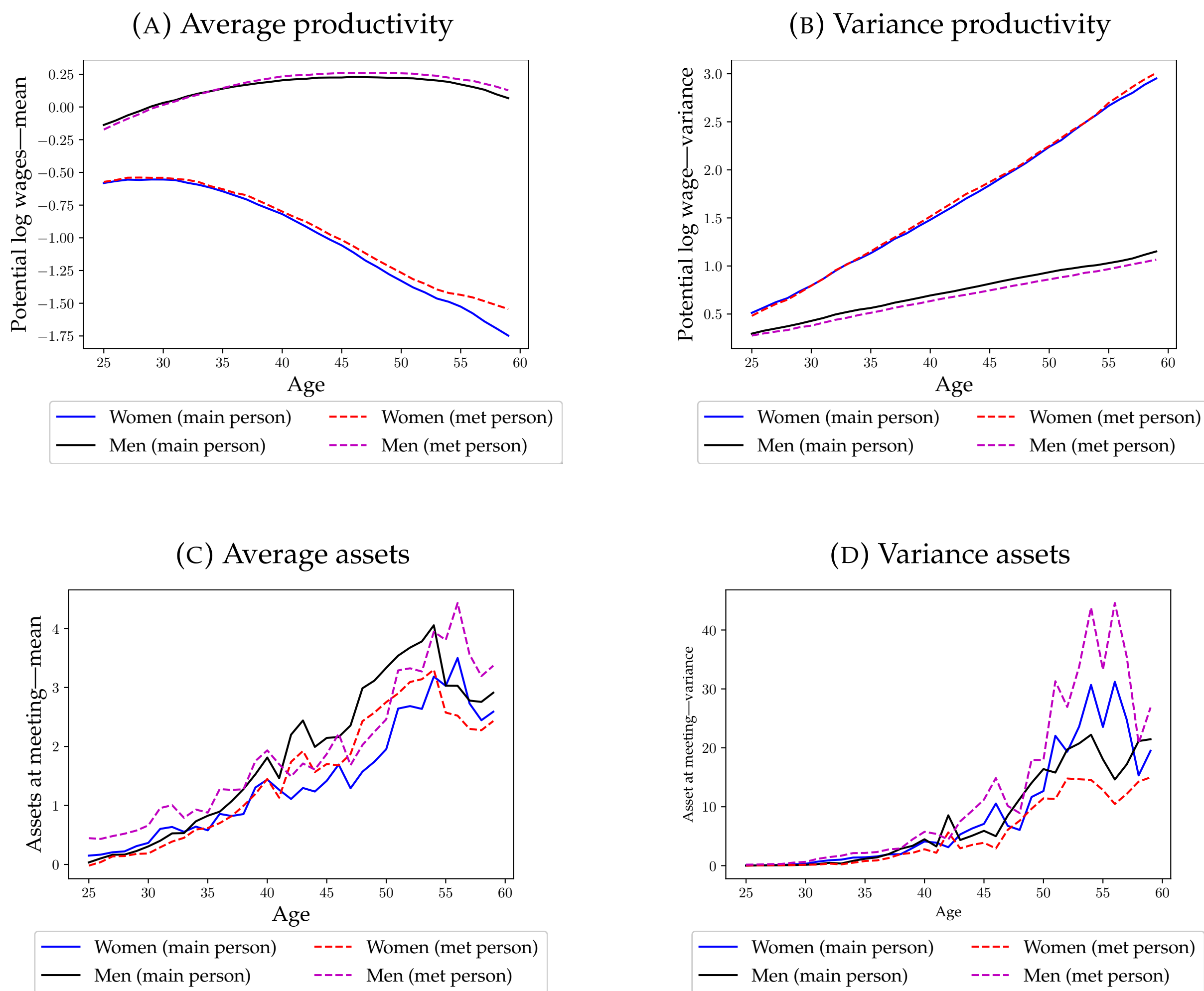
FIGURE G.3—Low wages over the life cycle: simulations and data



NOTES. This figure depicts simulated and empirical low wages over the life cycle. Data on wages are constructed by dividing the annual labor income by the total number of hours.

H Additional Figures and Tables

FIGURE H.1
Log Income and assets mean and variances by age—simulated data



NOTES. The figures display means and variances of simulated log wages and assets of men and women in a couple over their lifespan. We label as "main person" the variables that are computed from agents that are simulated and followed through their whole life-cycle, while we label as "met person" the variables constructed using the partners met by the people whose behavior is simulated for their whole life-cycle. Wage variables are constructed using couples at any point of their relationship, while for assets we use only the period when the couple met, where we can still distinguish the title of ownership of assets.

FIGURE H.2
Cumulative distribution of love shock ψ at meeting

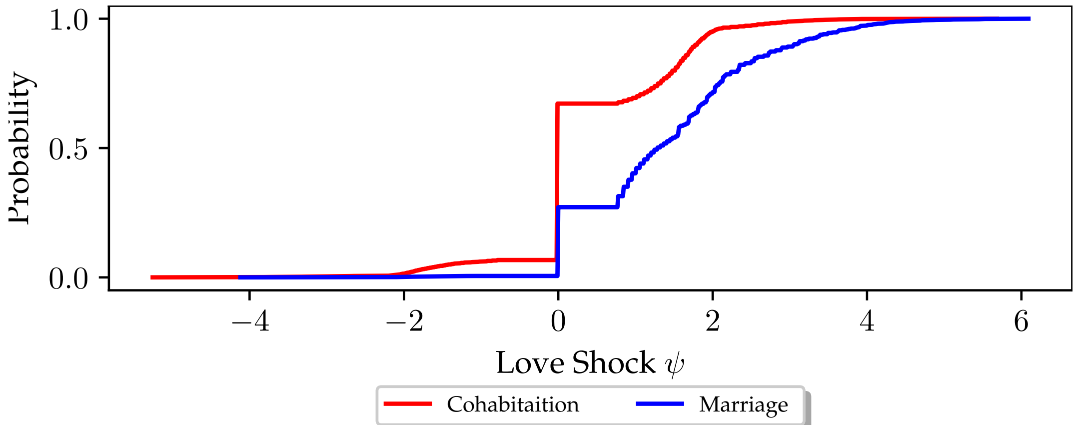
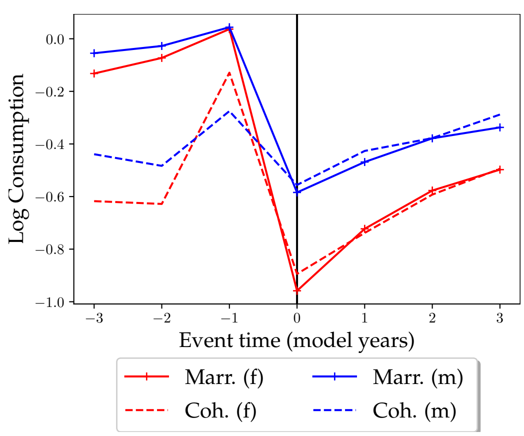
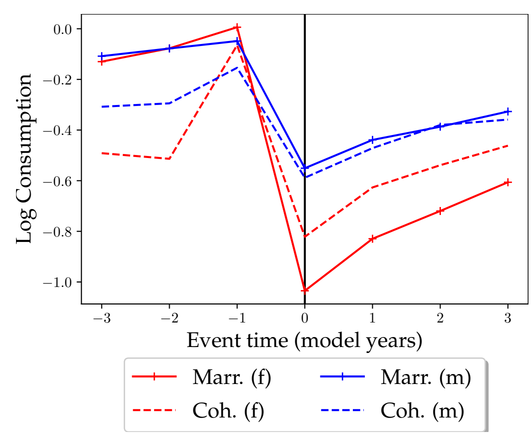


FIGURE H.3
Event studies of log consumption around divorce–simulated data

(A) Log consumption—unilateral divorce regime



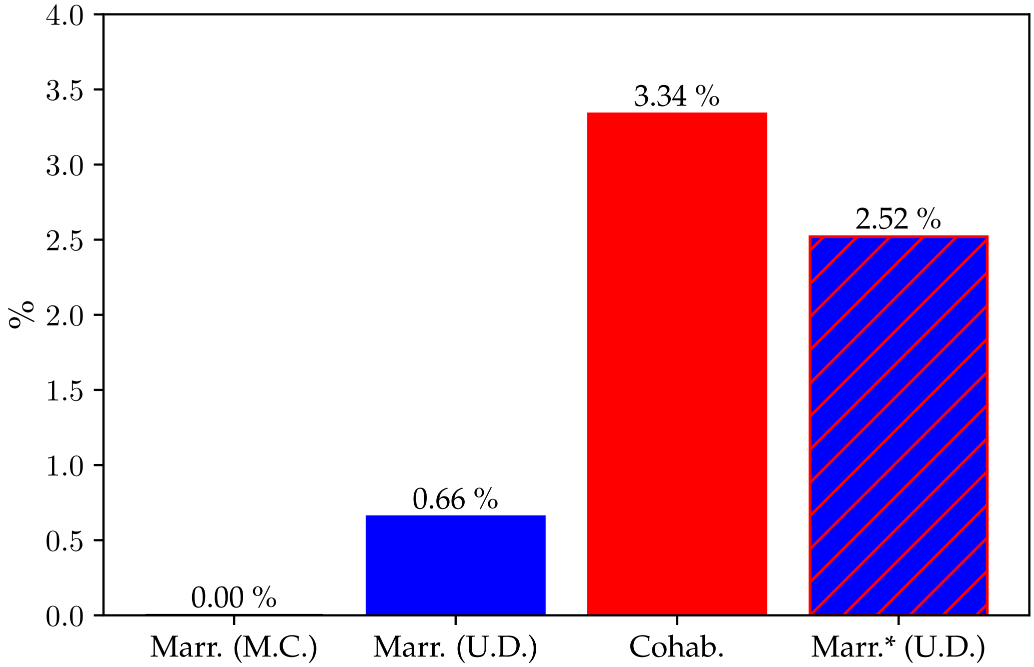
(B) Log consumption–mutual consent regime



NOTES. The figures display the evolution of simulated consumption around divorce and breakup. The displayed patterns are normalized coefficients from event studies around divorce/breakup.)

FIGURE H.4

% of periods t for which $\theta_t \neq \theta_{t+1}$



NOTES. The figures display the share of consecutive periods where the bargaining power in the couple changes. The abbreviation M.C. means mutual consent regime, the abbreviation U.D. stands for unilateral divorce regime. The values are computed using the first simulated partnership of individuals who spent their lives under the same divorce regime. The asterisk $*$ means that these marriages are obtained by imposing marriage (under unilateral divorce regime) as a partnership on couples that had decided to cohabit.

TABLE H.1

Partnership type and consumption insurance against income shocks

	Married and Cohabiting Women		
	Married		Cohabiting
	M.C.	U.D.	
Baseline	0.201	0.164	0.329
Only marriages preceded by cohabitation	0.097	0.183	-
Only marriages not preceded by cohabitation	0.210	0.160	-
Marriages with cohabitation selection	0.329	0.331	-

NOTES: the table reports the estimates of coefficients μ obtained from regression

$$\Delta \log c_{it} = \alpha + \mu \Delta \log(w_{it}) + \nu_t + \epsilon_{it}.$$

The sample includes the whole duration of the first relationship of simulated women i. The last row is run on a sample of women who decided to cohabit but we imposed marriage on them instead. This allows us to analyze the insurance within marriage controlling for selection into a relationship.