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Three Essays on Family Economics and Economic History

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To Raposa

Aknowledgemnts

I am sitting on my terrace in Brussels, searching for inspiration. While I am looking at the same threes for the umpteenth time, many memories come to my mind. Each time my eyes stopped on these trees, I was a different person: the people, the ideas, and the experiences of these Ph.D. years changed who I am, helping me mature both as a researcher and as a person. Now I would like to thank those standing by me during this journey, helping me directly or indirectly carry on my research.

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Contents

General Introduction

Why are some countries richer than others? What are the causes of economic inequalities within nations? These two important but different questions share one (partial) answer. Human capital has been shown to play a crucial role in determining the take-off to growth (Galor and Weil, 2000), but it also explains why college graduates earn higher incomes than noncollege-educated individuals (Goldin and Katz, 2009).

The determinants of human capital accumulation are multiple, but the role of families and institutions are vital. Parents play a crucial role in determining the flourishing of their children's cognitive and noncognitive abilities (Cunha et al., 2010). Schools and universities are institutions devoted to the accumulation and diffusion of knowledge.

Families and institutions sometimes create a positive feedback loop leading to human flourishing, while other times the feedback breaks down, leading to poverty and inequalities. This happened in the past and is happening nowadays. Nowadays, changes in family structure are affecting human capital accumulation. One example is the rise of unmarried cohabitation: its instability contributed to the upsurge 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). In the present, like in the past, some governments are protecting their interest by building institutions such as thought control and censorship. For example, in early modern times, the Catholic Church fought scholars and their novel ideas to maintain power.

Motivated by their relevance for the accumulation of knowledge and human capital, in this dissertation I study the determinants of cohabitation and the role of the Catholic Church's censorship in the demise of knowledge production in early modern Italy.

In the first chapter, I analyze the determinants of cohabitation in the US by education. I first document that: *i*) among singles, college graduates are less likely to cohabit than non-college educated individuals; *ii*) among cohabitors, college graduates are more likely to marry. I then provide a theoretical rationale for these stylized facts, building a life-cycle model of partnership formation. In the model, cohabitation can be both an investment good, useful to learn about the quality of prospective marriage partners, and a consumption good, namely a cheap substitute to marriage. A structural estimation of this model highlights the role of labor market earnings. Their composition accounts for most of the differential likelihood to cohabit and to marry of people with different education levels, by influencing their demand for commitment.

In the second chapter (coauthored with David de la Croix), we propose a new method to measure the effect of censorship on knowledge growth, accounting for the endogenous selection of agents into compliant vs. non-compliant ideas. We apply our method to the Catholic Church's censorship of books written by members of Italian universities and academies over the period 1400-1750. We highlight two new facts: once censorship was introduced, censored authors were of better quality than the non-censored authors, but this gap shrunk over time; the intensity of censorship decreased over time. These facts are used to identify the deep parameters of a model linking censorship to knowledge diffusion and occupational choice. We conclude that censorship reduced by 27% the average log publication per scholar in Italy. Interestingly, half of this drop stems from the induced reallocation of talents towards compliant activities, while the other half arises from the direct effect of censorship on book availability. One may wonder whether the Church's censorship also had a role in the *economic* decline of Italy. This possibility does not seem implausible, given that recent research highlighted the role of upper-tail human capital production for pre-industrial Europe's take-off (Squicciarini and Voigtländer, 2015; Cantoni and Yuchtman, 2014; Mokyr, 2011, 2016). Our analysis sets the stage for future research on this topic by directly linking the Church's censorship to upper-tail human capital production.

In the third chapter (coauthored with Egor Kozlov), we focus on 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.

Chapter 1

Cohabitation *vs.* Marriage: Mating Strategies by Education in the USA[1]

[1]The author wishes to thank David de la Croix, Fabio Mariani, Stefania Albanesi, Pierre-André Chiappori, Matthias Doepke, Fabian Kindermann, Rigas Oikonomou and Alessandra Voena for their useful comments. Egor Kozlov provided invaluable help for the development of the computational part. The author acknowledges financial support from the French speaking community of Belgium (ARC project 15/19-063 on *family transformations* and *mandat d' aspirant* FC 23613).

1.1 Introduction

The structure of American households is highly polarized. People without a college degree marry at a lower rate while experiencing a higher likelihood of divorce. At the same time, they are increasingly forming cohabiting unions, which are less stable than marriages, they are often repeated over the life cycle with different partners (Lichter et al., 2010), and they have been associated with the increasing number of single mothers (Bumpass and Lu, 2000). Among college graduates, cohabitation seems to work differently, being often followed by marriage. Understanding the mechanisms behind the different cohabitation patterns by education can help understand major life events such as marriage and divorce. For example, if premarital cohabitation is used to learn about prospective marriage partners' quality, it can lead to a low divorce risk. If, instead, premarital cohabitation signals a low initial level of commitment, it is most likely associated with a high divorce risk.

According to the seminal work of Becker (1991), people marry both for non-economic (i.e. love and companionship) and economic reasons, which include the sharing of public goods, the division of labor (Chiappori, 1997), and risk-sharing (Voena, 2015; Oikonomou and Siegel, 2015). This literature contributes to the explanation for why couples decide to live together, but it is silent about the choice between living under the same roof in a love relationship (henceforth cohabitation) and marrying. Can economic incentives also help explain the diverging patterns of marriage and cohabitation? Is cohabitation important for understanding the polarized structure of American households?

This paper addresses these questions by focusing on how economic incentives can explain the different mating strategies by education. I focus on the role of learning about the match quality and on its interaction with specialization in home production and risk-sharing. I first document that: i) among singles, college graduates are less likely to cohabit than noncollege-educated individuals; ii) among cohabitors, college graduates are more likely to marry. Then, I provide a theoretical rationale for the different mating behaviors by education, building a theory of partnership choice and endogenous dissolution. The main trade-off is that marriage enables better consumption insurance and labor specialization than cohabitation, thanks to divorce that, being costly, acts as a commitment device. Cohabitation is cheaper to dissolve, making it a preferred option for couples with the highest dissolution risk. Since a low relationship quality is associated with a higher instability of the couple, only well-matched partners select into marriage. In this framework, there is a logic to premarital cohabitation because it allows learning about the match quality, which modifies the risk of divorce and the couple's ability to cooperate and share risk. A structural estimation of this model highlights how heterogeneous income dynamics can trigger different mating strategies by influencing the demand for commitment. Since college graduates' income process triggers a larger demand for commitment, they use cohabitation as an investment good to gather information about the partner they eventually marry. Those with fewer years of education substitute marriage with serial cohabitation, which avoids the danger of a costly divorce while still delivering the gains of living together. This paper's main contribution is to show that a simple structural model of partnership choice can quantitatively account for the main patterns in partnership formation and dissolution and most of the differences in mating strategies by education. Counterfactual experiments show that changes in income dynamics over time contribute to explain the evolution of cohabitation, marriage, and divorce patterns.

I first document how mating strategies differ by education using data from the National Longitudinal Survey of Youth 1997. Using duration models in which I control for several confounding factors,[2] I show that: *i*) among singles, college graduates are less likely to cohabit (but not to marry) than noncollege-educated individuals; *ii*) among cohabitors, college graduates are more likely to marry. As in Jovanovic (1979), Brien et al. (2006) and Marinescu (2016), I find that the hazard of divorce is hump-shaped over the duration of the marriage and interpret this as evidence of the role of learning about the match quality. To further support the role of learning, I show that the longer the couples cohabit before marrying, the lower is their risk of divorce. By contrast, couples who married directly experienced a lower risk of divorce than those who cohabited for a short period. This evidence is consistent with the work of Lillard et al. (1995), who claim that better-matched couples select into marriage without first cohabiting.

To understand the mechanisms that drive the education-based differences in mating strategies, I build a dynamic model of intrahousehold decision making and search in the mating market. Agents differ in their education endowments (college or noncollege-educated) and

[2] I control for a range of socioeconomic characteristics that drive partnership choices. In particular, I control for religious affiliation and religiosity, which is linked to the likelihood of marrying according to Thornton et al. (2008). Partnership choice might be linked to religiosity through the stigma towards premarital sex (Fernández-Villaverde et al., 2014). More importantly, Mariani (2012) shows that the prevalence of premarital chastity, which is incompatible with cohabitation, increases with social status under some conditions.

make decisions according to the realization of idiosyncratic income shocks and perceived match quality, in the spirit of Jovanovic (1979). Single agents meet, with a certain probability, a potential partner with an imperfectly observed match quality. After the draw, they can decide whether to stay single, cohabit, or marry. Decision making process is modeled following the literature on limited commitment (see Pavoni et al. (2018) and Marcet and Marimon (2019)), which has been recently applied to dynamic inter-temporal household decisions by Oikonomou and Siegel (2015), Voena (2015), Low et al. (2018), Mazzocco (2007) and Foerster (2020), among others. In this framework, when one member of the couple wishes to split (for instance, because of a bad match quality draw), her bargaining power increases to make her indifferent between staying or leaving. When participation constraints cannot be met, a divorce or a breakup happens. Female labor force participation is endogenous, and women face a productivity penalty for not working. Public goods can be bought in the market or can be produced using female time. Public goods capture utility gains from children, durable goods, and services. The only difference between marriage and cohabitation is that the cost of divorce is higher than that of breakup. This is consistent with the evidence that divorce is more costly than breakup (Blasutto, 2020; Avellar and Smock, 2005) and that financial allocations upon divorce/breakup are the remaining differences between marriage and cohabitation (Lafortune and Low, 2020). Also, a high cost of divorce implies that more committed couples self-select into marriage, as suggested by sociologists (Poortman and Mills, 2012).

I do a structural estimation of the model to understand the quantitative relevance of the mechanisms that drive partnership choice. The model is estimated using the method of simulated moments to reflect a realistic mating market and female labor supply. With relatively few parameters, the model can match well all the targeted moments. In particular, a high divorce cost is important to match: *i*) the prevalence of cohabitation and marriage; *ii*) the higher instability of cohabitation than marriage; *iii*) the larger female labor supply of cohabitors. The results show that the model can also reproduce most of the educational gradient in marriage, cohabitation, and divorce without targeting these differences in the estimation. The ratio of college over non-college ever married (cohabited) at 35 is 1.18 in the data and 1.13 in the simulations (0.86 in the data and 0.55. in the simulations). I did not expect to match all the gradient by education since the model abstract from other mechanisms driving it. I then run a series of counterfactual experiments to understand the quantitative importance of the forces contributing to the different mating strategies by education. I find that the gender wage gap is more important than income volatility for explaining the differential mating strategies. Since the income process of college graduates is subject to a larger wage gap,[3] they value commitment more. When the gender wage gap is large, there is more scope for specialization in the household, which can be better exploited within marriage. Hence, college graduates prefer to marry, using cohabitation as an investment good to gather information about the partner they eventually marry. These mechanisms also contribute to explain the changes in marriage, cohabitation, and divorce over time. In particular, changes in the income process over the last thirty years can explain $\sim 100\%$ ($\sim 30\%$) of the evolution in the marriage (cohabitation) gap by education. These results are mainly driven by the reduction of the gender wage gap: if it were closed completely, the share of people ever married (ever cohabited) by the age of 35 would drop by 25% (increase by 19%). Finally, the role of learning and selection are also quantitatively relevant: if cohabitation were not accessible, marriages would be significantly more unstable.

This paper adds to three strands of the literature. First, this paper adds to the literature that studies the choice between marriage and cohabitation. I am the first to develop a dynamic structural model of partnership choice where: (*i*) opting for marriage or cohabitation is fully endogenous; (*ii*) patterns of cohabitation, marriage, and divorce by education are quantitatively accounted for. Different papers in the literature highlighted various advantages of marriage relative to cohabitation: commitment (Matouschek and Rasul, 2008; Blasutto, 2020), labor specialization within the couple (Gemici and Laufer, 2014), learning about match quality (Brien et al., 2006), and investment in children (Lafortune and Low, 2020). Gemici and Laufer (2014) were the first to develop a structural model of marriage and cohabitation where limited commitment triggers the gains of marrying. In their model, the marriage-cohabitation choice is not fully endogenous, and they abstract from persistent income shocks. This paper shows that persistent income shocks plays a role in partnership strategies when coupled with learning. Lafortune and Low (2020) highlight the role of assets as collateral that enforces commitment and allows for optimal investment in children within marriage. My model captures the idea that divorce acts as a commitment technology to enforce specialization within the household and consumption insurance. Brien et al.

[3]Cortés and Pan (2019) show that the gender pay gap declined more slowly at the top of the distribution.

(2006) develop a where agents cohabit to learn about the match quality, but they do not address the interaction between learning and commitment, and the cohabitation-marriage choice is not fully endogenous.

Second, this paper is tied to the extensive literature that studies the changes in the character of the American household over the last decades. Demographers and sociologists are the first to document the diverging patterns of partnership choice by education. Among others, Bumpass and Lu (2000) describe the uneven rise in cohabitation and its effect on children. Lichter et al. (2010) observe the emergence of serial cohabitation, which is particularly strong among disadvantaged populations. Perelli-Harris and Lyons-Amos (2016) show evidence of an educational divergence in marriage and union dissolution. In economics, various studies explored the role of education, technology, and the distribution of income to study the changes in household formation and dissolution (Chiappori et al. 2017; Greenwood et al. 2016; Ciscato 2019 among others), but they abstract from cohabitation. An exception is Lundberg et al. (2016), arguing that the socioeconomic gradient in partnership formation and dissolution might have emerged because of the uneven demand for marriage as a commitment technology, that allows intensive joint investment in children. Compared to these works, I show that the structure of income processes, interacting with learning and the demand for commitment, accounts for most of the gradient in mating patterns by education and (partially) their evolution over time. In particular, this paper is the first to show that the dynamics of the gender wage gap and the income process's volatility have a quantitatively important role in explaining the uneven retreat from marriage.

Third, my paper also relates to papers exploring the role of learning about match quality (Marinescu, 2016) and selection into marriage and premarital cohabitation (Lillard et al., 1995; Reinhold, 2010; Svarer, 2004). I extend their work allowing for the possibility that premarital cohabitation can have a role in marriage stability both through the extensive margin (linked to selection) and the intensive margin (linked to learning). Thanks to this distinction, I can claim that both selection into marriage and learning play are quantitatively important in explaining the relationship between premarital cohabitation and divorce.

The remainder of this paper is organized as follows. Section 2 documents the different mating strategies by education and the role of premarital cohabitation for divorce. Section 3 presents and develops the theoretical framework. Section 4 describes the estimation of the model. Section 5 presents a series of counterfactual experiments and the role of income dynamics for the changes in marriage and cohabitation patterns over time. Section 6 provides additional evidence about the model's mechanisms, while section 7 contains the conclusion.

1.2 Data and Stylized Facts

1.2.1 Sample

I use data from the "National Longitudinal Survey of Youth 97" (henceforth NLSY97), a national longitudinal survey of 8984 men and women born in the USA between 1980 and 1984. Participants were interviewed annually between 1997 and 2011, then biennially until 2017. I use this data for two main reasons. First, it reports the history of marriage and cohabitation with monthly frequency, thus capturing short cohabitation spells.[4] Second, marital and cohabitation status is checked at each wave, thus avoiding a recall bias.[5] Cohabitation is defined as marriage-like sexual relationship in which partners live together without being formally married. For my analysis, I select a sub-sample of the NLSY97 comprising all the observations for which I have complete information about the marital history, and the information about education is non-missing. A shortcoming of this data is that individuals are never observed while they are forty or older. Yet, a significant number of respondents have experienced cohabitation and marriage already. Over 60% (70%) of respondents had been married (cohabiting) at the time of the last survey's wave. It is not uncommon to use samples of young people to examine cohabitation. For instance, Brien et al. (2006) study cohabitation choices using a sample of women who are followed until age 32. My analysis targets understanding the behavior of the young, for whom cohabitation is very common. Table 1.1 provides some summary statistics of individuals, cohabitation, marriage, and singleness spells. One spell is defined as the period of time spent in a particular state (i.e. marriage, cohabitation or singleness). Notably, cohabitation spells are much shorter than marriages and many cohabitations become marriages.[6]

[4]Indeed 31% of the spells of cohabitation in my sample are shorter than twelve months.

[5]Hayford and Morgan (2008) study the quality of retrospective data on cohabitation using the National Survey of Family Growth and the National Survey of Family and the Household. They find that the recall bias leads to an underestimation of cohabitation rates.

[6]The duration reported in table 1.1 does not account for the fact that spells are censored. Hazard rates, reported in figure 3.H.1, account for this fact and show that cohabitations are much more unstable than marriages.

Table 1.1: Summary statistics

Variable	Mean	Std. Dev.	N
Individuals			
College (%)	0.264	0.441	6674
Female (%)	0.515	0.500	6674
Marriage Spells			
Duration (years)	7.720	4.527	4260
Ending in Divorce (%)	0.200	0.400	4260
Cohabitation Spells			
Duration (years)	3.300	3.065	7903
Ending in Marriage (%)	0.352	0.478	7903
Ending in Separation (%)	0.424	0.494	7903
Singleness Spells			
Duration (years)	7.863	5.378	9443
Ending in Marriage (%)	0.127	0.333	9443
Ending in Cohabitation (%)	0.734	0.442	9443

1.2.2 Facts

This subsection presents evidence that mating behavior differs significantly by education. Even though the sociological and demographic literature has already provided some evidence that serial cohabitation is more common among the least educated (Bumpass and Lu, 2000; Lichter et al., 2010; Perelli-Harris and Lyons-Amos, 2016), I present these differences also in this paper. I do that for two reasons: first, to be sure that the same patterns apply to my recent data. Second, I distinguish between the partnership choices of singles versus the likelihood that cohabitants transition into marriage. Throughout the section, I refer to Sample I when the spells of interest are constructed using all respondents in the NLSY97, counting as censored observations that leave the sample, or for which marital history is not available. Sample II is the one I already described above. Appendix 1.A describes in detail the econometric models that I am using in this section: the Cox proportional hazard model (Cox, 1972) and the Fine-Gray regression (Fine and Gray, 1999). These models analyze the likelihood that an event occurs (e.g. divorce) conditional on a set of controls. Their advantage over a linear regression is that they control for right censoring: I do not observe respondents until they die. The difference between the Cox model and the Fine-Gray model is that the second considers that observations face the risks of different types of events. For example, a cohabiting couple faces both the risk of breaking up and marrying. In the case of multiple risks, the model studies one main event of interest, while other events are referred to as competing events. The results of the Fine-Gray regressions are displayed in terms of relative sub-hazard ratios. A sub-hazard is defined as the probability that the event of interest happens conditionally on not having occurred already, while a competing event might already have happened. When the ratio is larger than 1, it means that the covariate increases the likelihood that the event occurs.

Fact 1: *Among singles, college graduates are less likely to cohabit than non-graduates.*

I study partnership choices of singles using the duration model by Fine and Gray (1999). The units of analysis are singleness spells,[7] while the event of interest is cohabitation.[8] I run six regressions that differ in the sample used (sample I or II) and the set of controls. I control for several individual characteristics, including religiosity and number of past relationships: Appendix 1.C has a complete list of the controls. Full results are presented in table 1.C.1, while panel (A) of figure 1.1 plots the sub-hazard ratio of our variable of interest, namely a dummy that takes value 1 when the respondent has a college degree. The sub-hazard ratio of being a college graduate ranges between 0.54 and 0.74 according to the specification. This results show that college graduates, when single, are less likely to end up cohabiting. In table 1.C.2 I use the same econometric model, covariates, and samples to study the risk of marriage, treating cohabitation as a competing risk. The results show that, if anything, college graduates are more likely to end up marrying than non-college educated individuals. Overall, my results show that college graduates are less likely to choose cohabitation over marriage when they are single. This does not automatically imply that they marry more and cohabit less overall. It could be that college graduates never transit from cohabitation to marriage, while non-graduates do it frequently. The stylized fact I present below rules out this possibility.

Fact 2: *Among cohabitors, college graduates are more likely to marry than non-graduates.*

Another interesting difference in the mating behavior by education is the likelihood of transitioning from cohabitation to marriage. In fact, it may signals whether the partnership is considered to be permanent or temporary. While I do not observe in my sample people who transitioned from marriage to cohabitation with the same person, the opposite is quite common: 35% of cohabiting couples end up marrying. An interesting feature of the data is that the probability of transitioning from cohabitation to marriage is more frequent among

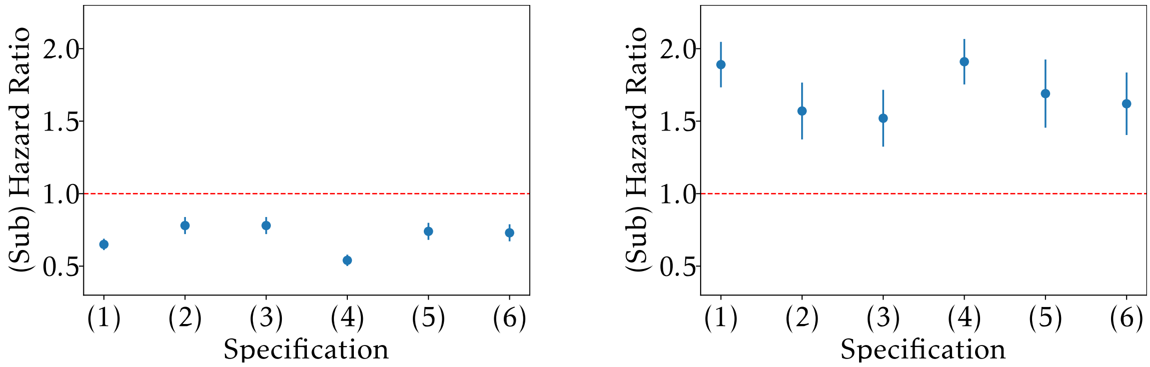
[7]These spells include both people who never experienced a relationship and those who are divorced or experienced a breakup.

[8]If a person decides to cohabit and subsequently marries the same partner, it is counted as cohabitation. I analyze the transition from cohabitation to marriage afterward.

Figure 1.1: Partnership choices by Education. Results from Duration Analysis

(a) Relative likelihood to cohabit for college over noncollege educated who are currently single

(b) Relative likelihood of marrying for college over noncollege educated who currently cohabit



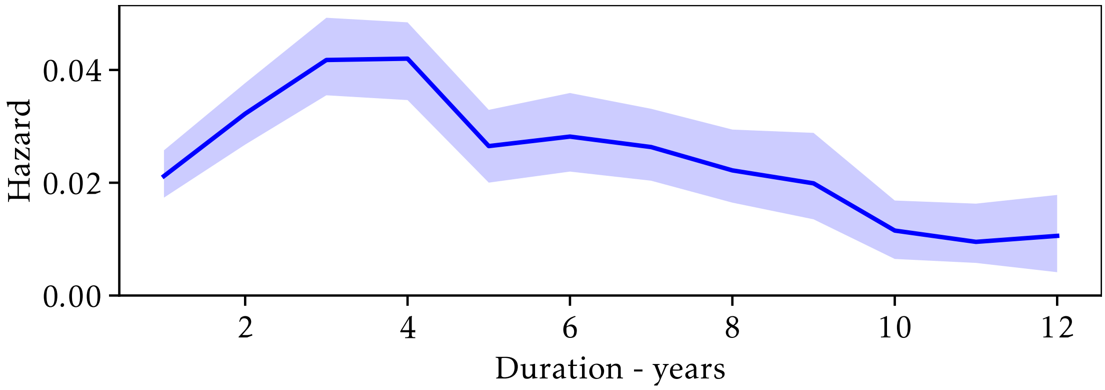
NOTES. The sub-hazard ratios are obtained using the Fine-Gray regressions for different specifications as displayed in tables 1.C.1 and 1.C.3 in the Appendix. When the ratio is larger than 1, the covariate increases the likelihood that the event occurs. This figure also reports the 95% confidence intervals for each sub-hazard ratio.

college graduates than those with fewer years of education. To show whether this difference is statistically significant, I use the Fine and Gray (1999) regression. The unit of analysis is a cohabitation spell, while the event I am interested in is marriage. I run six regressions that differ in the sample used (sample I or II) and the set of controls. Full results are presented in table 1.C.3, while panel (B) of figure 1.1 plots the sub-hazard ratio of the dummy college graduate. The sub-hazard ratio of being a college graduate ranges between 1.52 and 1.91 according to the specification. These results are consistent with the story that college graduates use cohabitation as a trial for marriage and therefore experience a transition from cohabitation to marriage more often. In appendix 1.C I propose a robustness check using a multinomial logit with spell-month observations for facts 1 and 2.

1.2.3 The Extensive and Intensive Margins of Cohabitation

In this subsection, I provide evidence that learning explains mating behavior. Jovanovic (1979) builds a model where employers and employees learn about the quality of their match. He concludes that the hazard rate of breakup is first increasing and then decreasing over time. The rationale for this result is that agents wait before splitting, since they are not sure of the match quality, which is learned over time. The hazard rate starts decreasing at a certain moment because of a self-selection effect: the relationships that have not broken up yet are the ones with the best match quality. Several works in the marriage and divorce literature (Brien et al., 2006; Marinescu, 2016) studied this pattern to provide evidence, or lack thereof, of learning. As figure 1.2 illustrates, for my sample, the hazard of divorce is first increasing and then decreasing over marriage duration.

Figure 1.2: Hazard of divorce by duration



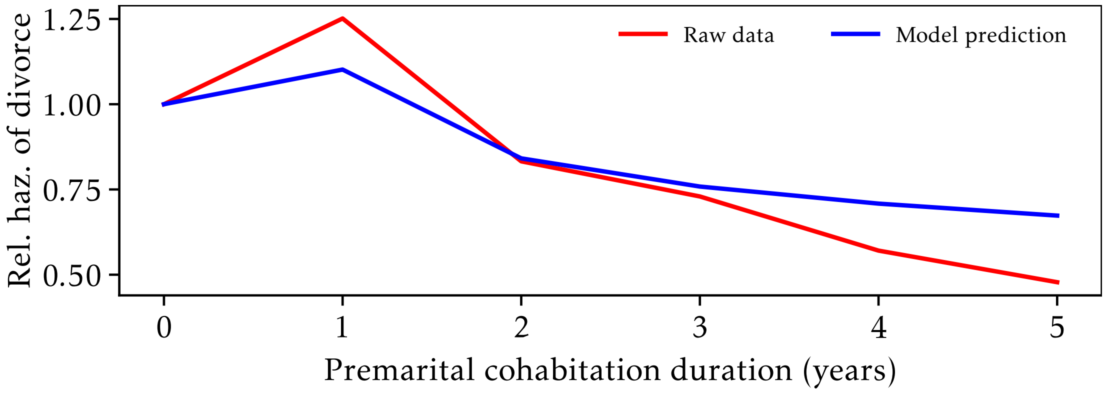
Notes. The hazard rate for each marriage year t is computed as the probability that a marriage ends in divorce in interval t, conditional on not having divorced until t.

It can still be that a composition effect also drives the shape of the hazard over time.[9] Hence, I collect additional evidence about learning. If there is learning, I expect longer periods of premarital cohabitation (intensive margin of cohabitation) to decrease the hazard of divorce. On the other hand, it might be that better matched partners marry directly, without the need to cohabit before (intensive margin of cohabitation). I might expect those people to have a lower hazard of divorce for a selection effect. I study the relative role of these margins estimating a Cox regression model. This model captures the extensive margin of cohabitation with a dummy taking value one if there was premarital cohabitation and log(Cohabitation length $+ \epsilon$) to capture the intensive margin,[10] where ϵ is a very small number. Table 1.C.4 shows the full results of six Cox regressions that differ in the sample used (sample I or II) and the set of controls. The extensive margin of cohabitation is associated with an increase in the risk of divorce in all specifications. The intensive margin is associated with a reduction in the risk of divorce. To provide a visualization of the overall effect, I plot in figure 1.3 the predicted risk of divorce given a certain number of years of cohabitation, relative to the risk of divorce of marriages with no previous cohabitation.

[9] Different distributions of people's characteristics linked to divorce can give rise to different shapes of the hazard of divorce over time.

[10] I use the log of cohabitation to capture a diminishing learning rate over time. ϵ is a small number that allows me to avoid dropping observations that did not cohabit in the sample.

Figure 1.3: Relative hazard of divorce by premarital cohabitation duration



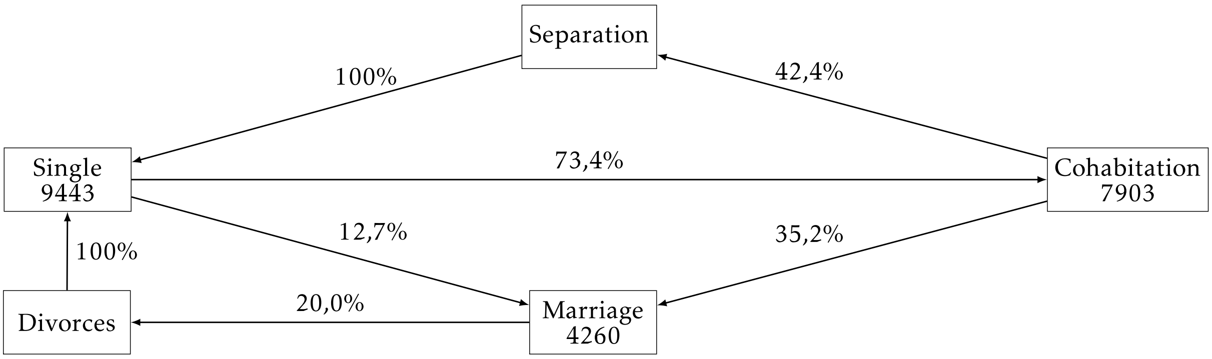
Notes. The line *Model Prediction* shows the effect of cohabitation on the hazard of divorce as derived from model (3) in table 1.C.4: couples that did not cohabit are taken as the reference. *Raw Data* represents the relative share of divorces by premarital cohabitation duration.

When premarital cohabitation increases, the hazard of divorce first increases and then decreases slowly. This fact suggests that the initial adverse self-selection effect fades out as the number of years of premarital cohabitation increases. The same graph shows the raw probability of divorce given a certain cohabitation length. In appendix 1.C I propose a robustness check using a logit with spell-month observations.

1.3 Theory

To identify the channels through which the mating strategies by education arise, I develop a dynamic life cycle model of partnership formation, female labor supply, home production, and learning about match quality. Couples act cooperatively 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, who belong either to the college group e or non-college group n, draw their productivities. If single, with some probability, they meet a potential partner. After drawing a match quality (which is imperfectly observed), modeled as a utility flow, they decide whether to marry, cohabit, or stay single. Couples draw productivity shocks and noisy observations of the match quality, and according to those, they decide whether to stay together or to split. Couples learn about match quality, making better predictions the longer they are in the relationship. Both singles and couples decide how to split their money between private consumption and public good expenditure. Couples also make female labor supply decisions and women's time can be used to produce public goods, but this comes at the cost of a loss in productivity.[11]

Figure 1.4



Notes. The number of spells and the transitions are those in the main sample.

1.3.1 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 also derive utility from match quality $\hat{\psi}_d$: in the next subsection, I detail how it evolves over a couple's duration d. The intra-period utility of a single agent of gender s is:

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

Note that the utility derived from the public and the private goods has the same shape. The utility for an agent in a couple is:

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

where the public good does not have a superscript because it is shared.

1.3.2 Match Quality

Agents in a relationship draw utility from a match quality shock $\hat{\psi}_d$, which can be interpreted as the utility value of companionship, and all the couple-specific attachment to the

[11] I assume that single women and all men work full time.

relationship.[12] I now describe how match quality evolves over relationship duration d.[13] Two singles that just met draw their long run match quality ψ from distribution $\mathcal{N}(0,\sigma^2)$. The potential couple only observe $\psi_1^f = \psi + \eta_1$, where the noise η_1 follows the distribution $\mathcal{N}(0,\sigma_\eta^2)$. Even in subsequent periods, couples observe $\psi_d^f = \psi + \eta_d$, where η_d has the same distribution for all durations d. While ψ is never observed directly, couples know this structure and can use past realizations to predict future match quality. I assume that agents use information efficiently, forming their beliefs through Bayesian learning. From these assumptions, it follows that beliefs are normally distributed,[14] with a variance $\hat{\sigma}_d^2$ that depends on duration d only, while the mean $\hat{\psi}_d$ depends on past realizations. I can write the recursion of the expectations for $d > 1$:

$$\begin{cases} \hat{\psi}_d = \hat{\psi}_{d-1} + K_d(\psi_d^f - \hat{\psi}_{d-1}) \\ \hat{\sigma}_d^2 = (1 - K_d)\hat{\sigma}_{d-1}^2 \\ K_d = \frac{\hat{\sigma}_{d-1}^2}{\hat{\sigma}_{d-1}^2 + \sigma_\eta^2}. \end{cases} \tag{1.1}$$

Note that $\hat{\psi}_1 = K_1\psi_1^f$, $\hat{\psi}_1 = (1-K_1)\sigma^2$, where $K_1 = \sigma^2/(\sigma_\eta^2+\sigma^2)$, which follows from Bayesian learning and agents knowing the distribution of ψ. Note that $\mathcal{C}_d = [\hat{\psi}_d, d]$ is a sufficient statistic for forming the expectations for match quality in $d+1$. In line with Brien et al. (2006) and Marinescu (2016), I assume that agents gain utility from $\hat{\psi}_d$ and not ψ_d^f: it is enough to capture the fact that couple's predictions become more precise as relationship duration increases, while it saves one state variable. For computational reasons I also assume that agents stop improving their beliefs in $d = 8$.

1.3.3 Home Production

In my model, each agent has one unit of time. I assume that singles and men in a couple supply inelastically a fraction $1 - \phi$ of their time to the labor market, and produce home goods with the rest of their time.[15] Women in a couple can devote their whole time endow-

[12]For instance, concerns about the well-being of children or the psychological fear of ending alone, are part of the match quality. Interestingly, religiosity is captured by this variable, being higher for couples with a high match quality. In fact, the model predicts that for the same economic incentives, couples with a high match quality are more likely to marry than to cohabit, less likely to divorce and more likely to have a stay-at-home female partner. In the NLSY97 data, people that reported to be more religious share the same characteristics.

[13]A couple transitioning from cohabitation to marriage continues its duration.

[14]When I refer to beliefs of match quality at duration d, I mean after η_d has been observed.

[15]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; Reynoso, 2020) and it is made to simplify the solution of the model, which allows me to enrich it with other features. Moreover, this assumption is

ment to producing the home good Q. The public good can also be produced by buying x goods in the market. Following Greenwood et al. (2016) I define the production function of home goods for singles of gender $s \in \{f, m\}$ as:

$$Q_t^s = [(x_t^s)^\nu + \chi\phi^\nu]^{\frac{1}{\nu}}, \text{ where } 0 < \nu < 1, \tag{1.2}$$

while for couples

$$Q_t = [x_t^\nu + \chi(2\phi + (1 - P_t^f)(1-\phi))^\nu]^{\frac{1}{\nu}}, \text{ where } 0 < \nu < 1. \tag{1.3}$$

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 (here normalized to 1) of durables decreases and when wages go up, women spend less time producing household goods and their employment outside the home increases. The variable P_t^f is a dummy variable that takes value 1 when women participate in the labor market.

1.3.4 Wages

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

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

where $f_t^{s,i}$ is a gender and education specific function that captures the evolution of productivity over age. Hence, both the gender wage gap and the financial returns to education depend on it. The permanent income component $z_t^{s,i}$ evolves over time according to:

$$z_t^{s,i} = z_{t-1}^{s,i} - (1 - P_t^{s,i})\mu + \zeta_t^{s,i}, \text{ where } \zeta_t^{s,i}\mathcal{N}(0, \sigma_\zeta^{2_{s,i}}), \text{ and } \zeta_1^{s,i} = z_1^{s,i}. \tag{1.4}$$

Parameter μ is the loss in productivity that affects women who are not participating in the labor market.[16] Importantly, the variance of the persistent shock depends on gender and

consistent with the fact that the large majority of men are in the labor force: only 8% of men between 20 and 60 were out of the labor force in the 2000s.

16. Parameter μ can be interpreted as a reduced form way of capturing i) the missed opportunity to accumulate human capital while working and ii)the skill atrophy from interruptions (Adda et al., 2017). It also captures the harsh economic situation that single mothers face: this corresponds to a wife divorcing after

education: for some groups, risk-sharing might be more important than for others.

1.3.5 Budget Constraints

Singles of gender $s \in \{f, m\}$ face the budget constraint:

$$w_t^s(1-\phi) = c_t^s + x_t^s, \tag{1.5}$$

where ϕ is the fixed time singles spend to produce the public good and w depends on gender, education, productivity shocks and age. Note that agents cannot save and borrow.[17] An agent that just divorced receives a one-shot loss in her disposable income, being left with a fraction κ of her wage.[18] The budget constraint for a couple is

$$w_t^m(1-\phi) + w_t^f P_t^f(1-\phi) = c_t^m + c_t^f + x_t. \tag{1.6}$$

1.3.6 Problem of the Singles and Divorcees

I start by describing the problem for a single agent of gender $s \in \{f, m\}$ and education $i \in \{e, n\}$ in t. In $t+1$ they meet a potential partner with probability λ_{t+1}^i and they can decide to enter a partnership, which also depends on whether the potential partner will agree. I The state variable of a single is $\omega_t^s = \{i, z_t\}$, while their choices are represented by the vector $\mathbf{q}_t = \{c_t, x_t\}$. I denote by $V_t^{S_s}(\omega_t^s)$ the value of being single and not entering in a relationship

having used her time to produce Q. It is sporadic in the model that stay-at-home female partners switch to employment while the couple is intact. As I already anticipated, $P_t^{m,i} = 1$.

[17] I abstract away from savings for three reasons. First, it allows me to keep the model computationally tractable while enriching it with learning about the match quality, which is important to understand cohabitation. Second, as Wu and Krueger (2018) and Blundell et al. (2016) point out, savings are not effective in self insuring against persistent income shocks compared to the role of the family for insurance. Moreover, as Guvenen and Rendall (2015) mention, the median liquid wealth of US households is less than \$5,000 and even less for young individuals, when most of the decisions regarding family formation and dissolution are taken. Third, while Lafortune and Low (2020) show that assets can act as a commitment technology to enforce cooperation in the family, in my model the monetary cost of divorce proportional to income plays a similar role.

[18] This assumption is consistent with the evidence that upon divorce i) wealthier people lose more asset (Blasutto, 2020) and ii) high income households in the UK experience more persistent losses in their standards of living (Fisher and Low, 2016).

in t as

$$
\begin{aligned}
&V_t^{S_s}(\omega_t^s) = \max_{\mathbf{q}_t} u(c_t, Q_t) + \beta E_t\Big\{ && (1.7)\\
&(1-\lambda_{t+1}^i)V_{t+1}^{S_s}(\omega_{t+1}^s)+ && \text{(no meeting)}\\
&\lambda_{t+1}^i(1-M_{t+1})(1-C_{t+1})V_{t+1}^{S_s}(\omega_{t+1}^s)+ && \text{(meeting, stay single)}\\
&\lambda_{t+1}^i M_{t+1} V_{t+1}^M(\Omega_{t+1})+ && \text{(meeting, marry)}\\
&\lambda_{t+1}^i C_{t+1} V_{t+1}^C(\Omega_{t+1})\Big\}, && \text{(meeting, cohabit)}\\
&\text{s.t. budget constraint of singles (3.9).}
\end{aligned}
$$

Note that if the couple 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. Variables M_{t+1} and C_{t+1} are not solely choices of the individual, but of the (potential) couple: I explain in subsection 3.4.7 how these are set. The value function of the divorcee $V_t^{D_s}$ is the same as that for singles, with the exception that in t only the agent is left with a fraction κ of their income. The vector of state variables Ω of couples is defined in the next subsection.

1.3.7 Marriage

Married couples maximize the weighted sum of the spouses' utilities under the constraint that both are better-off staying married than divorcing. This constraint implies that the Pareto weights of the wife θ_t^f and the husband θ_t^m are allowed to vary over time to meet the participation constraints. The state vector of this problem is $\Omega_t = \{i^f, i^m, z_t^f, z_t^m, d, \hat{\psi}_d, \theta_t^f, \theta_t^m\}$, where i^f and i^m are the education levels of the wife and of the husband, respectively. The variables over which the couple maximizes are summarized by $\mathbf{q}_t^M = \{x_t, c_t^f, c_t^m, P_t, D_t\}$, where D_t is a dummy variable indicating that the couple chose to divorce. The formal problem solved by a couple entering t as married is:

$$
\begin{aligned}
&V_t^M(\Omega_t) = \max_{\mathbf{q}_t^M}\Big\{ && (1.8)\\
&(1-D_t)\{\theta_t^f u(c_t^f, Q_t) + \theta_t^m u(c_t^m, Q_t) + \hat{\psi}_d + \beta E_t V_{t+1}^M(\Omega_{t+1})\}+ && \text{(stay married, } D_t = 0)\\
&D_t\{\theta_t^f V_t^{D_f}(\omega_t^f) + \theta_t^m V_t^{D_m}(\omega_t^m)\}\Big\} && \text{(divorce, } D_t = 1)\\
&\text{if } D_t = 1: \quad \text{s.t. } w_t^s\kappa(1-\phi) = c_t^s + x_t^s, \text{ for } s \in \{f, m\},\\
&\text{if } D_t = 0: \quad \text{s.t. budget constraint of the couple (3.10),}\\
&\qquad\qquad W_t^{M_s}(\Omega_t^M) \geq V_t^{D_s}(\omega_t^s) \text{ for } s \in \{f, m\}, && \text{(participation constraints)}
\end{aligned}
$$

where, whenever $D_t = 0$ and the participation constraint of person s binds, the new Pareto weight is: $\theta^s_{t+1} = \text{argmin}_{\theta^*}|\theta^s_t - \theta^*|$ such that participation constraints hold. When both participation constraints are satisfied the Pareto weights stay constant. Divorce happens when at leas one of the participation constraints cannot be met for any change in the bargaining power. The individuals' value of marriage, conditional on $D_t = 0$, is $W^{M_s}_t$ for $s \in \{f, m\}$, is defined as

$$W^{M_s}_t = u(\tilde{c}^i_t, \tilde{Q}_t) + \hat{\psi}_d + \beta E_t V^{M_s}_{t+1}(\Omega_{t+1}),$$

where $\tilde{\mathbf{q}}^M_t = \{\tilde{x}_t, \tilde{c}^m_t, \tilde{c}^f_t, \tilde{P}_t\}$ is the arg max of problem (3.14) conditional on having chosen $D_t = 0$. $V^{M_s}_{t+1}(\Omega_{t+1})$ instead 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 (3.11) accounting for the cost of divorce. Note that I follow the literature assuming that the planner evaluates the welfare of the two members of the couple if a divorce happens with the current Pareto weights.

This formulation of the problem for spouses implies that Pareto weights stay constant as soon as both spouses are better-off in marriage with the current Pareto weights. Instead, when one participation constraint is binding, her Pareto weight increases by μ^s_t, the smallest value such that a person is willing to stay married. In this case, the allocation corresponds to the solution of the above Pareto with the modified weights. Instead, when for any value of μ^s_t participation constraints are not met, divorce happens. Note that each time the martial contract is rebargained, the solution departs from the Pareto optimal allocation, making consumption more volatile and hence reducing the strength of risk-sharing. The fact that spouses cannot commit not to divorce also influences specialization within the couple. In fact, the efficient solution might entail full specialization in home production for women, but the actual allocation is such that both spouses work on the market. This can happen because the wife fears a future divorce and would have lower future productivity if she stays at home now. The husband cannot compensate her because utility is not perfectly transferable.

1.3.8 Cohabitation

The problem of cohabiting couples is similar to those who are married, but it departs from it because the choice set $\mathbf{q}^C_t = \{x_t, c^f_t, c^m_t, P_t, D_t, M_t\}$ includes i) the possibility of marriage and

ii) the outside option is not divorce but breakup. The utility of breaking up is the same as becoming single again.[19]

The fact that breakup is costless has important implications for the utility gains of cohabitation.[20] In fact, outside options bind more frequently for cohabiting couples and breakup is more likely. This fact makes both risk-sharing and specialization in home production less functional. Even if the allocation within cohabitation departs from the efficient allocation more often than it does for marriage, the low cost of breakup makes it the preferred relationship for those that would be at high risk of couple disruption, as a couple with a low match quality. Even more interestingly, learning about the match quality can make cohabiting before eventually marrying the best strategy, as the uncertainty around match quality decreases with the duration of the relationship. This means that couples learn about the gains of being together and the surplus of marriage with respect to cohabitation, which depends crucially on match quality.

1.3.9 Partnership Choice and the Mating Market

In each period t singles have some probability of meeting a potential partner. I assume that the productivity and the education of the prospective partner depend on the agents' characteristics.[21] Precisely, the probability $p_s^{i,i}$ that agents of gender s meets someone with their same education i is

$$p_s^{i,i} = \alpha_{edu} + (1 - \alpha_{edu})\text{Share}_{s*}^{i}, \tag{1.9}$$

where Share_{s*}^{i} is the share of agents of the opposite gender having education i. Instead, the productivity of the potential partner z_t^p is

$$z_t^p = f(\overline{z}_t^{s^*,i^*}, z_t^r, \epsilon), \tag{1.10}$$

where $\overline{z}_t^{s^*,i^*}$ is a number representing the average productivity of singles of gender s^* and education i^* in t, z_t^r is the productivity of the agent net of the gender and education spe-

[19] If I included the possibility for married couples to cohabit, this choice would never be taken. This happens because the divorce cost creates a wedge between the match quality threshold of divorce and that of singleness, with the latter being always larger.

[20] What matters for the choice between marriage and cohabitation is the difference between the cost of divorce and breakup. I tried to estimate the model with a positive breakup cost, but it was not well identified and tiny.

[21] Alternatively, I could have obtained the same level of assortative mating by education assuming random meeting probabilities and that the couple receives an additional utility flow based on her and her partner's education.

cific trend,[22]. In contrast, ϵ is a normally distributed shock. These assumptions capture in a reduced form fashion the fact that people are mating assortatively.[23] Once the meeting happens, agents have to decide whether to stay in a couple, which partnership contract to choose, and pick a Pareto weight. Following Mazzocco (2007), the initial Pareto weight is set through symmetric Nash bargaining, where the threat points are the utilities of staying single. The choice between marriage and cohabitation is resolved by picking the relationship that delivers the largest Nash product: this determines the value of C_t and M_t, which I introduced while presenting the value function of singles (3.11).

1.4 Estimation

I estimate the structural model using a two-step procedure. The first step is to set a subsample of the parameters either following the literature or by matching data features without simulating the model. This is the case of labor income processes of men and women by education. Estimating labor income processes outside the model is common in the literature because it reduces the burden on the structural estimation.[24] The second step is estimating the remaining structural parameters with the method of simulated moments. This section describes every step of the estimation, discusses the structural parameters' identification, and presents the results.

1.4.1 Income Processes

I estimate the income processes using the 1997-2017 waves of the core PSID sample, including men and women aged between 20 and 65.[25] For both men and women, I compute the hourly wage rate by dividing the annual labor income by the number of yearly working hours supplied. In this way, I avoid considering a variation in working hours as a productivity shock. This correction is particularly relevant for estimating women's income process, as the variance of hours worked per year is considerably larger than that of men. I drop individuals whose hourly wage is less than half of the minimum wage in some periods, as in Low et al. (2018).

For men m, I estimate the parameters of the following model separately for college gradu-

[22]This value is taken from the PSID.

[23]In the life cycle models featured in Ciscato (2019), Shephard (2019) and Reynoso (2020) assortative mating arises in marriage markets through the interactions of preferences, incentives, supply and demand forces.

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

[25]I retain men who are household heads and women who are either household heads or married.

ates e and not college graduates n:

$$\ln(w^{m,i}_{j,t,st,sur}) = \iota^{m,i}_0 + \iota^{m,i}_1 * t + \iota^{m,i}_2 * t^2 + \iota^{m,i}_3 * t^3 + \delta_s + \nu_{sur} + u^{m,i}_{j,t,st,sur}, \tag{1.11}$$

where j stands for individual, t is age, i for education, st for state and survey year is sur. Moreover, $u^{m,i}_{j,t,st,sur} = z^{m,i}_t + e^{m,i}_{j,t,st,sur}$, where $z^{m,i}_t$ follows equation (3.6), while $e^{m,i}_{j,t,st,sur}$ is the measurement error. δ_s are state fixed effects and ν_{sur} are year of the survey fixed effects. I use the estimated age-related coefficients and the intercept to build the trends in productivity by age and education in the model. I then use the regression residuals $\hat{u}^{m,i}_t$ to estimate through GMM the variance of the permanent components of income $\sigma^{2_{m,i}}_{\zeta}$ and one of the measurement errors $\sigma^{2_{m,i}}_e$ using the following conditions:

$$\begin{aligned} E((\Delta\hat{u}^{m,i}_t)^2) &= 2\sigma^{2_{m,i}}_{\zeta} + 2\sigma^{2_{m,i}}_e \\ E(\Delta\hat{u}^{m,i}_t \Delta\hat{u}^{m,i}_{t-2}) &= -\sigma^{2_{m,i}}_e \end{aligned} \tag{1.12}$$

Note that $\Delta u_t = u_t - u_{t-2}$. I am obliged to use two years differences because of a data constraint: the PSID collected information every two years after 1997. Results are reported in table 3.6.

The estimation of women's labor earning differs from the men's one since I need to take into account the endogeneity of female labor force participation. I do so by using a Heckman selection correction procedure. Female labor earnings of individual j with education i, age t,[26] from state st in survey year sur follows the equation:

$$\ln(w^{f,i}_{j,t,st,sur}) = \iota^{f,i}_0 + \iota^{f,i}_1 * t + \iota^{f,i}_2 * t^2 + \mathbf{X}_{j,t,st,sur}\beta + \delta_s + \nu_{sur} + u^{f,i}_{j,t,st,sur}, \tag{1.13}$$

where $u^{f,i}_{j,t,st,sur} = z^{f,i}_t + e^{f,i}_{j,t,st,sur}$. $z^{f,i}_t$ follows equation (3.6), while $e^{f,i}_{j,t,st,sur}$ is the measurement error. $\mathbf{X}_{j,t,st,sur}$ are demographic controls, which include the marital status and dummies for the number of children under 18 in the family unit. I observe women's wages only if they participate in the labor market, which happens if

$$\mathbf{Z}_{j,t,st,sur}\gamma + \pi_{j,t,st,sur} > 0, \tag{1.14}$$

where $\mathbf{Z}_{j,t,st,sur}$ includes all the regressors in equation (3.23), plus a dummy indicating

[26] Age is made consistent with the model.

whether the household contracted a mortgage. Following Blundell et al. (2016), this last variable is used as an exclusion restriction for women's labor force participation: Del Boca and Lusardi (2003) find a significant impact of mortgages on women's participation in the labor market. The first stages of the two-step Heckman selection model, reported in tables 1.D.1 and 1.D.2 for women with and without college respectively, shows the strength of the instrument. In the second step, I estimate equation 3.22 to obtain the age profiles of women's wages, controlling for the inverse of the Mills ratio of the prediction obtained from the first step. Finally, I use the regression residuals from the second step $\hat{u}_t^{f,i}$ to estimate the variance of the permanent component of income $\sigma_\zeta^{2f,i}$ and that of the measurement error σ_e^{2f} through GMM, using the following conditions:[27]

$$\begin{aligned}
&E(\Delta\hat{u}_t^{f,i}|P_t^{f,i}=1,P_{t-2}^{f,i}=1)=\sigma_\pi^{f,i}\frac{\phi(\tau_t)}{1-\Phi(\tau_t)},\\
&E((\Delta\hat{u}_t^{f,i})^2|P_t^{f,i}=1,P_{t-2}^{f,i}=1)=2\sigma_\zeta^{2f}+\sigma_\pi^{2f,i}\tau_t\frac{\phi(\tau_t)}{1-\Phi(\tau_t)}+2\sigma_e^{2f,i},\\
&E(\Delta\hat{u}_t^{f,i}\Delta\hat{u}_{t-2}^{f,i}|P_t^{f,i}=1,P_{t-2}^{f,i}=1,P_{t-4}^{f,i}=1))=-\sigma_e^{2f,i},
\end{aligned} \tag{1.15}$$

where $\phi()$ and $\Phi()$ stand for the density and the distribution function of a standardized normal respectively, while $\tau_t=-\mathbf{Z}_{j,t,st,sur}\gamma$. The results are displayed in table 3.6. It is interesting to notice that the variance of the permanent income shock is larger for college graduates, both for men and women. This result is consistent with previous research by Blundell et al. (2008), Meghir and Pistaferri (2004) and Hong et al. (2019), who also find a larger variance of the permanent component of income for college graduate men. Instead, the shocks for men are slightly larger than those for women. Note also that the gender earnings gap is larger for college graduates.[28] The variances of income in $t=1$ are taken directly from the data. I assume that the the innovations in the idiosyncratic productivity processes are not correlated within the household.[29]

This approach to estimating income processes has the disadvantage of not considering that those without a college degree face a higher risk of unemployment than those with a college

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

[28]The difference in log income between men and women is 0.45 for college graduates and 0.31 for non college graduates.

[29]The results are robust towards assuming a correlation of the shocks of 0.15, a value used by Heathcote et al. (2010). In fact, the ratio of college over non college ever married at 35 is 1.13 in the main version of the model and 1.15 in the robustness, while the ratio of college over non-college that ever cohabited at 35 is 0.55 in the main version of the model and 0.73 in the robustness. The main reason why I assume a correlation of the shocks of 0.0 in the main version is that it allows to discretize the persistent shocks following the generalization of Rouwenhorst's method to non-stationary AR(1) processes described in Fella et al. (2019).

degree. Section 1.E in the appendix tackles this problem. I provide an alternative estimation of the income processes where I consider the risk of working fewer hours because of unemployment. The results are robust to this sensibility analysis.

Table 1.2: Parameters of the income processes

Parameter	Symbol	Value
Variance of f's permanent income shock—college	$\sigma_{\zeta}^{2f,e}$	0.0261
Variance of m's permanent income shock—college	$\sigma_{\zeta}^{2m,e}$	0.0316
Variance of f's permanent income shock—no college	$\sigma_{\zeta}^{2f,n}$	0.0149
Variance of m's permanent income shock—no college	$\sigma_{\zeta}^{2m,n}$	0.0230

1.4.2 Preset Parameters

This section describes how I choose the parameters that are set externally from the model. Each period in the model lasts one year. I 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. I assume that male and female agents start making decisions at age 20 and 18 respectively,[30] while they retire at 65 and die with probability 1 at ages 82 for men and 80 for women.[31] The relative risk aversion γ and the discount factor β are set to 1.5 and 0.98 to match those in Attanasio et al. (2008). Recall that I impose the shape of the utility functions for the private and public good to have the same shape. I did this is because I do not have the data that could identify the shape of the utility for public goods. Instead, the parameters relative to the production of public goods, ν and κ are those in McGrattan et al. (1997). As far as pensions are concerned, I follow Heathcote et al. (2010), which mimics the US system's progressivity with the difference that the amount of retirement income depends on wages in the period before retirement only. The parameter ϕ is fixed to 0.196 to reflect the average time spent on housework relative to time spent working in the market.[32]

The parameters regarding the mating market, contained in equations (3.15) and (3.16), are

[30] Couples such that men are always 2 years older than females. This is also the median age difference between partners in the NLSY97.

[31] I also assume that agents cannot divorce nor renegotiate the marital contract after retirement. I made this assumption to make the model solution faster.

[32] By setting the workweek to 40 hours as in Greenwood et al. (2016), and by using the average number of hours spent on housework in my PSID sample, which is 9.76, I get ϕ=9.76/(40+9.76). The actual average working hours for people older than 25 in the NLSY97 are 39.6 for women, 45.3 for men, while the respective median values are 40 and 42.

pinned down to obtain a realistic degree of assortative mating with respect to wages and education. I decided to set these parameters before the structural estimation because individuals are not very "picky" with respect to these two dimensions and hence even varying considerably the deep parameters does not translate to sensibly different outcomes for the mating market.[33] Concerning labor productivities, I target the correlation in wages in the

Table 1.3: Preset parameters

Estimated Parameters	Symbol	Value	Source
Initial age		18-20	
Retirement age	T_R	65	
Age at death	T	80-82	
Years per period		1	
Mating market—productivities			PSID
Mating market—education	α_{edu}	0.53	NLSY97
Pensions system			Heathcote et al. (2010)
Fixed non working hours	ϕ	0.196	PSID Hours
Relarive Risk Aversion	γ	1.5	Attanasio et al. (2008)
Discount factor	β	0.98	Attanasio et al. (2008)

Function	Symbol	Value	Source
$Q_t = [x_t^{\nu} + \chi(\text{Home production time})^{\nu}]^{\frac{1}{\nu}}$	χ	3.76	McGrattan et al. (1997)
	ν	0.19	McGrattan et al. (1997)

PSID. Moreover, I chose the process that generates partners' wages to respect symmetry. For instance, men in a relationship at age t should have, on average, the same productivity regardless of being the person simulated for her whole life cycle or being a partner of a woman who is simulated for her whole life cycle.[34] I obtain a correlation in wages of 0.38 for married couples in the simulated sample, while the data value is 0.33. Regarding the assortative mating by education, I set $\alpha_{edu} = 0.53$ and I get that the simulated correlation in the dummy variable *college graduate* for partners is 0.44, while it is 0.41 in the data.

1.4.3 Method of Simulated Moments

I use the method of simulated moments (McFadden, 1989; Pakes and Pollard, 1989) to pin down the vector $\vartheta = (\alpha, \lambda_e, \lambda_n, \sigma, \sigma_\eta, \kappa, \mu,)$ of the 7 remaining parameters of the model. I use 78 moments for the structural estimation, which captures the process of marriage and

[33]This happens because, for reasonable values of the deep parameters, refusing a partner because she has a low wage it is very costly, since it implies waiting one or more periods without sharing the public good.

[34]Agents in the simulated sample meet potential partners. I follow the latter only while they are in a relationship with the person in the sample. Figure 3.I.1 shows that the average productivity and its variance by age are similar for "simulated sample" and "partner" agents.

cohabitation creation and dissolution and female labor supply.[35] More precisely, I include as targets the hazard of divorce (12), the hazard of breakup (6), the hazard of marriage (6), the share of people ever married over time (16), the share of people that ever cohabited over time (16), female labor supply for married (3) and cohabiting women (3) and the difference by education in the share of those in a relationship,[36] by age (16). I compute all the moments using the NLSY97, namely the sample I described in the empirical section.

The first step for the estimation is to solve the model for a vector of parameters ϑ, then simulating income and match quality shocks to obtain the behavior of 15000 fictional agents, which follows the same distribution over education, age, gender, and censoring age as in the data. This allows me to construct simulated moments and compare them with their empirical counterpart: the objective is to obtain ϑ such that this difference is as small as possible. Formally, the problem that I solve is

$$\hat{\vartheta} = \arg\min_{\vartheta} \qquad (\mathbf{m} - \mathbf{m}_{\vartheta})'\mathbf{W}(\mathbf{m} - \mathbf{m}_{\vartheta}), \tag{1.16}$$

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. I minimize this objective function using the global minimization algorithm TikTak. According to Arnoud et al. (2019), TikTak outperforms an array of global and local optimizers when the target is a complex objective function.[37]

1.4.4 Identification

This section describes how the structural parameters of the model are identified heuristically. The parameter α is identified from total female labor supply. When this parameter is large, the household wants to produce more public goods, which requires women's time.

[35] I normalize female labor supply in the data such that if women work full-time, female labor supply is equal to one. The assumption that working full-time corresponds to 40 hours of work per week was also used for calibrating ϕ. Alternatively, I 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.

[36] By having been in a relationship I mean having cohabited, married or both.

[37] This algorithm is composed of (i) a global stage, where the fit of the model is evaluated at quasi-random points; (ii) a local stage, where local minimizers are run starting from a combination of the initial points and the solution of local minimizers that already converged.

Instead, μ affects the gap in female labor force supply for married and cohabiting couples. When this parameter is large, specialization within cohabitation becomes harder without having a commitment technology. This implies a large gap in labor supply between the two relationship types. The parameters λ_e and λ_n are intuitively identified by the share of people in a relationship over age and education. The parameter σ is crucial for determining the share of people 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. All these changes make cohabitation relatively more attractive than marriage. The share of labor income left after divorce κ plays an essential role in the surplus of marriage with respect to cohabitation, but it is mainly identified by the difference in the hazard of divorce and breakup. The larger it is, the stronger selection on marriage and cohabitation with respect to match quality, which creates a gap in the stability of those relationships. Finally, σ_η is identified by the shape of the hazard of divorce, as described in the empirical section.

1.4.5 Model Fit

Table 3.8 reports parameters estimates. The estimated standard deviation σ of the initial match quality shock is 0.063, while standard deviation σ_η of the noise for match quality is 0.163. The probability of meeting partners λ_e and λ_n are respectively 0.165 and 0.608, while the share of labor income left after divorce κ is 0.622. The weight on the public good α is 14.14, while the productivity loss parameter linked to nonparticipation μ is 0.297.

Table 1.4: Estimated structural parameters

Estimated Parameters		Value
Standard deviation of the true match quality	σ	0.063
Standard deviation of the noise	σ_η	0.163
Probability of meeting a partner, College	λ_e	0.165
Probability of meeting a partner, No College	λ_n	0.608
Weight of public good	α	14.14
Loss in productivity while not working	μ	0.297
Share of income left after divorce	κ	0.622

NOTES: The parameters in the table are estimated by the Method of Simulated Moments.

I report the fit of the model in table 3.9. The share of people who ever cohabited and married over time are well matched. The difference in the share of ever been in a relationship by education is overall well fitted. The moments regarding female labor supply are also well matched, capturing both the level and the difference for married and cohabiting couples. The hazard of divorce and breakup are well matched. As in the data, the hazard of breakup is and decreasing over time and larger than the hazard of divorce. The hazard of marriage is slightly lower than in the data.

The model is validated according to its ability to fit a set of external moments which are tightly linked to the mechanisms underlying my model, and hence are relevant for the formation of different mating strategies by education. I check the ability of the model to match the relative hazard of divorce by premarital cohabitation duration and by education, obtained using a Cox regression.[38] The fit for the margins of premarital cohabitation is shown in figure 1.B.5, where the direction and the size of the effect of the intensive and extensive margins is the correct one. Instead, the relative risk of divorce of college graduates in simulated data is 0.64, which lies slightly above the 95% confidence interval of the relative hazard obtained using actual data. Since both insurance and specialization within the couple depend on the selection of women into the labor force, I check the model's ability to fit the wage of working women by age in the PSID. Figure 1.B.4 displays the empirical and simulated wages over the life cycle by education and by gender. I can see that simulated wages for women resemble those in the data, meaning that the participation margin works correctly.[39]

[38]The relative hazards are obtained using two Cox regression using yearly data for consistency between the model and the data. One regression analyzes the two margins of premarital cohabitation as the only two regressors, while the second regression contains a dummy for being college and age as regressors.

[39]For example, if the selection of women into the labor force was too dependent on women's potential wages in the simulations, I would have observed that the average wage of working women was higher than in the data.

Table 1.5: Model fit, validation and results

Estimated Moments	Model	Data	95% CI
Hazards over Time	fig. 3.H.1	fig. 3.H.1	fig. 3.H.1
Share Ever Cohabited and Married	fig. 1.B.2	fig. 1.B.2	fig. 1.B.2
Ever in a relationship—difference by education	fig. 1.B.2	fig. 1.B.2	fig. 1.B.2
Female Labor supply—marriage and cohabitation	fig. 1.B.3	fig. 1.B.3	fig. 1.B.3
External Moments	**Model**	**Data**	**95% CI**
Women wages by age	fig. 1.B.4	fig. 1.B.4	fig. 1.B.4
Premarital Cohabitation and Divorce Risk	fig. 1.B.5	fig. 1.B.5	fig. 1.B.5
Risk of of Divorce of College over Non College	0.64	0.48	[0.37,0.60]
Results—Mating Strategies by Education	**Model**	**Data**	**95% CI**
Likelihood to Cohabit, college over noncollege singles	0.34	figure 1.1	figure 1.1
Likelihood to Marry, college over noncollege cohabitors	1.19	figure 1.1	figure 1.1

The main reason why I built and estimated a structural model of partnership choice was to understand to what extent economic incentives can explain the differences in mating strategies by education. Hence, now I estimate the same duration models I used in the empirical section, but with simulated data. I find that the sub-hazard of cohabiting for college graduates is 0.34 of the rest of the population. This number is slightly lower than its empirical counterpart, which I estimated to be 0.54-0.74 depending on the controls included. Finally, using simulated cohabitation spells, I find that the sub-hazard of marriage of college graduates is 1.19 of the rest of the population. This number is lower than its empirical counterpart, which was 1.62-1.91 depending on the specification, but the effect's direction is the same.[40] Overall, the ratio of college over non-college ever married at 35 is 1.18 in the data and 1.13 in the simulations. Also, the ratio of college over non-college that ever cohabited at 35 is 0.86 in the data and 0.55. in the simulations.

I can conclude that my model can account for most of the differences in mating strategies by education. The differences by education regarding singles' choices are larger than in the data. It can be that non-economic forces—such as values and religion—are pushing their choice in the opposite direction. In the next section, I disentangle which mechanism accounts for these differences and quantify their importance.

[40]See figure 1.1 for the details.

1.5 Mechanisms and Counterfactual Experiments

This section aims to understand better the quantitative importance of each mechanism present in my model. It also aims to understand which features of the income dynamics matter for triggering different mating strategies by education. To do so, I examine the outcomes obtained simulating the estimated model and through a series of counterfactual experiments.

Selection into partnership and learning. One result of my model is that, as it is more costly to divorce than breakup, the match quality threshold for marrying is larger than that for cohabiting, as described by figure 1.5, panel (a). This is because for intermediate values of the match quality, the risk of divorce is too large. In fact, my simulations show that the distribution of match quality at meeting for married couples dominates that of cohabitors, as depicted in figure 1.5, panel (b). This effect is quantitatively large: only 16.9% of marriages started with a value of the match quality that is lower or equal to its structural mean value,[41] as opposed to 35.2% of cohabiting couples. Figure 1.B.5 suggests that both the extensive and intensive premarital cohabitation margins play an important role in divorce through selection and learning. In table 1.6 I report the results of a counterfactual experiment where cohabitation is no longer an option:[42] the share of marriages surviving up to their seventh anniversary is 67% in the experiment versus 79% in the baseline.[43] This suggests that the rise in cohabitations (due to a decreasing stigma) contributed to the decrease in the US divorce rate. The decline of divorce started at the beginning of the 1980s,[44] a period during which the share of women that ever cohabited increased substantially and the rate of marriage declined.[45] Among couples that chose to cohabit as their first relationship in the baseline, 82% would have married according to the counterfactual policy functions. This suggests that cohabitation is closer to marriage than singleness. Given this result, it is not unexpected to observe that the share of couples ever married at 35 moves from 63% in the baseline to 94% in the counterfactual.

[41] The median value of the match quality for a potential couple that just met is 0, while the median of observed match quality of formed couples is larger because of selection.

[42] This can be interpreted as an extreme increase of stigma towards cohabitation, which can be modeled as a utility penalty that hits cohabiting couples every period. Lower values of utility penalty deliver intermediate results.

[43] Survival is computed applying to individuals the duration specific hazard rate.

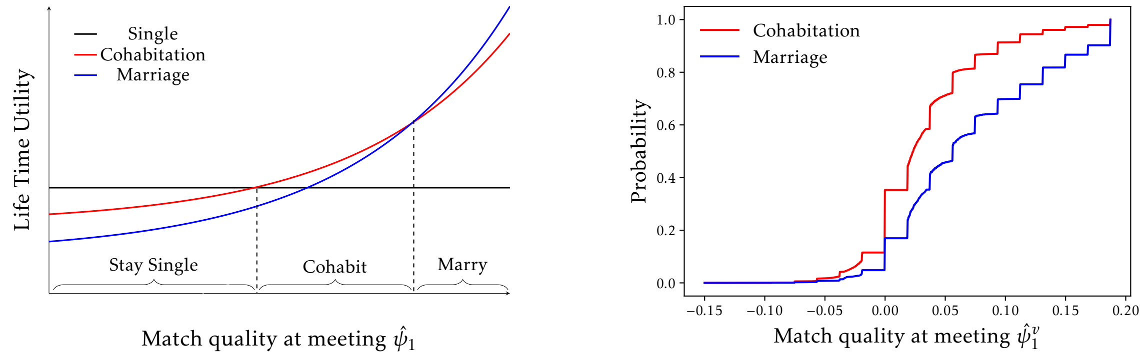
[44] Note that a rise in cohabitation that happens for reasons other than stigma might cause the divorce rate to vary differently.

[45] According to Manning (2013) the share of women that ever cohabited in the United States moved from 33% in 1987 to 60% in 2010.

Figure 1.5: Selection into partnership

(a) Value functions and partnership choices

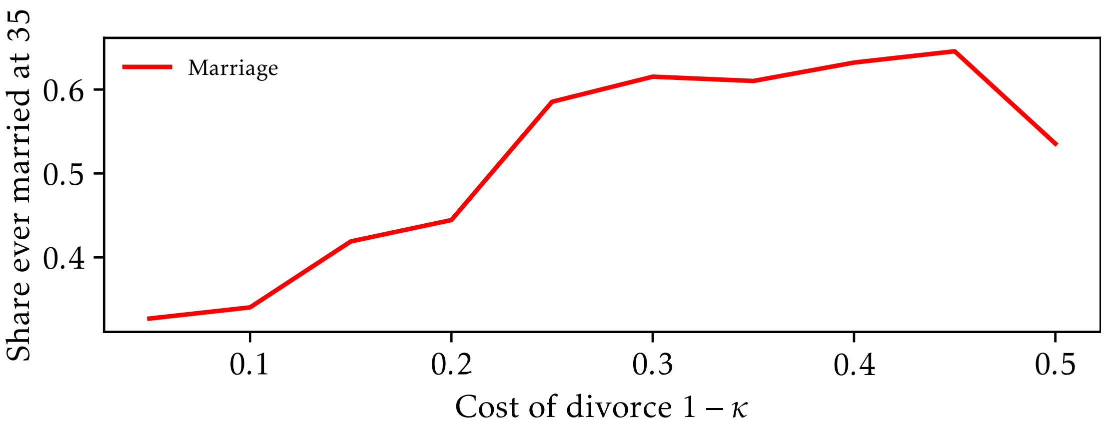
(b) Cumulative distribution of match quality



NOTES. The left pane shows the life time utility to a person of remaining single or start cohabiting or marrying a partner she just met. Note that the threshold for marriage is larger than for cohabiting. The right pane represents the cumulative distribution of the match quality at first meeting conditional on agents' choices that arise from the model's simulation.

The role of divorce. The only difference between marriage and cohabitation in the model is the cost of splitting. I already explained that divorce acts as a commitment technology that generates marriage gains with respect to cohabitation. Simultaneously, a high cost of divorce might discourage people from marrying since the utility loss in case of splitting would be larger. In figure 1.6 I show that the share of people deciding to marry is hump-shaped over the cost of divorce, pointing to the existence of this trade-off.

Figure 1.6: Share of people ever married at 35 and the cost of divorce



The role of income dynamics. College graduates and those with fewer years of education face different income processes: which features of the wage dynamics are the most important for explaining the rise in mating strategies by education? I perform two experiments to answer this question.

First, I equalize the productivity of women to that of men.[46] The results, displayed in table 1.6, show a decrease (increase) in the share of people that ever married (cohabited). The reason is that shrinking the gender wage gap reduces the scope of specialization within the household, thereby reducing marriage gains with respect to cohabitation. Interestingly, this change is mainly driven by college graduates, for whom the gender productivity gap is larger.[47]. Moreover, as male college graduates have a comparative advantage in market production relative to most women, they could form traditional marriages to enjoy the gains from specialization. Hence, a shrinking gender wage gap affects college educated men more by reducing the pool of women with whom they have a comparative advantage in market production. In line with my reasoning, table 1.6 shows that in the counterfactual, the College-NoCollege gradient in the likelihood of singles to marry shrinks significantly. Education differences in the transition from cohabitation into marriage remain constant, even if the incentives to marry for college graduates are eroded in the counterfactual. This is due to a selection effect, as the flow of cohabitors that marry decreases substantially. Note also that the uneven decrease of the gender gap over time is consistent with the rise in cohabitation and the polarization in partnership choices by education. The size of this effect is large: erasing the productivity gap would cause the share of people ever married by the age of 35 to decrease by 25% (see again table 1.6). 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).

As a second experiment, I equalize college graduates' persistent shock variance to that of those with fewer years of education. The results in table 1.6 show that the difference in the likelihood of marrying between College and Non College graduates shrinks. This is because the demand for consumption insurance mechanisms of college graduates decreases as their income process becomes less volatile. Consistent with this, the aggregate share of people that ever married (cohabited) decreases (increases) in the counterfactual, driven by college graduates' behavior.

Changes in cohabitation and marriage over time. What is the role of income dynamics for the changes in partnership choices by education over time? To answer this question, I per-

[46]Precisely, I close the gender pay gay within education classes.

[47]Cortés and Pan (2019) show that the gender pay gap declined more slowly at the top of the distribution.

form a backward experiment asking what would be the partnership patterns for the NLY97 young birth cohorts (Y.C.) if they faced the income processes of the 1951-1955 older birth cohorts (O.C.).[48] Through simulations, I find that the share of people that ever married at 35 years old in the O.C. (Y.C.) is 74.1% (63.7%), while the share of people that ever cohabited is 62.1% (72.3%). Hence, the model can partially explain the changing partnership patterns. In the data, the share of people in the O.C. (Y.C.) that respectively ever married and cohabited at 35 is 92.6% (63.7%) and 42.2% (73.9%). In the simulations, the share of marriages that lasted seven years or more is 0.79 (0.78) for the Y.C. (O.C.), which is consistent with the decrease in divorce rates observed in the USA over the last twenty years. I also analyze the evolution over time of the college-non college gap in partnership choices for marriage (Gap_{M}) and cohabitation (Gap_{C}):

$$\begin{aligned}\text{Gap}_{\text{M}} =& \text{\% Coll married ever married at 35 (Y.C.)} - \text{\% NoColl ever married at 35 (Y.C.)}\\ &- (\text{\% Coll ever married at 35 (O.C.)} - \text{\% NoColl married (O.C.)}),\\ \text{Gap}_{\text{C}} =& \text{\% Coll ever cohabited at 35 (Y.C.)} - \text{\% NoColl ever cohabited at 35 (Y.C.)}\\ &- (\text{\% Coll ever cohabited at 35 (O.C.)} - \text{\% NoColl ever cohabited at 35 (O.C.)}).\end{aligned} \tag{1.17}$$

Gap_{M} is 9.6% in the data and 10.3% in the simulations, while Gap_{C} is -16.6% in the data and -5.3% in the simulations. This confirms that changes in income processes by education and gender can partially explain the uneven changes in the family structures by education. The results are driven by the fact that both the persistence of income shocks and the gender wage gap by education were more similar in the O.C. than in the Y.C.: see section 1.F in the appendix for the details.

[48] I measure the marriage and cohabitation patterns of the 1951-1955 birth cohorts using the National Survey of the Household (NSFH) wave III. Their income process is estimated using the PSID following the procedure described in section 1.F in the appendix.

Table 1.6: Counterfactual experiments

Experiment: Cohabitation is not a Choice	Baseline	Exp.
Share Marriages that lasted 7 years or more	0.79	0.67
Share ever married at 35 y.o.	0.63	0.94
Experiment: Closing the Gender productivity Gap	Baseline	Exp.
(Coll: % singles that marry)/(NoColl: % singles that marry)	2.80	1.07
(Coll: % cohabitors that marry)/(NoColl: % cohabitors that marry)	1.25	1.25
Share ever married at 35 y.o.	0.63	0.47
Share ever cohabited at 35 y.o.	0.72	0.86
Experiment: Productivity Shock of Coll-NoColl equalized	Baseline	Exp.
(Coll: % singles that marry)/(NoColl: % singles that marry)	2.80	2.26
(Coll: % cohabitors that marry)/(NoColl: % cohabitors that marry)	1.25	1.08
Share ever married at 35 y.o.	0.63	0.56
Share ever cohabited at 35 y.o.	0.72	0.72
Experiment: Income dynamics and partnership choices over time	Data	Model
Marriage gap by education over time (Gap_{M}, eq. 1.17)	9.6%	10.3%
Cohabitation gap by education over time (Gap_{C}, eq. 1.17)	-16.6%	-5.3%

NOTES. Exp. stands for "experiment". By saying that I equalized a variable between College (Coll) and non-College (NoColl), I mean that I substituted the original value for College with that of Non-College. The share of singles that marry is better understood as a choice between marriage and cohabitation, since censoring happens only through death and very few people retire while single.

1.6 Additional evidence on the uneven retreat from marriage

According to my analysis, the gender wage gap plays an important role in shaping the incentives to cohabit and marry. When the gender wage gap shrinks, the couple's incentive to marry decreases. Do we observe a negative association between the gender wage gap and the prevalence of cohabitation across US states and European countries? While the divergence in the gender wage gap by education coincided with the uneven retreat from marriage by education in the USA (Lundberg et al., 2016), it is useful to provide additional evidence about the association between the gender wage gap and the prevalence of cohabitation.

Panel (A) of figure 1.B.7 documents that the US States where the gender wage gap decreased the most are also those where the couples' propensity to cohabit increased the most.[49] Moreover, panel (B) of figure 1.B.7 documents that in the US States, where the gender pay gap decreased relatively more for those without a college degree, the share of

[49]For more details about how these variables are constructed, see the footnote of figure 1.B.7.

couples that cohabit grew faster for those without than with a college degree. Additional evidence that is consistent with my analysis comes from Anelli et al. (2019). They 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).

The international evidence is also consistent with my analysis. Adamopoulou (2010) documents that European countries where the gender wage gap decreased the most are also those countries where the couples' propensity to cohabit increased the most. She also shows that in all the countries she considers except Denmark, married women are more likely to be housewives than women who cohabit. This is consistent with the model's prediction that the gender wage gap affects the incentive to marry and cohabit by changing the returns to household specialization within each partnership.

Can other mechanisms explain the uneven retreat from marriage? An alternative explanation is provided by Lundberg et al. (2016), according to which marriage acts as a commitment device that allows intensive joint investment in children. Since college graduates are more likely to make time and money investments in children, their demand for commitment is larger.[50] Their story is consistent with the observed gap in time and money investment in children by education. The timing at which the gap in childcare by education exploded between the mid-eighties/mid-nineties (Ramey and Ramey, 2010) coincides with the appearance of a marriage gap by education, while the gap in money investment in children was present already in the seventies (Kornrich and Furstenberg, 2013). Furthermore, the gap in childcare time is also present in Europe (Dotti Sani and Treas, 2016), but in some European countries the marriage and cohabitation gap by education is smaller or even reversed, such as in France, Spain and Slovenia (OECD, 2016). Overall, the mechanism proposed by Lundberg et al. (2016) can help to explain the uneven retreat from marriage, but its quantitative relevance is not clear yet. I suggest that, while my and their mechanisms alone cannot explain all the gaps in partnership choice by education, these two stories can be complementary to explain the facts. For example, my model explains only some of the gap by education in the transition from cohabitation to marriage. This can be because educated couples find it particularly important to learn about the quality of their match before engaging in a coordination-intensive task as raising a baby, which they plan to have after

[50] Note that my model also captures the idea that marriage act as a commitment device that allows intensive joint investment in public goods, which include children. Yet, I do not assume that the demand for public goods differs by education.

getting married.

1.7 Conclusion

I study how economic incentives affect partnership choices by education. Using NLSY97 data, I show that *i*) college graduates are less likely to cohabit than to marry among singles and *ii*) among cohabitors, college graduates are more likely to marry. The data also show that people marrying directly have a lower risk of divorce than those who cohabited for a short period before the marriage. In particular, the lowest risk of divorce is observed among people who cohabited for a long time before marrying, suggesting that premarital cohabitation might be linked to learning about the match quality. To understand the mechanisms behind these patterns, I build a structural life cycle model of partnership choice that introduces a specific role for learning. In my theory, marriage gains with respect to cohabitation come from a better risk-sharing and specialization within the household, enforced through a costly divorce. Since the commitment gains of marriage crucially depend on the match quality, learning creates a rationale for premarital cohabitation.

I use moments concerning the mating market and female labor supply to estimate the model by the method of simulated moments. The structural estimation shows that the model's mechanisms can explain almost entirely the mating patterns by education. In particular, I run a series of counterfactual experiments to disentangle the strength of the competing mechanisms. I find that college graduates are more likely to marry because: *i*) their larger gender pay gap causes men to have a comparative advantage in market production, which is best exploited in a marriage with traditional roles; *ii*) their income is more volatile, which requires a well-functioning consumption insurance. Finally, I show that changes in the income process over the last thirty years can explain $\sim 100\%$ ($\sim 30\%$) of the evolution in the marriage (cohabitation) gap by education. These results are mainly driven by the reduction of the gender wage gap: if it were closed completely, the share of people ever married (ever cohabited) by the age of 35 would drop by 25% (increase by 19%). The role of learning and selection are also quantitatively relevant: if cohabitation were not accessible, marriages would be significantly more unstable.

Beyond what is studied in this paper, it would be interesting to model children as consumption commitments.[51] This feature, which has been shown to explain how income volatility

[51] Consumption commitments are goods whose consumption is infrequently adjusted because they involve

and age at marriage are linked (Santos and Weiss, 2016), has the potential to partially account for the different age at first marriage and cohabitation of people with different levels of education.

costly transaction costs.

Appendix

1.A Duration Models

The Cox proportional hazard model (Cox, 1972) relates some covariates to the hazard rate, defined as the rate at which the event of interest is realized at time t, conditional on not having happened until that moment. The unit of observation is a spell, defined as the time spent in a particular state (i.e. marriage and singleness). The Cox regression assumes that covariates $\mathbf{X}$, together with the vector of parameters β, shift proportionally the baseline hazard $\lambda_0(t)$, which captures unobserved heterogeneity. Formally

$$\lambda(t|X_i) = \lambda_0(t)\exp(\mathbf{X}_i\beta), \tag{1.18}$$

where λ is the hazard. I use this model because it controls for right censoring, avoiding the bias that would arise from considering some observations as having a low risk just because they were observed for fewer periods. Since the Cox regression is not convenient when individuals face additional types of risk,[52] called competing events, I use the Fine-Gray regression (Fine and Gray, 1999). This model is represented by Equation 1.18, which is the same as the Cox model with the difference that $\lambda(t|X_i)$, in this case, is a so-called *sub-hazard*. This is defined as the probability that the event of interest happens conditionally on having survived until that moment, while a competing event might already have happened.[53] This model is used when the unit of analysis is a singleness or cohabitation spell, since in

[52] When analyzing multiple risks with a Cox regression, the competing event is treated as if it was right censoring, which implicitly assumes that the main and the competing event are independent. This implies that independence of irrelevant alternatives is also assumed to hold.

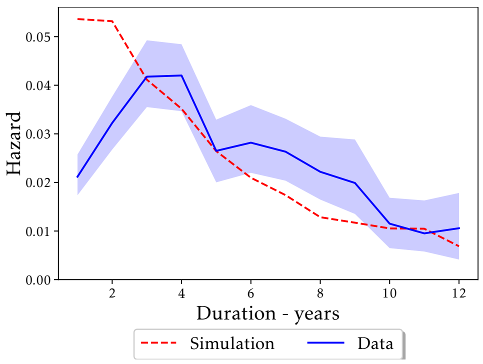
[53] This means that individuals who experienced the competing event are present in the risk set with diminishing weights, which allows to relax the assumption of independence of the competing event. While this characteristic seems unnatural, Putter et al. (2020) shows that the same parameters as the Fine-Gray regression can be obtained fitting a Cox regression corrected with a reduction factor. This is defined as the share of spells in the Fine–Gray risk set that has not yet experienced a competing event.

these cases, the competing risks are cohabitation vs. marriage and marriage vs. breakup, respectively.

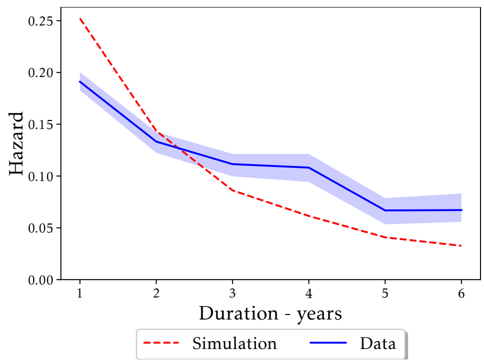
1.B Figures

Figure 1.B.1: Hazards by duration of spells: simulations and data

(a) Hazard of divorce



(b) Hazard of breakup



(c) Hazard of marriage

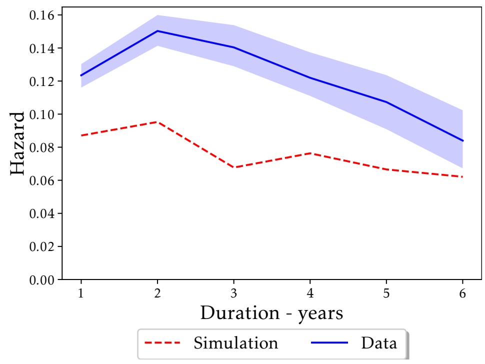
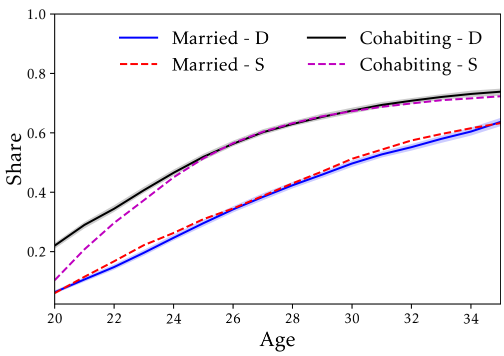
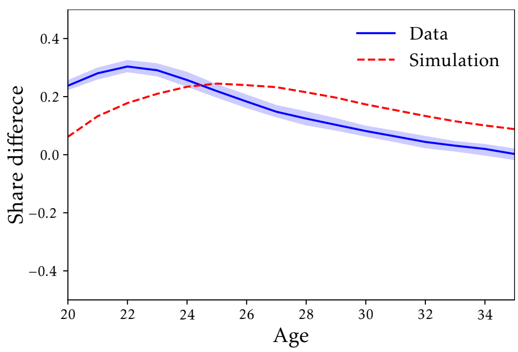


Figure 1.B.2: Ever in a relationship by age: simulations and data

(a) Ever cohabited and married

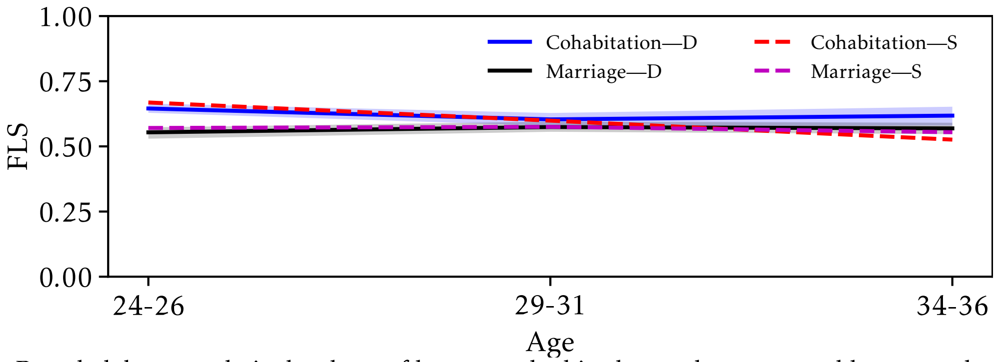


(b) ΔEver in a relationship noncollege-College



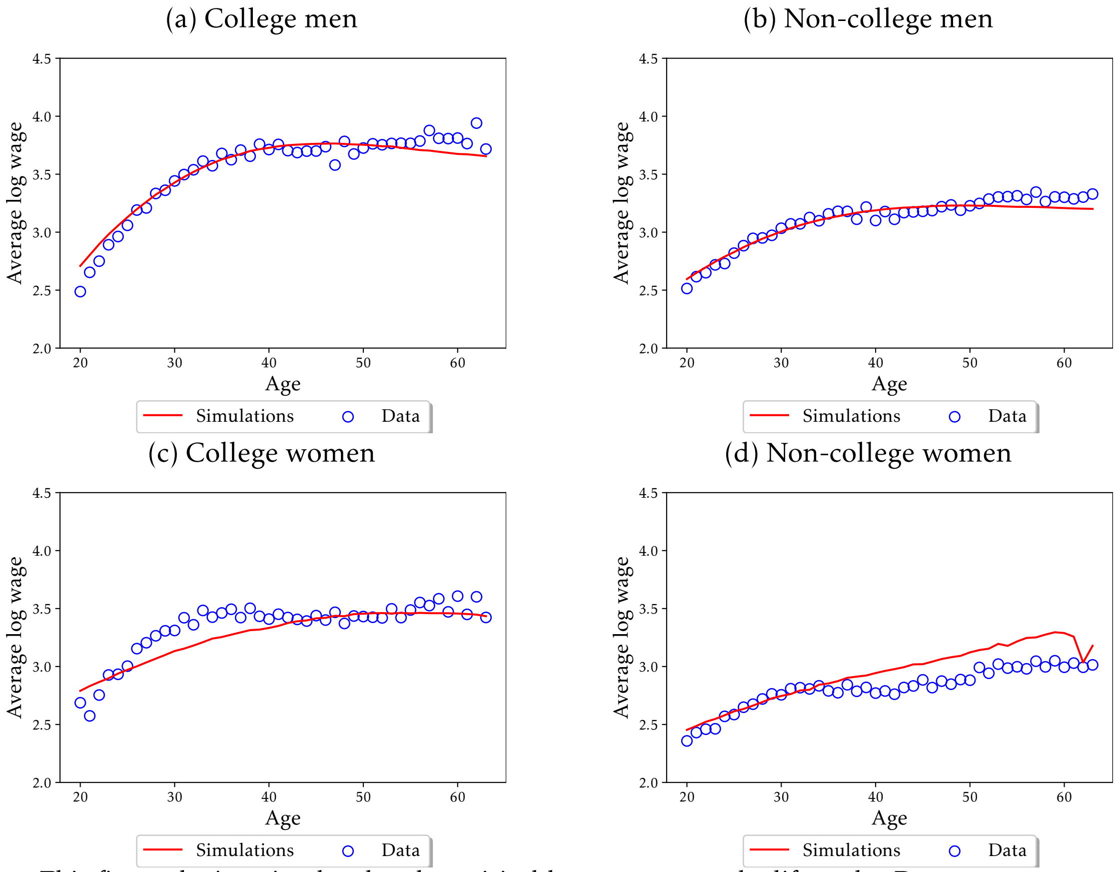
NOTES. The left panel displays the share of people ever cohabited and ever married by age. Note that cohabitation has to be interpreted as living together without being married: hence marrying someone directly does not count as cohabitation. Instead, a couple that cohabited before marrying will be part of both the ever married and ever cohabited group. The right panel shows the College-NoCollege difference in people's share ever in a relationship, by age. Being ever in a relationship means having being married or having cohabited or both.

Figure 1.B.3: Female labor supply: simulations and data



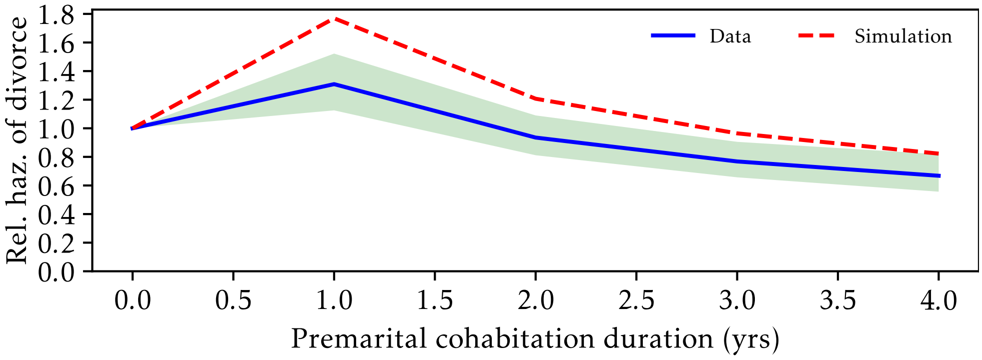
Notes. Female labor supply is the share of hours worked in the market over total hours worked, both in the market and at home. In the model is obtained by averaging $P_t^f(1-\phi)$. In the data, female labor supply is obtained by dividing average working hours in a week by (40+9.76), where 40 are the hours in a workweek and 9.76 the average time spent per week on housework by singles in the PSID.

Figure 1.B.4: Low income 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 dividing the annual labor income by the total number of hours.

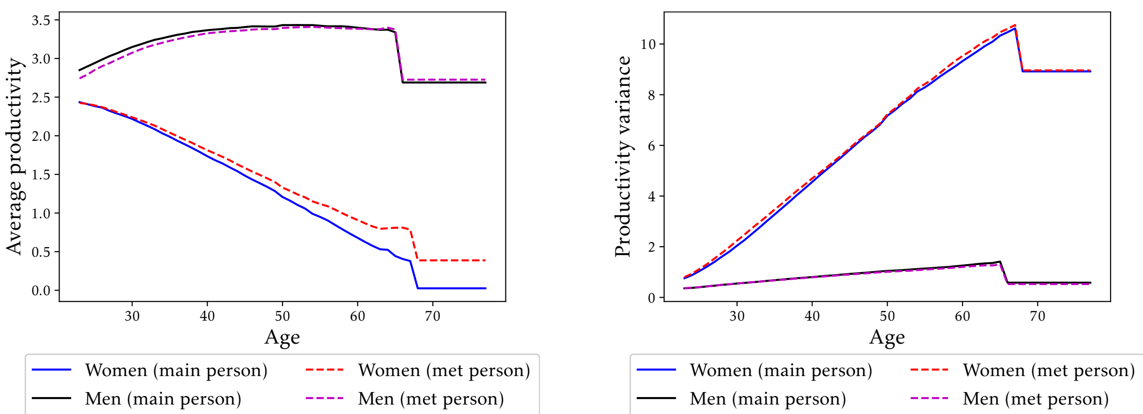
Figure 1.B.5: Relative hazard of divorce by premarital cohabitation duration.



Notes. The relative hazards are obtained using Cox regression using yearly intervals for duration. The intensive and extensive cohabitation margins are specified as in the empirical section, and a dummy for being a college graduate is also included among the regressors.

Figure 1.B.6: Labor productivities means and variances by age: simulated data

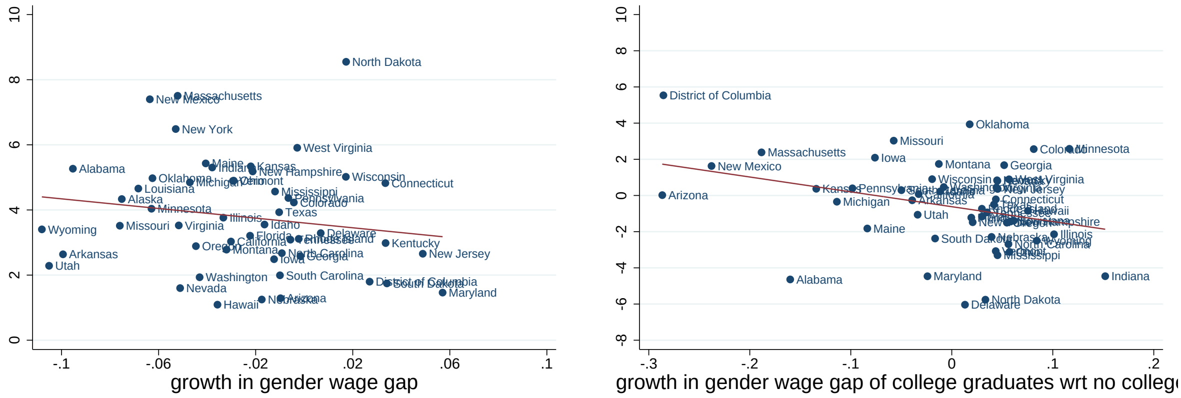
(a) Average Log income (b) Variance Log income



Notes. The figures display means and variances of productivities (potential log income) of men and women in a couple over their age. I label as “main person” the variables computed from agents that are simulated and followed through their whole life-cycle, while I label as “met person” the variables constructed using the partners met by the people whose behavior is simulated for their whole life-cycle. I construct potential wage variables using couples at any point of their relationship. Note that the values for women are extreme because they include women that never participate in the labor market. These are hit by the negative productivity drift μ each period.

Figure 1.B.7: The gender wage gap and cohabitation in the US States

(a) Growth in the % of cohabitants (b) Growth in the % of cohabitants with college wrt no college



Notes. The left panel shows the association between changes over time in the gender wage gap and changes over time in the share of couples who cohabit. The variables are built using the CPS, considering couples whose household head's age is between 21 and 65 years old. The growth in the % of cohabitants ΔCO is built as the difference in the share of couples choosing cohabitation over marriage in the periods 2019-2020 minus those in 2007-2008. The growth in the gender pay gap ΔPAY is build as the difference in the average log (hourly) wage gap in 2019-2020 minus that of 2007-2008. The right panel shows the same variables for college graduates (those who has a bachelor degree or more) with respect to those without a degree. The variable on the horizontal axis is ΔCO(college graduate) minus ΔCO(no college), while the variables on the vertical axis is ΔPAY(college graduate) minus ΔPAY(no college).

1.C Duration Analysis

This section presents the complete results of the duration models used in this paper and the robustness checks. First, I describe the controls used in the regressions that are non self-explanatory.

- *Church*: Number of worship attended last month—averaged over the years the respondent answered the question.
- *Initial Number of Children*: Number of children at the beginning of the spell.
- *Relationship Number*: Dummies of the number of marriages and cohabitations experienced by the individual before the current spell began.
- *Geographic Controls*: Dummy variables capturing the census region and whether the respondent lives in a Metropolitan statistical area.
- *Educational Homogamy*: Dummy taking the value of 1 when both partners have a college degree or whether both partners do not have a college degree.
- *Shotgun Marriage*: Dummy variable taking value one if a child was born during the first 8 months of the marriage.

- *Religion Dummies*: Dummies of religious affiliation.

Robustness check—Multinomial logit. I use the multinomial logit reshaping the dataset to be of the spell-month type as a robustness check for the results obtained with the Fine-Gray model. The multinomial logit has the following advantages: (i) does not rely on proportionality assumptions; (ii) it takes into account that duration is discrete; (iii) it allows controlling for attrition by modeling it as an additional competing risk.[54] Tables 1.C.6 and 1.C.7 show that the association between graduating from college and the risk of interest is the same as that obtained with the Fine-Gray model.[55]

Robustness check—Logit. Results in table 1.C.4 might not precise because my data is discrete or because the risk of divorce over time is not proportional to all of the covariates I control for.[56] Hence, I run a robustness check, where I use a hazard model in discrete time. Following Jenkins (1995), I reshape my dataset to be of the marriage spell-month type, then estimate a logit model where the dependent variable is a dummy that takes value one if a divorce was observed in that month and zero otherwise. The results with this method, presented in table 1.C.5, are close to the ones obtained from the Cox regression.

[54]The drawback of the multinomial logit is that it relies on the strong assumption of independence of irrelevant alternatives.

[55]Note that the magnitudes of the results of the two models are not directly comparable, since Fine-Gray estimates sub-hazard ratios, while I obtain hazard ratios with the multinomial logit. Instead, Austin and Fine (2017) point out that the direction of the sub-hazard ratio and the relative hazard ratio are the same.

[56]I performed the proportionality test by Grambsch and Therneau (1994) in all the specifications. The assumption holds for the intensive and extensive margins of cohabitation in all the specifications, except for the intensive margin in (5) and (6). For some controls, the proportionality assumption does not hold.

Table 1.C.1: Fine and Gray (1999) duration model. Observations: singleness spells.

	Sample I			Sample II		
	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Variable: Sub-Hazard of Cohabitation						
Completed College (0/1)	0.65***	0.78***	0.78***	0.54***	0.74***	0.73***
	(0.02)	(0.03)	(0.03)	(0.02)	(0.03)	(0.03)
Age		1.39***	1.39***		1.42***	1.41***
		(0.07)	(0.07)		(0.08)	(0.08)
Age Squared		0.99***	0.99***		0.99***	0.99***
		(0.00)	(0.00)		(0.00)	(0.00)
Female		1.37***	1.38***		1.34***	1.35***
		(0.04)	(0.04)		(0.04)	(0.04)
Hispanic		0.95	0.95		0.94	0.95
		(0.04)	(0.04)		(0.04)	(0.04)
Church		0.85***	0.85***		0.85***	0.85***
		(0.01)	(0.01)		(0.01)	(0.01)
Black		0.96	0.97		0.95	0.95
		(0.04)	(0.04)		(0.04)	(0.04)
Rural		0.99	0.99		0.93	0.93
		(0.05)	(0.05)		(0.05)	(0.05)
Smoke		1.44***	1.45***		1.40***	1.41***
		(0.05)	(0.05)		(0.06)	(0.06)
Initial Nr. of Children			0.93*			0.91*
			(0.04)			(0.05)
Religion Dummies		✓	✓		✓	✓
Relationship Number		✓	✓		✓	✓
Geographic Controls		✓	✓		✓	✓
Observations	12365	12133	12133	9443	9415	9415

Notes: The results are displayed in terms of relative sub-hazard ratios. A sub-hazard is defined as the probability that the event of interest happens conditionally on not having occurred already, while a competing event might already have happened. When the ratio is larger than 1, it means that the covariate increases the likelihood that the event occurs. The composition of samples I and II is described in the text. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.2: Fine and Gray (1999) duration model. Observations: singleness spells.

	Sample I			Sample II		
	(1)	(2)	(3)	(4)	(5)	(6)
DEP. VARIABLE: SUB-HAZARD OF MARRIAGE						
Completed College (0/1)	1.80***	1.15**	1.15**	1.75***	1.03	1.04
	(0.11)	(0.08)	(0.08)	(0.12)	(0.07)	(0.07)
Age		1.22	1.22		1.16	1.16
		(0.15)	(0.15)		(0.15)	(0.15)
Age Squared		1.00	1.00		1.00	1.00
		(0.00)	(0.00)		(0.00)	(0.00)
Female		1.01	1.00		1.04	1.02
		(0.06)	(0.06)		(0.07)	(0.07)
Hispanic		1.23**	1.21**		1.21**	1.19*
		(0.10)	(0.10)		(0.12)	(0.11)
Church		1.49***	1.49***		1.49***	1.49***
		(0.03)	(0.03)		(0.03)	(0.03)
Black		0.43***	0.41***		0.39***	0.37***
		(0.04)	(0.04)		(0.04)	(0.04)
Rural		0.65***	0.64***		0.64***	0.63***
		(0.06)	(0.06)		(0.07)	(0.07)
Smoke		0.54***	0.53***		0.51***	0.50***
		(0.05)	(0.05)		(0.05)	(0.05)
Initial Nr. of Children			1.52**			1.69***
			(0.27)			(0.30)
Religion Dummies		✓	✓		✓	✓
Relationship Number		✓	✓		✓	✓
Geographic Controls		✓	✓		✓	✓
Observations	12365	12133	12133	9443	9415	9415

NOTES: The results are displayed in terms of relative sub-hazard ratios. A sub-hazard is defined as the probability that the event of interest happens conditionally on not having occurred already, while a competing event might already have happened. When the ratio is larger than 1, it means that the covariate increases the likelihood that the event occurs. The composition of samples I and II is described in the text. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.3: Fine and Gray (1999) duration model. Observations: cohabitation spells.

	Sample I			Sample II		
	(1)	(2)	(3)	(4)	(5)	(6)
DEP. VARIABLE: SUB-HAZARD OF MARRIAGE						
Completed College (0/1)	1.89***	1.57***	1.52***	1.91***	1.69***	1.62***
	(0.08)	(0.10)	(0.10)	(0.08)	(0.12)	(0.11)
Educational Homogamy (0/1)		1.13*	1.14*		1.21***	1.22***
		(0.08)	(0.08)		(0.09)	(0.09)
Age		1.29***	1.29***		1.35***	1.36***
		(0.11)	(0.11)		(0.12)	(0.12)
Age Squared		1.00**	1.00**		1.00**	1.00**
		(0.00)	(0.00)		(0.00)	(0.00)
Female		0.97	0.98		0.95	0.97
		(0.04)	(0.04)		(0.04)	(0.04)
Hispanic		0.68***	0.70***		0.71***	0.73***
		(0.04)	(0.04)		(0.05)	(0.05)
Church		1.15***	1.15***		1.14***	1.14***
		(0.02)	(0.02)		(0.02)	(0.02)
Black		0.40***	0.42***		0.41***	0.43***
		(0.02)	(0.03)		(0.03)	(0.03)
Age Difference of Partners		0.97***	0.97***		0.97***	0.97***
		(0.00)	(0.00)		(0.00)	(0.00)
Smoke		0.57***	0.58***		0.58***	0.58***
		(0.03)	(0.03)		(0.03)	(0.03)
Rural		0.78***	0.78***		0.77***	0.77***
		(0.05)	(0.05)		(0.06)	(0.06)
Initial Nr. of Children			0.92***			0.92***
			(0.02)			(0.03)
Year Relationship Starts Dummies		✓	✓		✓	✓
Religion Dummies		✓	✓		✓	✓
Geographic Controls		✓	✓		✓	✓
Observations	9707	9616	9616	7910	7841	7841

NOTES: The results are displayed in terms of relative sub-hazard ratios. A sub-hazard is defined as the probability that the event of interest happens conditionally on not having occurred already, while a competing event might already have happened. When the ratio is larger than 1, it means that the covariate increases the likelihood that the event occurs. The composition of samples I and II is described in the text. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.4: Cox model. Observations: marriage spells.

	Sample I			Sample II		
	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Variable: Hazard of Divorce						
Cohabited (0/1)	16.87***	3.31***	4.35***	19.22***	3.60***	4.64***
	(5.66)	(1.36)	(1.91)	(7.19)	(1.65)	(2.29)
Log(Cohabitation Length)	0.74***	0.88***	0.85***	0.74***	0.87***	0.84***
	(0.03)	(0.04)	(0.04)	(0.03)	(0.04)	(0.05)
Completed College (0/1)		0.61***	0.64***		0.59***	0.63***
		(0.07)	(0.07)		(0.07)	(0.08)
Educational Homogamy (0/1)		1.25**	1.21*		1.24*	1.20
		(0.13)	(0.13)		(0.14)	(0.13)
Age		1.25	1.24		1.27	1.26
		(0.29)	(0.29)		(0.33)	(0.33)
Age Squared		0.99	0.99		0.99	0.99
		(0.01)	(0.01)		(0.01)	(0.01)
Female		1.10	1.09		1.10	1.08
		(0.08)	(0.08)		(0.09)	(0.09)
Hispanic		0.90	0.86		0.88	0.85
		(0.09)	(0.09)		(0.10)	(0.10)
Church		0.90***	0.90***		0.89***	0.89***
		(0.02)	(0.02)		(0.03)	(0.03)
Black		1.22*	1.12		1.22*	1.11
		(0.13)	(0.13)		(0.15)	(0.14)
Age Difference of Partners		1.00	1.00		1.00	1.00
		(0.01)	(0.01)		(0.01)	(0.01)
Rural		1.43***	1.43***		1.56***	1.56***
		(0.18)	(0.18)		(0.22)	(0.22)
Smoke		1.74***	1.68***		1.75***	1.68***
		(0.18)	(0.17)		(0.20)	(0.19)
Initial Nr. of Children			1.17***			1.18***
			(0.06)			(0.07)
Nr. of Children—Cohabitation			1.23***			1.24**
			(0.10)			(0.11)
Shotgun Marriage			1.12			1.18
			(0.12)			(0.14)
Religion Dummies		✓	✓		✓	✓
Year relationship starts Dummies		✓	✓		✓	✓
Geographic Controls		✓	✓		✓	✓
Observations	5127	4948	4948	4260	4118	4118

Notes: The results are displayed in terms of relative risk. For example, if the number next to the variable *College* is α, it means being a college graduate have a risk of marriage which is α% of the risk of the rest of the population. Standard errors are clustered at the individual level. The composition of samples I and II is described in the text. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.5: Logit model. Observations: each month of any marriage spell.

	Sample I			Sample II		
	(1)	(2)	(3)	(4)	(5)	(6)
DEP. VARIABLE: DIVORCE DUMMY						
Cohabited (0/1)	17.17***	3.24***	4.64***	19.59***	3.52***	5.00***
	(5.80)	(1.34)	(2.07)	(7.38)	(1.62)	(2.50)
Log(Cohabitation Length)	0.74***	0.88***	0.84***	0.73***	0.87***	0.84***
	(0.03)	(0.04)	(0.04)	(0.03)	(0.04)	(0.05)
Completed College (0/1)		0.62***	0.64***		0.60***	0.62***
		(0.07)	(0.07)		(0.07)	(0.08)
Educational Homogamy (0/1)		1.24**	1.21*		1.22*	1.19
		(0.13)	(0.13)		(0.14)	(0.14)
Age		1.41	1.34		1.49	1.42
		(0.31)	(0.30)		(0.37)	(0.36)
Age Squared		0.99*	0.99*		0.99**	0.99*
		(0.00)	(0.00)		(0.01)	(0.01)
Female		1.10	1.08		1.11	1.08
		(0.08)	(0.08)		(0.09)	(0.09)
Hispanic		0.89	0.85		0.88	0.84
		(0.09)	(0.09)		(0.10)	(0.10)
Church		0.89***	0.90***		0.88***	0.89***
		(0.02)	(0.02)		(0.03)	(0.03)
Black		1.20	1.08		1.20	1.08
		(0.13)	(0.12)		(0.15)	(0.14)
Age Difference of Partners		1.00	0.99		1.00	1.00
		(0.01)	(0.01)		(0.01)	(0.01)
Rural		1.43***	1.46***		1.55***	1.60***
		(0.18)	(0.19)		(0.22)	(0.23)
Smoke		1.74***	1.67***		1.74***	1.66***
		(0.18)	(0.17)		(0.20)	(0.19)
Initial Nr. of Children			1.16***			1.17***
			(0.06)			(0.07)
Nr. of Children—Cohabitation			1.21**			1.21**
			(0.10)			(0.11)
Shotgun Marriage			1.53***			1.58***
			(0.18)			(0.20)
Nr. of Children—Marriage			0.69***			0.69***
			(0.04)			(0.04)
Religion Dummies		✓	✓		✓	✓
Marriage Duration—poly.		✓	✓		✓	✓
Year Relationship Starts—poly.		✓	✓		✓	✓
Geographic Controls		✓	✓		✓	✓
Observations	426745	416018	416018	370314	361040	361040

NOTES: The results are displayed in terms of relative risk. For example, if the number next to the variable *College* is α, it means being a college graduate have a risk of marriage which is α% of the risk of the rest of the population. Standard errors are clustered at the individual level. The composition of samples I and II is described in the text. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.6: Multinomial Logit. Observation: person-month of singleness

	(1)	(2)	(3)
	Risk of Marriage relative to Singleness		
College	0.57***	0.16**	0.16**
	(0.07)	(0.07)	(0.07)
Hazard Ratio—College	1.77	1.18	1.18
	Risk of Cohabitation relative to Singleness		
College	−0.12***	−0.22***	−0.21***
	(0.04)	(0.04)	(0.04)
Hazard Ratio—College	0.88	0.81	0.81
Polynomial—Duration	✓	✓	✓
Individual Controls		✓	✓
Relationship Specific Controls		✓	✓
Children Controls			✓
Geographic Controls		✓	✓

Notes: the estimation of this model is performed using the *R* package mlogit developed by Croissant (2012). The dependent variable includes 4 possible outcomes: marriage, cohabitation, censored observation or observation lost for attrition. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.C.7: Multinomial Logit. Observation: person-month of cohabitation

	(1)	(2)	(3)
	Risk of Marriage relative to Cohabitation		
College	0.95***	0.82***	0.72***
	(0.05)	(0.07)	(0.07)
Hazard Ratio—College	2.59	2.26	2.06
Polynomial—Duration	✓	✓	✓
Individual Controls		✓	✓
Relationship Specific Controls		✓	✓
Children Controls			✓
Geographic Controls		✓	✓

Notes: the estimation of this model is performed using the *R* package mlogit developed by Croissant (2012). Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

1.D Female Labor Force Participation

Table 1.D.1: Probit Regression. Observation: females college graduates in year t.

	(1)
Dep. Variable: Female Labor Force Participation	
Mortgage	0.14^{***}
	(0.05)
Survey Year Fixed Effects	✓
State Fixed Effects	✓
Demographic Controls	✓
Labor Market Experience Controls	✓
Observations	11467

Notes. The demographic controls include marital status and number of children in the family unit, as well as age and age squared. Standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

Table 1.D.2: Probit Regression. Observation: females without college in year t.

	(1)
Dep. Variable: Female Labor Force Participation	
Mortgage	0.25^{***}
Survey Year Fixed Effects	✓
State Fixed Effects	✓
Demographic Controls	✓
Labor Market Experience Controls	✓
Observations	21419

Notes. The demographic controls include marital status and number of children in the family unit, as well as age and age squared. Standard errors are clustered at the state level. Coefficients that are significantly different from zero are denoted by *10%, **5% and ***1%.

1.E Alternative estimation of the income processes

In this section, I provide an alternative strategy for estimating men and women's income processes by education. The only difference with respect to my primary estimation strategy in subsection 1.4.1 is that I consider variations in the annual income, not in the hourly income. Hence, this approach captures the decrease in income arising from lower hours worked because of unemployment.[57] Note that this approach is less appropriate for women, since they might move to a part-time working schedule to care for children. Using the same sample as in 1.4.1, I obtain that the standard deviations of the persistent income shocks are: 0.1778 (non-college men), 0.2678 (college men), 0.2678 (non-college women), 0.3268 (college women). The ratio of the standard deviations of the persistent income shock between college and noncollege graduates decreases slightly with respect to my baseline estimation: it moves from 1.133 to 1.173 for men and from 1.323 to 1.220 for women.

The results are robust towards the slight shrinking uncertainty gap by education implied by this section's estimation. To show this, I provide a robustness check where I decrease the standard deviation of the persistent income shock of college men and women according to the estimation presented in this section.[58] I find that the ratio of college over noncollege ever married at 35 is 1.13 in the main version of the model and 1.14 in the robustness. Also, the ratio of college over non-college that ever cohabited at 35 is 0.55 in the main version of the model and 0.59 in the robustness.

1.F Income process for older cohorts

The main sample of my analysis is made of 1980-1984 birth cohorts taken from the NLSY97, whose income processes are estimated using the 1997-2017 waves of the PSID. Similarly, I estimate the 1951-1955 birth cohorts' income from the NSFH wave III using the 1968-1988 waves of the PSID. The procedure for estimating the income processes of men and women by education is similar to that described in subsection 1.4.1, with the following differences: *i*) the moments used for the GMM takes into account that the income data is collected every year and not every two years; *ii*) I use the interaction between the introduction of unilateral divorce and property regimes upon divorce as an exclusion restriction for women's participation in the labor market. This procedure is identical to that described in Blasutto (2020), with the only difference that I split the sample of men and women in two according

[57] Note that I drop observations where the number of hours is zero because labor earnings are also zero. However, among those who were unemployed at least for one week in the reference year, only a tiny fraction worked zero hours. This fraction is 0.126 for noncollege graduate men, 0.113 for college graduate men, 0.18 for noncollege graduate women, and 0.133 for college graduate women. The gap by education is low for men, but large for women. The different participation rates by education drive this. Among never unemployed women, the fraction with zero hours worked is 0.284 for noncollege graduates and 0.208 for college graduates.

[58] In practice I divide the standard deviations of the persistent income shocks of college graduates by 1.173/1.133 (men) and 1.323/1.220 (women). I do not directly use the alternative standard deviation estimates since the model's deep parameters are estimated conditionally on much less uncertain income processes for all the education and gender categories.

to whether individuals completed college or not. The results concerning the income shocks are presented in table 1.F.1 below. Income shocks are larger than those of the younger cohorts for women, and that there is almost no difference concerning the uncertainty faced by college and non college-educated men. Moreover, the gender wage gap was: i) larger than for the younger cohorts, ii) similar for college graduates and noncollege educated. The difference in log wages between men and women was 0.52 for college graduates and 0.45 for the others.

Table 1.F.1: Parameters of the income processes

Parameter	Symbol	Value
Variance of f's permanent income shock—college	$\sigma_\zeta^{2f,e}$	0.0348
Variance of m's permanent income shock—college	$\sigma_\zeta^{2m,e}$	0.0349
Variance of f's permanent income shock—non-college	$\sigma_\zeta^{2f,n}$	0.0262
Variance of m's permanent income shock—non-college	$\sigma_\zeta^{2m,n}$	0.0343

Chapter 2

Catholic Censorship and the Demise of Knowledge Production in Early Modern Italy[1]

[1]This chapter is coauthored with David de la Croix (UCLouvain, IRES/LIDAM and CEPR). This project has received funding from the European Research Council (ERC) under the European Union's Horizon 2020 research and innovation programme, under grant agreement No 883033 "Did elite human capital trigger the rise of the West? Insights from a new database of European scholars." Blasutto acknowledges financial support from the French speaking community of Belgium (*mandat d' aspirant* FC 23613).

2.1 Introduction

Italy's primacy in knowledge creation was undisputed in the fifteenth and sixteenth century. Yet, it was overtaken by North and Western Europe in the following two centuries, during which scholars and the knowledge they produced are believed to have played an essential role in the rise of the West (Mokyr, 2016). The first candidate to explain such a reversal of fortune is the fight led by the Catholic Church against novel ideas (Landes, 1999), such as heliocentrism (Gingerich, 1973), infinitesimal calculus (Alexander, 2014), and atomism (Beretta, 2007).[2] These novel ideas were at the root of the scientific revolution in Europe. What is the role of the Catholic Church in the demise of knowledge production in early modern Italy?

This paper tackles this issue by focusing on the role of one weapon in the Church's arsenal, namely the power to censor books published by scholars. The list of such prohibited books is called *Index Librorum Prohibitorum*. The question we address is whether such censorship was key in altering the growth path of the generation of new knowledge in the Italian peninsula.

We answer the question with three contributions. First, we build a database of scholars active in Italy from 1400 to 1750, and highlight two new facts: *i*) once censorship was introduced, censored authors were of better quality than the non-censored authors, but this gap shrunk over time; *ii*) the intensity of censorship decreased over time. Second, we use these facts to identify the deep parameters of a novel model linking censorship to knowledge diffusion and occupational choice. Third, we perform a counterfactual experiment to assess quantitatively the role of censorship on the decline in total publications per scholar in Italy. Our strategy to measure the impact of censorship results in a new method that explicitly accounts for agents' endogenous selection into compliant vs. non-compliant ideas. Thus, this paper consists of three parts, one for each contribution, which we now describe in more detail.

In the first part of the paper, we build a database of Italian scholars active in the Renaissance academies and universities from 1400 to 1750. For each scholar, we identify whether his (or her) work was subject to censorship by the Church. We also measure the "quality"

[2]Probably Newton would have had issues developing his particle theory of light in a country averse to atomism.

of each scholar by his (or her) quantity of written output in today's library catalogs. Using this new database, we document the drop in publications per person over the period 1400-1750. Studying the distribution of the publications per person, we highlight that, in the sixteenth century, the censored authors are of much better quality, on average, than the non-censored authors. Moreover, this difference shrunk over time. The intensity of censorship decreased as well, after it was first introduced in the sixteenth century. This pattern may reflect either a deliberate choice of the best authors to switch from non-compliant to compliant publications, a change in the Church's policy, or both.

In the second part of the paper, we design a structural model linking censorship to knowledge diffusion. The model explicitly includes the two mechanisms described in the first part. In the model, knowledge is codified in books and can be of two types: conformist and non-conformist. Following the literature on knowledge diffusion (Lucas, 2009; Lucas and Moll, 2014; De la Croix et al., 2018), we assume that authors randomly draw ideas from the stock of knowledge left by the previous generation, retaining the best one. We introduce a novel occupational choice made by printers between printing compliant/conformist books or revolutionary/non-conformist books. Revolutionary books are less likely to be printed if they are of lower quality or rarer than compliant books.[3] We show that, by censoring revolutionary books, the Church can not only reduce the share of people in the revolutionary occupation, but, more importantly, alter the development path of knowledge drastically. An initial temporary blow to the revolutionary ideas can force society to convergence towards a compliant steady state. Since setting up a censorship apparatus is costly, if the Church displays patience she has the incentive to delay censorship. This rationalizes why the Church waited several decades after the rise of Protestantism before setting up an Index of forbidden books.[4]

In the third and last part of the paper, we use the facts highlighted in the first part to identify the deep parameters of the structural model. The most important parameter, namely the rate of censorship, is intuitively identified by the share of censored authors. The model implies a one-to-one relationship between the share of censored authors and the two sectors' relative quality, conditional on parameter values. Hence, the model's ability to also

[3]In a robustness exercise, we also consider the possibility that authors and printers self-censor because of the fear of being persecuted by the Inquisition.

[4]Donato (2019) notice that "As is well known, in Italy the Inquisition was reorganized relatively late (1542), in response to the 'infection' of Protestantism." The first roman Index was published even later (1559), and censorship was institutionalized in 1564 by the Council of Trent.

match the (not targeted) dynamics of the quality of censored and non-censored authors support the model's mechanisms.[5] The fixed cost necessary to impose censorship is picked to match the timing of the creation of the first Index of forbidden books. Simulations show that imposing a censorship rate of 18% on the non-conformist books was sufficient to decrease the share of non-conformist authors from 51% in 1470-1550 to 25% in 1680-1750. We conclude that censorship reduced by 28% the average log publication per scholar in Italy. Interestingly, half of this drop stems from the induced reallocation of talents towards compliant activities, while the other half arises from the direct effect of censorship on book availability.[6]

The development and estimation of the structural model constitute a new methodology to measure the effect of censorship on knowledge growth. Not only we account for the effect of censorship on the availability of already written books, but also for its repercussions on the sector and quality of future knowledge. This is done by modeling the endogenous selection of agents into the compliant vs. non-compliant sectors, which depends on past knowledge and censorship. The introduction of censorship is also endogenized. Overall, the model's structure and its estimation allow us to build a counterfactual path of knowledge dynamics characterized by the absence of censorship.

Literature This paper relates to three strands of the literature. First, we add to the existing literature that studies the effects of censorship. Motivated by the fact that a large share of the world population is currently subject to censorship,[7] previous research studied how autocratic governments strategically impose censorship (King et al., 2013; Zhuang, 2019) and its effectiveness in the spread of non-compliant ideas (Roberts, 2014). This paper contributes to this literature by proposing a novel method to study censorship, accounting for the endogenous selection of agents into compliant vs. non-compliant knowledge. Another strand of the literature explores the fight of government/religious institutions against novel ideas in early modern Spain (Vidal-Robert, 2011), Europe (Anderson, 2015), Imperial China (Koyama and Xue, 2015), and the Islamic world (Iyigun, 2015; Chaney, 2016;

[5]Note that we also target the dynamics of the overall quality of authors. The relative productivity in the two sectors is identified by the level and dynamics of the share of censored authors.

[6]The effect of censorship is also due to the interaction between i) its direct effect and ii) the induced reallocation of talents. We reported the size of i) and ii) assuming that the effect of the interaction is shared between i) and ii) proportionally, according to their relative "pure" effects.

[7]According to the report "Freedom of the Press 2017" by Freedom House, only 13% of the world population enjoys a free press: https://freedomhouse.org/report/freedom-press/2017/press-freedoms-dark-horizon.

Rubin, 2017). Relative to these works, this paper differs by distinguishing the effect of censorship from that of the inquisition. Censorship affects knowledge production by making some ideas unavailable to future generations, while the inquisition is the armed arm of the Church, responsible for punishing heretics. Censorship can be effective even if heretic authors do not risk their life. This paper is also the first work in economics about the effect of Catholic censorship, except for Becker et al. (2021). Their work studies the effect of censorship on the number of printed books. We focus on scholars, their choice to comply with the Church's ideology, and the dynamic effects of censorship via diffusion of knowledge to future generations.

Second, our paper contributes to the literature on changes and persistence in institutions and the ruling class (Acemoglu and Robinson, 2001; Acemoglu, 2008; Acemoglu and Robinson, 2008). More closely related to our work, Bénabou et al. (2015) focus on the persistence of religiosity in a framework where belief-eroding innovations can be censored, and religious institutions can adapt the doctrine to the new knowledge. Ekelund, Hebert, and Tollison 2002; 2004 study the behavior of the Catholic Church before and after the rise of Protestantism by interpreting her action as an incumbent monopolistic firm. Compared to this literature, we propose a dynamic approach to understanding the persistence of the Catholic Church's power, where her decisions to impose censorship depend on the current and future (endogenous) distribution of authors' quality by sector. Our framework allows us to rationalize both the Church's late reaction to the rise of Protestantism and that several books censored in the sixteenth century could circulate freely in the previous centuries.

Finally, this paper is tied to the literature that studies the causes at the root of the decline of Italy. The hypothesis regarding the demise of Italy include the excessive control by the guilds (Cipolla, 1994), the inability of Italy to seize the new profitable trade routes leading across the Atlantic (Landes, 1999; Braudel, 2018), and the fight of the Catholic Church against novel ideas (Landes, 1999; Gusdorf, 1971). We focus on the latter argument by examining the role of the Catholic Church's censorship on knowledge diffusion. Compared to the literature on knowledge diffusion in the Malthusian epoch (De la Croix et al., 2018), in which knowledge is embodied into craftsmen, we model a complementary vector of idea transmissions by focusing on codified/written knowledge. We do not seek to make a direct link between censorship and economic growth, even though recent research highlights the importance of upper-tail human capital for pre-industrial Europe's take-off (Squicciarini and Voigtländer, 2015; Cantoni and Yuchtman, 2014; Mokyr, 2011, 2016).

The remainder of this paper is organized as follows. In Section 2, we present the data sources, and we highlight two novel facts about censorship and scholars' quality. In Section 3, we develop a model linking censorship to knowledge diffusion. In Section 4, we estimate the structural model and present its implication for the role of censorship on Italy's accumulation of knowledge. The conclusion is in Section 5.

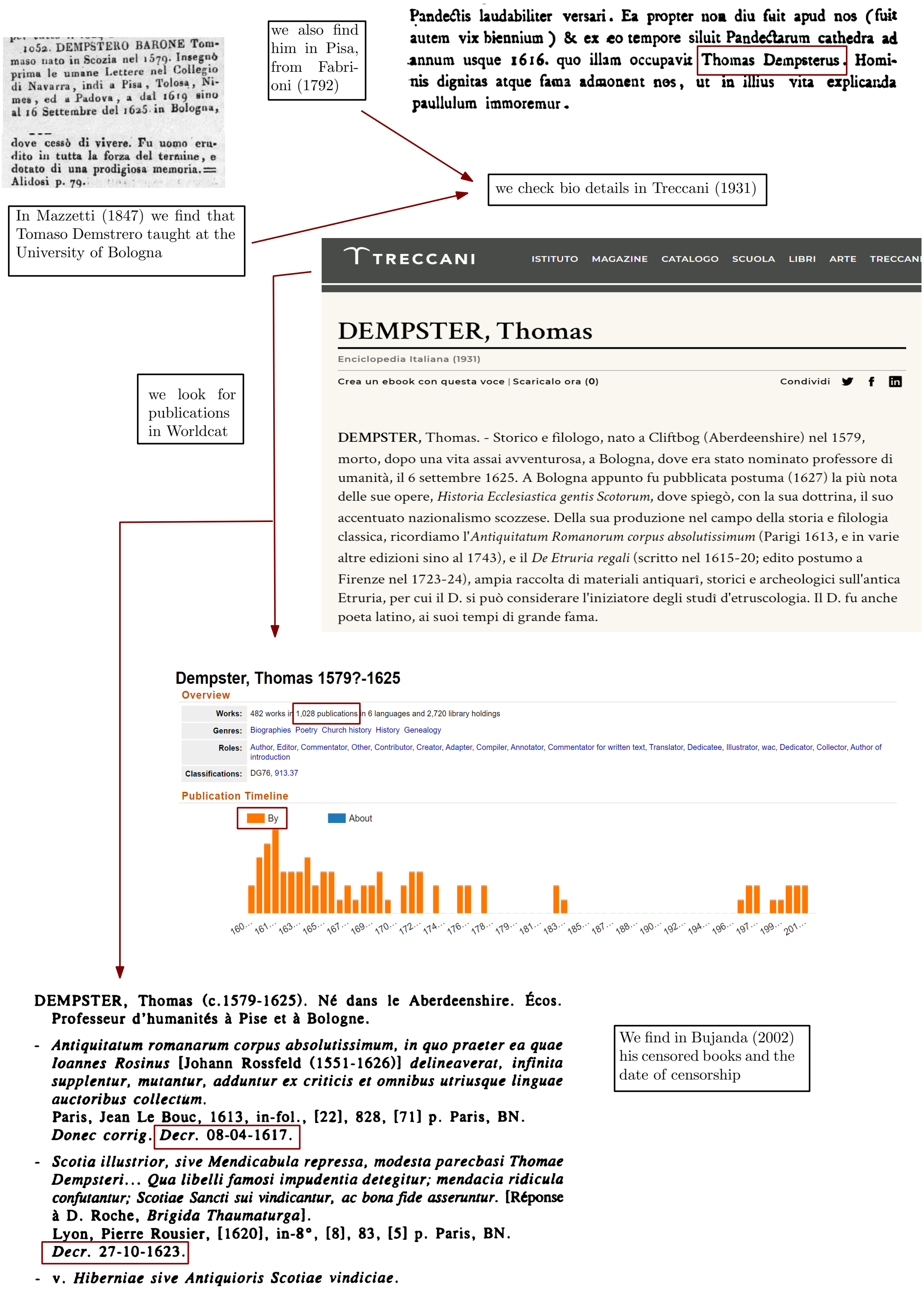
2.2 Data

2.2.1 Academies, Scholars, Publications, and Censorship

The database is built in three steps. An example is shown in Figure 2.1.

First, we collect information on all scholars who were appointed to an Italian university or were nominated to an Italian academy over the period 1450-1750. For universities, the main sources are as follows. An extensive coverage of the university of Bologna is provided by Mazzetti (1847). The university of Padova is covered by Facciolati (1757): we complete its information with the works by Casellato and Rea (2002) and Pesenti (1984). Professors at the university in Rome (Sapienza) were found in Renazzi (1803). The professors at university of Naples are covered by Origlia Paolino (1754). Pavia is another well-document university: Raggi (1879) listing all its professors. Pisa is covered in Fabroni (1791). The smaller university of Macerata also benefits from a full coverage by Serangeli (2010). For academies, we use the database "Italian Academies 1525-1700, the first intellectual networks of early modern Europe" made available by the British Library in 2013. Among the academies covered, the Gelati and the Ricovrati are two important ones. We complete these data with Parodi (1983) for the language academy "La Crusca" and with Maggiolo (1983) for a full coverage of the biggest Academy, the Ricovrati. In appendices 2.A.1 and 2.A.2 we discuss how representative are university professors and academicians in our data and how much of the Italian university/academy population is covered.

Figure 2.1 shows that Tommaso Dempstero is in the list compiled by Mazzetti (1847) of professors at the university of Bologna. We also find him in the history of the university of Pisa by Fabroni (1791), under his Latin name, Thomas Dempsterus. Checking the Italian encyclopedia Istituto dell'Enciclopedia Italiana (1929), we corroborate the information on Bologna.



1052. DEMPSTERO BARONE Tommaso nato in Scozia nel 1579. Insegnò prima le umane Lettere nel Collegio di Navarra, indi a Pisa, Tolosa, Nimes, ed a Padova, a dal 1619 sino al 16 Settembre del 1625 in Bologna,

dove cessò di vivere. Fu uomo erudito in tutta la forza del termine, e dotato di una prodigiosa memoria. = Alidosi p. 79.

Pandectis laudabiliter versari. Ea propter non diu fuit apud nos (fuit autem vix biennium) & ex eo tempore siluit Pandectarum cathedra ad annum usque 1616. quo illam occupavit Thomas Dempsterus. Hominis dignitas atque fama admonent nos, ut in illius vita explicanda paullulum immoremur.

DEMPSTER, Thomas

Enciclopedia Italiana (1931)

DEMPSTER, Thomas. - Storico e filologo, nato a Cliftbog (Aberdeenshire) nel 1579, morto, dopo una vita assai avventurosa, a Bologna, dove era stato nominato professore di umanità, il 6 settembre 1625. A Bologna appunto fu pubblicata postuma (1627) la più nota delle sue opere, *Historia Ecclesiastica gentis Scotorum*, dove spiegò, con la sua dottrina, il suo accentuato nazionalismo scozzese. Della sua produzione nel campo della storia e filologia classica, ricordiamo l'*Antiquitatum Romanorum corpus absolutissimum* (Parigi 1613, e in varie altre edizioni sino al 1743), e il *De Etruria regali* (scritto nel 1615-20; edito postumo a Firenze nel 1723-24), ampia raccolta di materiali antiquarî, storici e archeologici sull'antica Etruria, per cui il D. si può considerare l'iniziatore degli studî d'etruscologia. Il D. fu anche poeta latino, ai suoi tempi di grande fama.

Dempster, Thomas 1579?-1625

Overview

Works:	482 works in 1,028 publications in 6 languages and 2,720 library holdings
Genres:	Biographies Poetry Church history History Genealogy
Roles:	Author, Editor, Commentator, Other, Contributor, Creator, Adapter, Compiler, Annotator, Commentator for written text, Translator, Dedicatee, Illustrator, wac, Dedicator, Collector, Author of introduction
Classifications:	DG76, 913.37

DEMPSTER, Thomas (c.1579-1625). Né dans le Aberdeenshire. Écos. Professeur d'humanités à Pise et à Bologne.

- ***Antiquitatum romanarum corpus absolutissimum, in quo praeter ea quae Ioannes Rosinus* [Johann Rossfeld (1551-1626)] *delineaverat, infinita supplentur, mutantur, adduntur ex criticis et omnibus utriusque linguae auctoribus collectum.***
  Paris, Jean Le Bouc, 1613, in-fol., [22], 828, [71] p. Paris, BN.
  ***Donec corrig. Decr.* 08-04-1617.**
- ***Scotia illustrior, sive Mendicabula repressa, modesta parecbasi Thomae Dempsteri... Qua libelli famosi impudentia detegitur; mendacia ridicula confutantur; Scotiae Sancti sui vindicantur, ac bona fide asseruntur.* [Réponse à D. Roche, *Brigida Thaumaturga*].**
  Lyon, Pierre Rousier, [1620], in-8°, [8], 83, [5] p. Paris, BN.
  ***Decr.* 27-10-1623.**
- **v. *Hiberniae sive Antiquioris Scotiae vindiciae*.**

Figure 2.1: Data collection: example of Thomas Dempster

Second, for each scholar, we use the Worldcat search engine, which provides references to the collections of thousands of libraries around the world, to assign to each person all the written output he/she generated, including post mortem editions. More precisely, we count the number of "publications", including different editions of the same work. We only record publications by the author, excluding publications about the author, which are also available through Worldcat.

Figure 2.1 shows the Worldcat Page for Thomas Dempster, with the total count of publications (by or about). We can identify the two types of publications by scrapping the page. From the graph on the webpage, we can see that all publications are by him.

In a third step, we look at the list of forbidden books provided by De Bujanda and Richter (2002) and De Bujanda et al. (1996). We find an entry for Thomas Dempster, with a short biography, and the list of books that were forbidden, with the date of the corresponding decrees.

We now show some statistics on the number of scholars and on their publications. In Table 2.1 the period 1400-1750 has been divided into five periods of 70 years each. The first line covers all of Europe, from the database built by de la Croix (2021), and includes both universities and academies. Columns (1) to (5) contain the number of "published" scholars per period, i.e. those having some work referenced in Worldcat. Columns (6) to (10) show the average number of publications per person. The second line covers the subset of scholars affiliated to an Italian institution.

The number of publications per person illustrates perfectly the decline of Italy. Until period 2 (1470-1540), published scholars in Italy produced an output similar to the average European scholar. Then, a gap appears in period 3 (1540-1610) and becomes really wide in period 5 (1680-1750). The appearance of the gap coincides with the formalization of censorship through the first index published by the university of Paris in 1544, and the first Roman Index, also known as *Pauline Index*, promulgated by Pope Paul IV in 1559 (De Bujanda and Richter, 2002). Note that the Catholic Church also censored scholars who never vistited Italy, but the Church struggled to enforce censorship outside Italy.[8]

[8]One explicative example is the story of Copernicus' *De revolutionibus*, which was first published in 1546 and appeared in the Index in 1616 under the list of books to be corrected. Gingerich (2004) analyzed 277 copies of the first edition and 324 of the second of *De revolutionibus*. He found that about two-thirds of the copies of *De revolutionibus* in Italy were "corrected." [9] However, virtually none of the copies outside Italy, among which Spain and France, were touched.

Table 2.1 also shows the European numbers by individual country.[10] For countries like France, Germany and Austria we can observe that until period 2 (1470-1540) published scholars produce a similar or lower output than Italy, while a gap appears in the following periods. Note that these countries reach a level of output unknown to Italy. This fact goes against the possibility that Italy was overtaken because it had less room for growth and that Europe grew more just because of the spreading of knowledge in more remote areas. A similar pattern can be observed for Great Britain, Ireland, Denmark, and Sweden, with the caveat that we have very few observations for these countries in the first two periods. The case of Spain and Portugal is different, as these countries do not overtake Italy. This is not surprising given the intensity of the Spanish Inquisition (Vidal-Robert, 2011).

The following lines in Table 2.1 disaggregate the Italian numbers by (important) institution. The decline from period 3 to period 5 is present in the universities of Bologna, Padua, Pavia, Pisa, and in the two Roman universities. The university of Naples escapes decline in period 5 thanks to one person, Gianbattista Vico (a renowned political philosopher and historian of the Age of Enlightenment). The academies do better, in particular the Ricovrati, but this is not enough to compensate for the overall decline at the Italian level.

One can argue that the decline in knowledge production in Italy might be due to a weakening of the barriers to entry into the scholar profession. In fact, if published scholars are positively selected and the barriers to entry weaken, the average quality of scholars goes down. One way to control for this problem is to look at the dynamics of top scholars, who are less affected by changes in the barrier to entry. Hence, in table 2.A.2 in the appendix, we show that Italy loses to Europe in terms of knowledge production also if we consider only scholars whose Wikipedia page is longer than 5000. Moreover, in appendix 2.A.3 we show that Italy is overtaken by Europe within all the scholars' fields that we are able to identify, ruling out the possibility that the this observation was driven by a composition effect across fields.

2.2.2 Two Features of Authors Censorship

On May 23 1555, a new Pope was elected. Cardinal Caraffa became Paul IV. This election witnessed the return of the conservatives. In 1559, Paul IV had published the first long

[10]We did not show the results for all European countries because some have too few observations or contained scholars coming from one University/academy only (this is the case of Belgium and the University of Louvain).

Table 2.1: Total number of scholars & publications by period

	Number of published scholars					Mean publications per person				
Period	1	2	3	4	5	1	2	3	4	5
Europe	525	1274	3186	4435	5862	159	279	318	232	236
France	64	144	493	796	871	256	214	337	263	343
Germany & Austria	77	335	892	1059	2127	138	256	398	277	250
Great Britain & Ireland	14	46	158	328	753	25	697	248	393	382
Denmark & Sweden	1	3	45	119	215	2	722	228	150	231
Spain & Portugal	13	66	271	243	172	415	161	166	105	70
Italy	297	531	969	1167	766	161	272	238	160	111
Ubologna-1088	53	81	74	64	41	74	184	225	84	46
Unapoli-1224	4	20	31	14	19	632	123	57	39	245
Upadua-1222	59	116	146	85	75	75	179	175	140	63
Upavia-1361	35	68	59	22	8	113	170	159	63	46
Uroma-1303	32	67	53	47	39	518	420	490	164	91
Upisa-1343	1	43	55	65	37	4	92	123	79	63
UromaGregoriana-1556			40	62	49			416	253	62
Umacerata-1540			25	18	14			26	15	29
AcadRicovrati-1599			39	107	188			51	162	183
AcadCrusca-1583			14	98	100			136	126	153
AcadGelati-1588			8	71	27			193	116	57

Note: periods: 1:1400-70, 2:1470-1540, 3:1540-1610, 4:1610-80, 5:1680-1750 list of prohibited books, the Index. The idea was refined further by the Council of Trent, which established in 1564 the Index librorum prohibitorum. The Index comprised three parts. The first part contained the name of the heretical authors whose entire output, past and future, was condemned (opera omnia, the works). The second part contained a list of censored publications by authors who still belonged to the Church. The third part dealt with anonymous publications.

This attempt to control publications by the Catholic Church is probably the biggest experiment in the whole history of censorship. To read or to keep censored books could lead to excommunication and eternal damnation. It lasted four centuries, as the last version of the Index was published in 1948, still affecting the second authors' grandparents.

The establishment of the Index follows from a change of attitude of the Church concerning novel ideas, including scientific ones. One aim was to defend Aristotelian physics against alternative views. Aristotle had to be defended, because his theory on substance and accidents is compatible with the dogma of transubstantiation according to which, during the Mass, the bread and wine are literally converted into the body and blood of Christ at the consecration. Everything that goes against Aristotle's physics became suspect.

The Copernicus case best illustrates the reversal of attitude. The idea of his heliocentric system was developed around 1505, and first written in an unpublished book intended for his friends. The Pope Clement VII learned about these ideas in 1533 and liked them. Several highly ranked clerics asked Copernicus to publish his treaty. One advantage of Copernicus system was to provide more accurate computations for astronomical events. Then, after the conservative revolution, Copernicus writings were blacklisted. What appeared to be a legitimate hypothesis in 1543 became in 1616 a foolish thesis, absurd in philosophy, and formally heretic. The Church took more than three centuries to accept heliocentrism and remove Copernicus' works from the Index in 1846.

The Church's fight did not spare the most notable forerunners of the varied flow of novel ideas that spread all over Italy and Europe. Galileo Galilei was condemned, and his books were censored not only for his astronomical views, but also for his support of atomism. According to atomism, the physical world comprises fundamental, indivisible components known as atoms, violating the Aristotelian view of a continuous matter. Atomism and its proponents, such as the French philosopher Descartes, were censored by the Church until at least the beginning of the eighteenth century. In a world where religion and philosophy were intertwined with natural sciences, the aversion towards atomism is likely to have affected scientific knowledge. Perhaps it is not a case that the particle theory of light, which relies on an atomist view of the matter, was developed by Newton and not by an Italian.

The Church's fight had some consequence on how to think about the continuum, indivisibles, and the actual infinite. The Jesuits were particularly active in these mathematical controversies, fighting against the idea that a continuous line is composed of distinct and infinitely tiny parts (Alexander, 2014). In his book, Alexander (2014) considers what the world would have been like without infinitesimals. "If the Jesuits and their allies had had their way, there would be no calculus, no analysis, nor any of the scientific and technological innovations that flowed from these powerful mathematical techniques." Now, this is perhaps exaggerated, and Alexander claims more than he is able to prove. Grabiner (2014) defends the view that seventeenth-century mathematics had far too much momentum and too many demonstrable successes to be stopped by philosophical arguments about the nature of the continuum.

Another landmark of the reversal in the attitude of the Church is the censorship of all the works by and the burning at the stake of Giordano Bruno. Bruno had accumulated many reasons to be condemned to death, but one point of his theory that did not fit at all with the Church's view was the theoretical possibility of an infinite universe and the plurality of worlds. Bruno has become the symbol of the scientist persecuted by religious authorities. Other times authors were punished with imprisonment. For example, Galilei was sentenced to house arrest for the rest of his days.

Looking at the data in the Index librorum prohibitorum, one should admit that censorship does not necessarily imply that the author risks his life. While sometimes, as for Bruno and Galilei, censorship went together with severe consequences for the author, in other cases, the consequences were mild. For example, the poet John Barclay's name, whose works contained satirical descriptions of the Jesuit school, appeared in the Index in 1608. Upon invitation by the Pope himself, he went to Rome in 1616 and resided there until he died in 1621. Moving to Rome was a way to testify that he was a good Catholic and avoid further consequences. Not all his writings were blacklisted, and he was able to publish again after he was first censored. In other cases, there were no consequences for the author simply because the heresy was identified after her/his death. This is the case of Bernardino

Ciaffoni, who used to be the rector of Rome's college San Bonaventura. He died in 1684 but was censored in 1701 because his works contained insulting claims against the Jesuits. Scholars developed different strategies to avoid consequences because of their writings. Many authors used pseudonyms to protect themselves. This is the case of Copernicus, who first exposed his theories anonymously in the *Commentariolus*. Only once he realized that his work was well-received, he revealed his identity by writing his theory under his real name (Rosen, 1977). In sum, censorship did not always bear consequences for the authors, while posterity indeed paid for more complicated access to the revolutionary's wisdom, at least that embodied in forbidden books.

Being a clergyman was not a safe antidote towards censorship. One particularly striking case is Serry Jacobus Hyacinthus. Professor in Padova, he contributed to the Dominicans-Jesuits controversy on grace, and several of his works appeared on the index. Not only he was a Dominican, but also he was a member of the *Congregation of the Index*, the organ responsible for the creation and management of the Index. Censorship did not spare even the members of the company of Jesus, who had a primary role in the Counter reformation and who were the "*soldiers* of God [...] for the *defense* and propagation of the faith.[11]" In our database, 10 out of 173 published scholars belonging to the Jesuit university Gregoriana were censored. Among them, Achille Gagliardi was censored in 1703 for his writings about the annihilation of the will during mystical states. These ideas were found not compatible with free will, which is a cornerstone of Catholic theology.

We now describe the impact of censorship quantitatively. Figure 2.2 shows how authors belonging to our dataset are distributed according to the number of their publications. In red, we mark authors who were censored at least once, and in green non-censored authors. We provide five histograms, one for each period. Censorship started at the end of the second period, but also hit works that were published in the past. From these five histograms, it is clear that censorship was concentrated on top scholars for the first two to three periods, and then becomes more uniformly distributed over the quality of scholars. Or, as we wrote earlier, once censorship was introduced, censored authors were of better quality than the non-censored authors, but this gap shrunk over time.

This shift in the identity of who is hit by censorship reflects behavioral changes. The top

[11]This is a translation of the words of the *Exposcit Debitum* Papal bull, that gave to the foundation of the order in 1550.

scholars who had the potential to publish non-compliant ideas and become famous (as in the first three periods) decided to be more compliant, and published conventional material instead. Bruno, Copernicus and Galilei were at the top of the distribution and were all censored, and sometimes burned. Their reincarnation in the last two periods might have preferred to be mediocre poets.

Table 2.2: Moments per period

Moment description	Period				
	1400-70	1470-1540	1540-1610	1610-80	1680-1750
Number of authors	192	376	745	755	634
Log publications per scholar, median (1)	4.55	4.54	4.13	3.53	3.48
Log publications per (censored) scholar, median (2)	7.64	7.07	6.64	5.62	5.89
Gap in median publications (2)-(1)	3.09	2.53	2.51	2.09	2.41
% censored scholars	7.29	11.17	7.25	6.49	4.26
Log publications per scholar, 75^{th} percentile	5.93	6.04	5.44	5.1	4.98
Log publications per (censored) scholar, 75^{th} perc.	7.80	7.86	8.02	7.14	6.67

We show in Table 2.2 the key moments of these distributions. It confirms what we expected from the figures: the gap in median publications between censored authors and all authors shrunk from about 3 to 2.4 (the numbers should be interpreted as log of number of publications). The Table shows two additional features. First, the percentage of censored authors is shrinking over time. Second, overall quality, measured by median publications per person, is declining over time as well. This also holds for the top of the distribution, as the 75th percentile also diminishes. Those two trends are very much compatible with the idea of top innovative people becoming progressively compliant and mediocre over time.

Our data also allows looking at possible geographical patterns in censorship. Figure 2.3 shows the place of birth of the scholars in the database, distinguishing the censored (red) from the not censored (green) ones. Geographical coordinates have been slightly randomized, so that people born in cities still appear distinctly. From Italy's map, we retain that our data cover the whole peninsula and its island, without obvious bias. Moreover, censorship seems to affect all regions.

Some members of Italian universities and academies were born outside Italy (as Thomas Dempster in our example above). Hence the interest in having a map of Europe. Figure 2.4

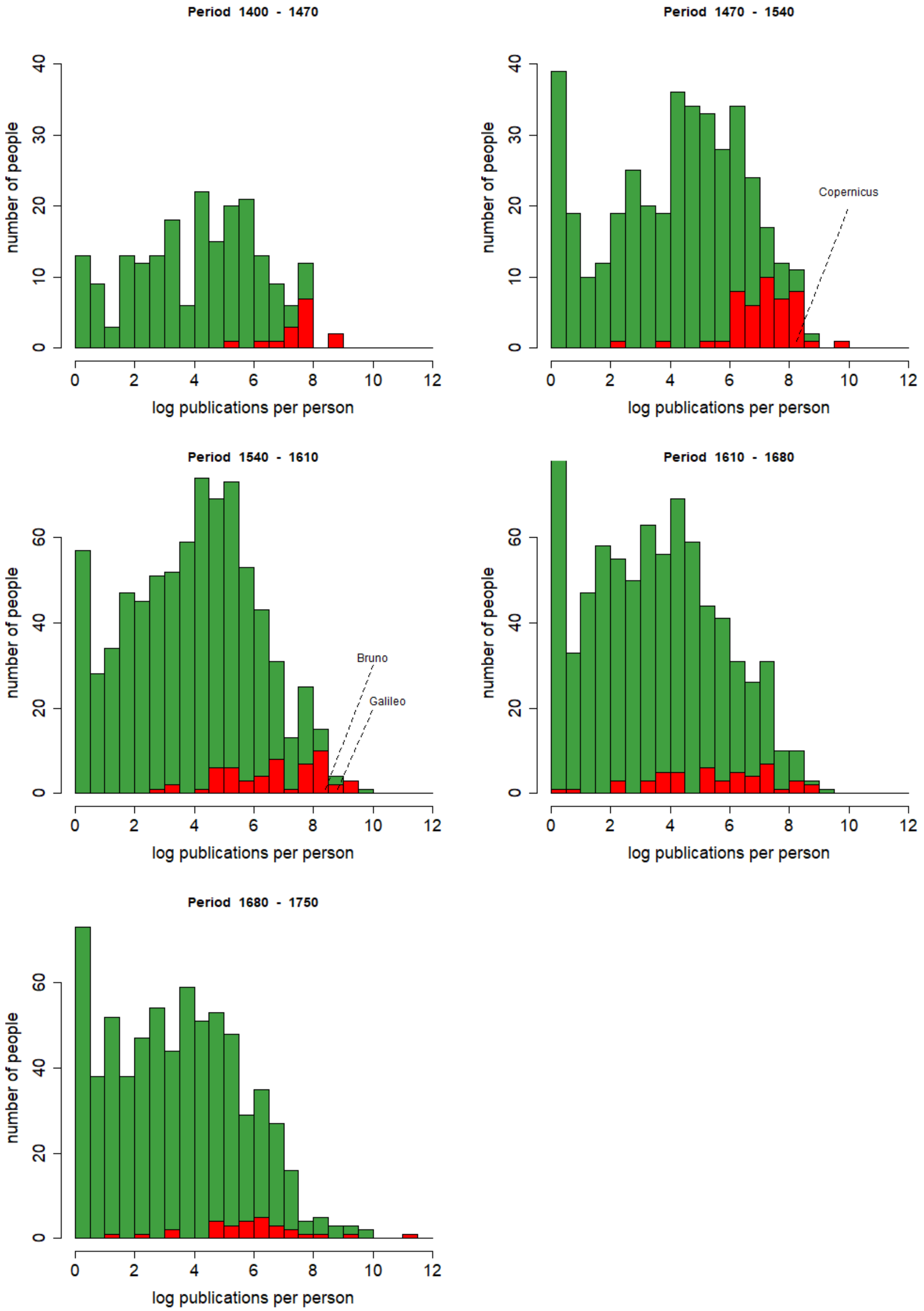


Figure 2.2: Distribution of published authors by quality. Red: censored. Green: non-censored.

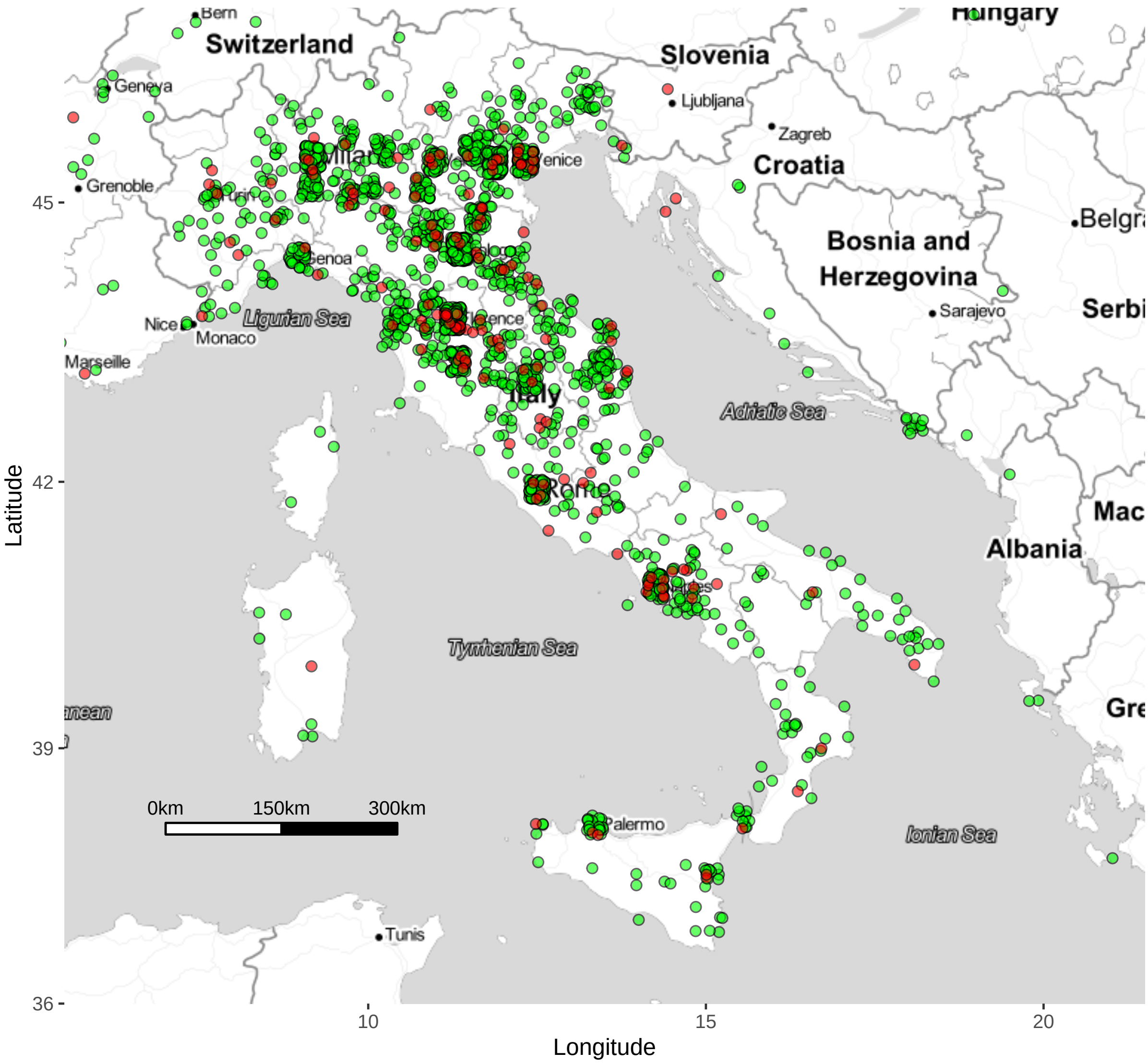


Figure 2.3: Place for birth of censored (red) and non censored (green) members of Italian universities & academies – Italy.

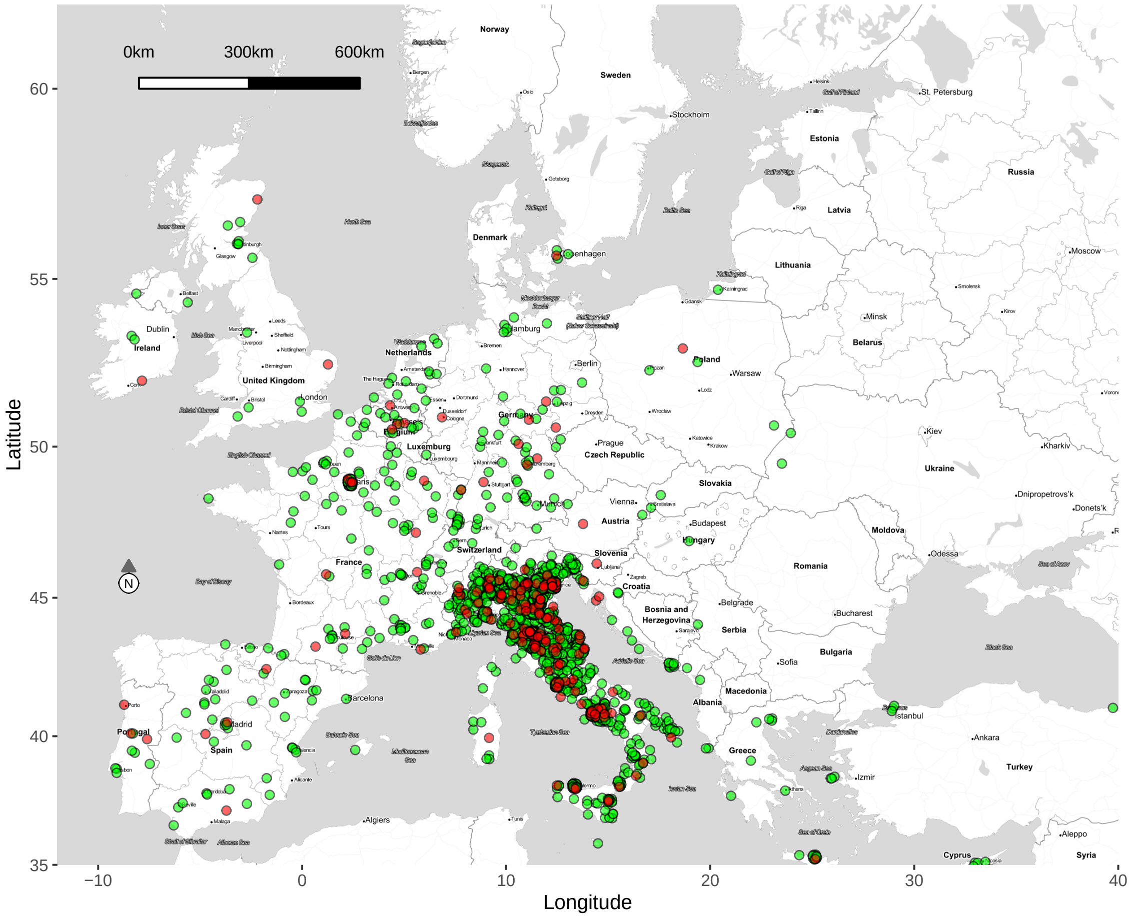


Figure 2.4: Place for birth of censored (red) and non censored (green) members of Italian universities & academies – Europe.

provides a European view of the places of birth of our scholars. Some of them are foreign members (or corresponding members) of some academies, such as the Ricovrati. They might have never come to Italy, requiring us to have a specific robustness exercise excluding those foreigners.

2.3 Occupational Choice and Knowledge Diffusion

In this section, we develop a theory of accumulation of knowledge and occupational choice. Authors, building on the knowledge left by the previous generation, write books that can be compliant to the Roman Church's ideology or revolutionary (in the sense of the humanistic and scientific revolutions). Printers decide whether to be active in the revolutionary or compliant sector. They make this choice according to the quality of the books of each type that they encounter. Therefore, if the revolutionary knowledge grows faster than the compliant one, revolutionary books' share will also increase. The Roman Church dislikes revolutionary ideas and might decide to censor them, which would decrease their share but also alter the accumulation of the total stock of knowledge in the economy.

2.3.1 Knowledge Diffusion

Time is discrete. At each date t one generation is alive. Knowledge is embodied in books and is transmitted between the successive generations through them.At the beginning of each period, the individuals first learn from μ books. μ is a parameter representing the number of books one can read during her life. Books include more or less relevant content to produce goods and services. Irrelevance of a book read by person i is denoted h_i. The quality of a book is a decreasing function of its irrelevance, with elasticity θ:

$$q_i = h_i^{-\theta}, \quad \theta \in (0,1). \tag{2.1}$$

Books are of two types, which define different distributions from which their relevance is drawn. *Compliant* books, indicated by the superscript C, embodies the type of knowledge which is acquiescent with the ideology of the Roman Church.[12] *Revolutionary* books, denoted by the superscript R, indicates that the knowledge is considered to be heretical by the Roman Church. Taking examples from Alexander (2014), geometry books would be compliant while books using infinitesimal calculus would be revolutionary. Both of them are of variable quality, what we call relevance.

At the beginning of time t, the irrelevance of book i of type j follows an exponential distribution

$$h_i^j \sim \exp(k_t^j), \quad \text{with } j \in \{C,R\} \text{ and } i \in \{1,..,N\}. \tag{2.2}$$

Note that the scale parameter k_t^j depends on the book type. As

$$\mathrm{E}[h_i^j] = \frac{1}{k_t^j},$$

k_t^j measures the average usefulness of knowledge in sector j.

Since the irrelevance of books is exponentially distributed and given Equation (2.1), the distribution of book quality follows a Fréchet distribution, see Appendix 2.C.1. This allows

[12]Note that being compliant does not necessarily mean to produce using the official Roman Church doctrine as an input: this is true just for the production of religious books or religious services in general. Instead, it just means that the knowledge should not contradict the Roman Church doctrine.

us to write the average book quality q^j by sector as:

$$E(q_i^j) = \text{cst}(k^j)^\theta \text{with } j \in \{C, R\} \text{ and } i \in \{1, .., N\}. \quad (2.3)$$

The number of revolutionary books that each agent will read in t depends on their availability in bookshops. The share of printers that produced revolutionary books in the previous generation is denoted by m_t. Therefore, a individual will read $\lfloor \mu m_t \rfloor$ revolutionary books and $\lfloor \mu(1-m_t) \rfloor$ compliant books, draw from their respective distribution. Each individual retains the best book coming from each one of the two distributions.[13] Formally, the process of retaining the best books by sector is described as

$$\begin{aligned} \hat{h}_i^C &= \min\{h_1^C, .., h_{\lfloor (1-m_t)\mu \rfloor}^C\}, \\ \hat{h}_i^R &= \min\{h_1^R, .., h_{\lfloor m_t \mu \rfloor}^R\}. \end{aligned}$$

For the sake of simplicity, from now on we will approximate $\lfloor (1-m_t)\mu \rfloor$ and $\lfloor m_t \mu \rfloor$ to respectively $(1-m_t)\mu$ and $m_t\mu$, so that we will be able to proceed our analysis treating the number of books read as a continuous variable.

Note that the exponential distribution satisfies the minimum stability postulate: if x and y are mutually independent random variables, exponentially distributed with parameter λ, then $\min(x, y)$ is exponentially distributed with parameter 2λ. Hence, we have:

$$\begin{aligned} \min\{h_1^C, .., h_{(1-m_t)\mu}^C\} &\sim \exp(k_t^C(1-m_t)\mu), \quad \text{and} \\ \min\{h_1^R, .., h_{m_t\mu}^R\} &\sim \exp(k_t^R m_t \mu). \end{aligned}$$

We can now deduce that the distribution of actual relevance of the best book read by person i follows

$$\hat{h}_i^j \sim \exp(b_t^j), \quad \text{with } j \in \{C, R\}, \quad (2.4)$$

[13]It is reasonable to think that better books have more readership. This possibility can be embodied in our model by reinterpreting parameter μ. Assume that readers visit a number B of bookshops and buy the best among $\hat{\mu}$ books for each bookshop. Then, readers retain the best book among all the books they bought. This situation is equivalent to our model if $B\hat{\mu} = \mu$ and high-quality books will have relatively more readership.

where b_{t+1}^C and b_{t+1}^R are defined as

$$b_{t+1}^C = k_t^C(1-m_t)\mu,$$
$$b_{t+1}^R = k_t^R m_t \mu.$$

Later in life, the generation t writes new books, combining their inherited knowledge with a new idea. This new idea is drawn from a distribution whose scale parameter depends on the average quality of the books they have read:

$$h_N^j \sim \exp(\nu b_t^j), \quad \text{with } j \in \{C, R\}.$$

Taking the best of their acquired and new knowledge leads to a book with irrelevance distributed as:

$$\tilde{h}^j = \min(h_N^j, \hat{h}^j) \sim \exp((1+\nu) b_t^j).$$

We can now summarize the dynamics of the two types of knowledge by the dynamics of the scale of their distribution:

$$k_{t+1}^C = (1+\nu)k_t^C(1-m_t)\mu, \tag{2.5}$$
$$k_{t+1}^R = (1+\nu)k_t^R m_t \mu. \tag{2.6}$$

2.3.2 Occupational Choice

To finish describing the dynamics, we need to define how the share of printers producing revolutionary books evolves over time. We suppose that printers meet authors randomly, but have to decide where to be active in the compliant sector or in the revolutionary sector at the beginning of their activity. Once they have chosen a sector,[14] they would print any author they meet randomly. They will thus determine their sector of activity based on the first author they meet. This author has written books of relevance $\tilde{h}^C$ and $\tilde{h}^R$. Printers decide their sector taking into account both the relative relevance of the two books. Printers also take into account that customers of the bookshop might value differently two books

[14] Assuming that printers have to choose a sector is consistent with Dittmar and Seabold (2015). In Germany, the official city printers were not advocates of the Reformation because they "did not want to endanger official work orders or antagonize city governments." Moreover, according to Grendler (1975), printers in Venice faced the risk of having their bookshops in Rome sized by the Vatican if they printed revolutionary content, which implies that they had to choose a sector.

with the same quality that belong to two different sectors. This might happen because of consumer preferences or for how book quality translates into consumption goods[15]. We summarize these two effects assuming that the relative price at which revolutionary books are sold is represented by p. The probability that the revolutionary book is best is:

$$\text{Prob}\{q^C < pq^R\} = \text{Prob}\{\tilde{h}^C > p^{-1/\theta}\tilde{h}^R\} = \frac{b_t^R}{b_t^R + b_t^C p^{-1/\theta}} = m_t. \tag{2.7}$$

Using the law of large numbers, this probability also defines the share of printers active in the revolutionary sector m_t. From now on we will refer to $\hat{p}$ as $\hat{p} = p^{-1/\theta}$.

The dynamics of knowledge quality (2.5) and (2.6), together with the occupation choice

$$m_t = \frac{k_t^R}{k_t^R + \hat{p}k_t^C} \tag{2.8}$$

and initial conditions k_0^C and k_0^R, determine m_0 and the equilibrium path $\{m_t, k_t^C, k_t^R\}_{t\geq 0}$.

2.3.3 Censorship

So far, the Church did not play any role in the model. As we discussed in the introduction, there is historical evidence that the Roman Church tried to limit the spread of revolutionary books issuing the *Index Librorum Prohibitorum*. We model this behavior of the Church, assuming that she can interfere with the process of occupational choice imposing a rate of censorship on revolutionary books. More precisely, she can limit the number of revolutionary titles that an author can read, making unavailable a fraction β of the volumes that she would have read without censorship. Formally, the process of censorship limits the number of revolutionary books that individuals encounter during their life to $\mu m(1-\beta)$ and therefore alters the process of accumulation of revolutionary knowledge, which now follows

$$k_{t+1}^R = (1+\nu)(1-\beta)k_t^R m_t \mu, \quad \text{with } \beta \in [0,1]. \tag{2.9}$$

Note that in this way, the Church can directly decrease the share of revolutionary books m and will also make it less likely that revolutionary works will be written in the future. This

[15] Books can be used to produce consumption goods, and books belonging to different sectors can have different productivity in this respect. For example, the production of consumption goods through books can be represented as $c = \alpha \sum^{N_R} q_i^R + \sum^{N_C} q_i^C$, where α would be the relative productivity of revolutionary books quality, while N_R and N_C are respectively the number of revolutionary and compliant books owed by the customer.

is because the process of accumulation of revolutionary knowledge slows down.

History tells us that book's censorship was not the only way the Church used to limit the spread of revolutionary books: in fact, in the second half of the 16th century, the Roman Church developed a system of tribunals, called *Roman Inquisition*, aimed at persecuting both authors and printers accused of heresy. This institution affected the work of scientists and thinkers. One notable example is the process of Galileo Galilei, who was tried by the Inquisition in 1633. The Inquisition matters for our analysis because it can slow down the accumulation of revolutionary knowledge through self-censorship: even if one author writes a high-quality revolutionary book, she still might prefer not to hand it in it to the printer for fear of being processed by the Inquisition. Similarly, even if the best books are revolutionary, the printer might still prefer to be compliant for the same reason. This mechanism can be easily incorporated in our framework, assuming that the Inquisition makes publishing and writing revolutionary books less desirable. Individuals take this into account discounting q^R by a factor $\gamma \in [0,1]$. We can also interpret γ as the probability that authors decide not to write revolutionary books or that printers do not publish them for fear of receiving punishment. Under this new mechanism, the probability that a printer chooses the revolutionary sector is:

$$\text{Prob}\{q^C < \gamma p q^R\} = \text{Prob}\{\tilde{h}^C > (\gamma p)^{-1/\theta} \tilde{h}^R\} = m_t. \tag{2.10}$$

2.3.4 The Dynamics under an Exogenous Church's Behavior

So far we mentioned that the Church can limit the share of revolutionary books through censorship, but we did not mention how the Church is choosing β. Clearly, the choice of β over time will depend on the behavior of agents described in the previous section and on the objective of the Roman Church. On the one hand, the Church wanted to have the smallest possible number of heretical books circulating to maintain her power. On the other hand, we do not know what prevented her from imposing the highest level of censorship in any period. The Church was probably trading-off censorship with other reasons. These reasons could be time spent in other activities of her interest, or the fact that an overly harsh censorship could create damage to the Church herself,[16] or something else.

[16] As an example, we can think that if the censorship is overly harsh, the Roman Church might lose in terms of competition with the Protestant Church. This reasoning is plausible if devotees dislike a too harsh censorship. While rulers had the final say about the religion of their territory, their decision was not completely

Here we treat β as is it was exogenous, and we study the dynamics under this assumption. We start defining $z = k^R/k^C$: note that the share or revolutionary ideas m can assume one and only one value given z, which means that once we know the dynamics of one of the two variables, we also know the dynamics of the other. From equation (2.8) we get

$$m_t = \frac{z_t}{\hat{p} + z_t}. \tag{2.11}$$

We decided to make m_t rather than z_t our main variable for describing the model dynamics because its domain is a bounded set. The dynamics of m are defined formally below.

Definition 1. *Given β, an equilibrium path is a sequence $\{m_t\}_{t\geq 0}$, describing the share of revolutionary books over time. It is such that:*

- *Each author of each generation write books whose quality and type is defined by the current state of knowledge.*
- *Each printer of each generation chooses her sector according to the most productive book presented by the first randomly met author.*
- *Each printer of each generation, once she chose her sector, prints all the authors she meets randomly.*
- *The probability of being exposed to revolutionary book in $t+1$ depends on the share of revolutionary titles written in t.*
- *The books printed in t embody the stock of compliant and revolutionary knowledge available to generation $t+1$.*

The equilibrium described in definition 1 depends on the whole theory that we described in the previous subsection, but we are able to summarize in a single equation the law that governs the dynamics of m. Dividing Equation (2.6) by (2.5) side by side, and substituting the resulting z_{t+1} in (2.11) at time $t+1$, we get the equation that governs the equilibrium dynamics of m:

$$m_{t+1} = \frac{(1-\beta)m_t^2}{1 - m_t((\beta-2)m_t + 2)} = f(m_t). \tag{2.12}$$

independent from the common people's believes. Protestantism could spread thanks to the invention of the printing press, which aroused popular support by distributing pamphlets Eisenstein (1980); Rubin (2014). Probably it would not be the best choice for a ruler to impose Catholicism if a large majority of its population already converted to Protestantism.

Equation (2.12) together with an initial condition m_0, allow us to determine the equilibrium path $\{m_t\}_{t\geq 0}$. The equilibrium path $\{m_t\}_{t\geq 0}$ satisfies:

Proposition 1. *Given the initial condition* $m_0 \in [0,1)$, *we have*

i) $\lim_{t\to\infty} m_t = 0$ *if* $m_0 < 1/(2-\beta)$ *(Compliant steady state),*

ii) $\lim_{t\to\infty} m_t = 1$ *if* $m_0 > 1/(2-\beta)$ *(Revolutionary steady state),*

iii) $\lim_{t\to\infty} m_t = m_0$ *if* $m_0 = 1/(2-\beta)$ *(Unstable steady state).*

Proof. See Appendix 2.C.2 □

2.3.5 The Dynamics under an Optimizing Church's Behavior

In the previous subsection, we described the dynamics under a constant rate of censorship β_t. A simple way to go beyond this approach would be to assume a rule of thumb behavior of the type: the Church chooses the lowest rate of censorship that allows converging to a world with no revolutionary ideas. We analyzed this case in Appendix 2.C.3. This approach has two main shortcomings. Firstly, it is stringent in defining how the Church trades-off the prevalence of revolutionary books and censorship. Secondly, it leaves unexplained the discontinuity over time in censorship and its delay with respect to the rise of Protestantism. Here we propose a model that can address this additional fact and also explain the two features of authors censorship that we illustrated in Section 2.2.2. We assume that setting up an apparatus able to create a list of forbidden books and enforce its application represented a large fixed cost to bear for the Church. The Church cannot enforce any censorship before having paid a fixed cost ψ. After having paid ψ, she can impose a rate of censorship up to $\overline{\beta}$. The Church cares about the share of compliant books in the economy: her utility function is given by $u()$, which is differentiable, bounded, and strictly increasing in $1-m_t$, while $\delta < 1$ is the discount factor. We can now define the value function of the Church recursively. In the case she still did not establish a censorship structure yet, her value function is

$$V(m_t) = \max[V^N(m_t), V^C(m_t) - \psi],$$

where V^N is the value of not imposing censorship and equals

$$V^N(m_t) = u(1-m_t) + \delta V(m_{t+1})$$
$$\text{s.t.} \quad m_{t+1} = f(m_t, 0) = \frac{m_t^2}{1-m_t(-2m_t+2)},$$

while V^C is the value of having a censorship apparatus set up and equals

$$V^C(m_t) = \max_{0\leq\beta_t\leq\overline{\beta}} u(1-m_t) + \delta V^C(m_{t+1}),$$
$$\text{s.t.} \quad m_{t+1} = f(m_t, \beta_t) = \frac{(1-\beta)m_t^2}{1-m_t((\beta-2)m_t+2)}.$$

We can write this last value function like that since $V^N(m_t)$ equals $V^C(m_t)$ if $\beta = 0$ is chosen. Moreover, it is straightforward to see that, once ψ has been paid, the Church will always set β_t to its maximum level[17]. In this model, the Church has to choose between paying a fixed cost today for enjoying a lower share of revolutionary books in the future and postponing such payment. Postponing censorship would be less costly because of discounting, but it would also imply a higher share of revolutionary books in the future. This trade-off implies that the Church would be more prone to implement censorship immediately when the fixed cost ψ is low and when the effectiveness of censorship $\overline{\beta}$ is high. Moreover, she is less likely to start censoring the more she is impatient. When $\delta = 0$, the Church cares only about what happens in 0, and therefore she will never pay a cost ψ that affects only the future share of revolutionary books. The Church's decision to start censoring also depends on the initial level of revolutionary books m_0. In fact, m_0 influences the dynamics with and without censorship. To understand why the initial condition matters, consider the extreme case $m_0 = 0$. Proposition 1 states that in this case, m stays constant over time, regardless of the value of $\overline{\beta}$, which makes censorship useless. Proposition 2 allow us to understand better when it is not optimal for the Church to censor:

Proposition 2. *If $\psi > 0$, then there exist $\tilde{m} > 0$ and $1 > \check{m} > 0$ such that*

i) If $m_0 < \min(1/2, \tilde{m})$ then $\beta_t = 0$ for each $t \geq 0$ (No need to censor),

ii) If $m_0 > \max(1/2, \check{m})$ then $\beta_t = 0$ for each $t \geq 0$ (Too late to censor).

[17]This holds because $\partial f(m_t, \beta_t)/\partial\beta_t \leq 0$ and $\partial u(1-m_t)/\partial m_t < 0$, which implies $\partial V^C(m_t)/\partial\beta_t \geq 0$.

Proof. See Appendix 2.C.4. □

Proposition 2 makes the point that for some m_0 it can be optimal for the Church to never impose Censorship, which can be due to opposite reasons. In fact, for a low enough m_0, the Church knows that revolutionary ideas would naturally disappear. Therefore, there is no need to censor. Symmetrically, when m_0 is large enough, the Church knows that even imposing censorship, she would converge fast to the revolutionary steady state. In this case, it is too late to censor. Proposition 3 improves further our understanding of the Church's censoring behavior.

Proposition 3. *There exists $\overline{\psi}$ such that for each $\psi < \overline{\psi}$, there also exists $\overline{m}, \hat{m}$ such that for $\hat{m} > m_0 > \overline{m}$, $\beta_0 = \overline{\beta}$ holds (Window of censorship).*

Proof. See Appendix 2.C.5. □

Proposition 3 tell us that the areas under which the Church is willing to censor are not isolated points, but form *windows* of the domain of m. This result is intuitive if we think that two conditions should hold to make the Church willing to censor. First, censorship should be able to alter the time path of revolutionary books significantly. Second, the Church should lose the opportunity to change the equilibrium path dramatically if she waits for one additional period. Censorship decisions are not taken in isolated points because contiguous points have a very similar value of postponing censorship and of altering the dynamics of m.

Note that we could not characterize a closed form of the equilibrium time path $\{m_t\}_{t\geq 0}$. Censorship *windows* can be placed anywhere in $[0,1]$ unless some strict assumptions are made. The model leaves open the possibility that revolutionary ideas were growing or declining before the Church implemented censorship. One test of the (estimated) model would be to find that revolutionary ideas were growing before censorship. This possibility would be consistent with the historical fact that the Protestant reformation started before the first issue of the Index.

2.3.6 Discussion of Model Assumptions

Our model of censorship introduction under an optimizing Church's behavior relies on a set of assumptions that allows us to make it tractable. In this subsection, we discuss such assumptions, and we compare them with some alternative modeling choices that we could have taken.

One shot fixed-cost of censorship A different modeling choice that we could have taken is to assume that there is no one-shot fixed cost ψ to set up censorship, but rather a convex cost that depends on the size of the chosen rate of censorship β_t.[18] A model based on this alternative assumption would have a hard time fitting the data for two reasons. First, it would predict a gradual increase or decrease in the share of forbidden books, while in the data there is a discontinuous increase in censored books given by the creation of the Index. Second, it would generate a faster reaction to the diffusion of Protestantism than the one observed in the data. This is because revolutionary authors' dynamics display inertia, which makes acting soon with a low rate of censorship a good idea. Finally, the one-shot nature of the cost ψ helps to rationalize why the Church kept updating the index until the 20^{th} century. The Church would have removed censorship much sooner if she had to pay ψ each period. In fact, once censorship can shift dynamics towards the compliant steady state, the gains of censorship decrease rapidly.

Maximal level of censorship A point that is worth discussing is why the Church is bounded above by $\overline{\beta}$ concerning the level of censorship that she can impose. We assume this for two main reasons. First, the process leading to censorship was largely bottom-up and grounded on external denounce.[19] If the arrival rate (frictions) of new books to be checked is low enough, then the Church can not have the opportunity to censor all revolutionary books. This mechanism explains why many books were censored decades after being first published. It also hints at why some books might have never been censored. Also, it justifies why we assume that the Church censor a *share* and not a *number* of censored authors. The latter would make sense if the Church's bottleneck was a too large number of books to be checked, which is unlikely to be the case.[20] Second, dissimulation to avoid censorship was far from uncommon Spruit (2019). Heretic authors could cloak their dissident beliefs either by pretending to comply with the Church (*simulatio*) or by hiding their heterodox views to authorities (*dissimulatio*). Decartes' quote "Like an actor wearing a mask, I come forward,

[18]If the cost was concave enough, the model would behave similarly than with the fixed cost.

[19]By external denounce, we mean that the Congregation of the Index did not initiate the process most of the time. Wolf (2006) enumerate members of the clerics, aristocrats, and bourgeoisie as the categories of people who were bringing suspicious books to Rome to denounce them.

[20]The Congregation of the Index was formed of eight cardinals only in 1587, while it previously had five members. Moreover, Wolf (2006) observes that between 1571 and 1596, the number of yearly meetings of the Congregation oscillated between 0 and 34. The Church could have easily increased these numbers if needed.

masked, on the stage of the world," means that he was conscious of the risks ahead of him and found in dissimulation a valuable tool to overcome them Snyder (2012). Since books' revolutionary content was seldom hidden, it is reasonable to think that the Church could identify only a share of the heretic books.

Censorship enforcement One additional assumption we made is that the Roman Inquisition (the police) was able to enforce the application of the Index outside the Papal State at a constant rate over time. The primary weapon of the Church for enforcing the application of the Index was a dense peripheral organization of censorship. It was constituted by 41 inquisitorial tribunals, primarily distributed in northern Italy, and by bishops in the center-south (Balsamo and Fragnito, 2001).[21] The Church had an additional economic weapon for enforcing censorship: in Venice, the Church enforced compliance threatening printers to seize their stores within the Papal State (Grendler, 1975). The Church could enforce censorship throughout the XVIII century, even though carrying out this task became increasingly more complicated (Prosperi et al., 2010). In fact, under the papacy of Pope Benedict XIV, which started in 1740, some inquisitorial tribunals were closed. Therefore, we can claim that our assumption about censorship enforcement is valid until the first half of the XVIII century.

In Appendix 2.D we provide several robustness checks regarding the Church's ability to enforce censorship over time and space. The sensitivity checks results, summarized in Table 2.4, indicate that our assumptions are not crucial for our baseline results.

2.4 Quantitative Results

2.4.1 Identification Strategy

In this section, we estimate the parameters of the model of knowledge diffusion under the optimizing Church's behavior described in section 2.3, using the data and stylized facts described in section 2.2. We follow a three-step estimation strategy. The first step is to set one parameter following the literature. The second step is to estimate 6 parameters using a minimum distance estimation procedure, under the assumption that censorship kicks in mid 16^{th} century as in the data. The last step is to set one last parameter such that it rationalizes the timing of the introduction of censorship.

[21] In the south and the highlands the Inquisition was carried on by the Spanish, except for Naples, where the archbishop was responsible for censorship (Prosperi et al., 2010).

Before going into the estimation details, we specify the relationship between model periods and their empirical counterpart. We consider 5 model periods that correspond to 1400-1470, 1470-1540, 1540-1610, 1610-1680, and 1680-1750. We made this choice following four criteria. First, we want each period to correspond to an equal number of years. Second, we want to stop in 1750 because the Church might have lost the capacity to censor after this date. Third, we want a year close to 1544 (first edition of the Index) to be the threshold between two consecutive model periods. In this way, we can claim that censorship started in the second of these two periods. Finally, we don't want each period to be too short. If this was the case, the number of authors per period would be small, causing the moments' standard errors to be large.

Preset Parameter. We set the discount factor δ to 0.06, which corresponds to a quarterly discount factor of 0.99: $0.06 \approx 0.99^{280}$. This parameter's role is minimal: conditionally on censorship starting on $t = 3$ (which depends on the fixed cost of censorship ψ), it does not affect dynamics.

Minimum Distance Estimation. We estimate the array of six parameters $\vartheta = [k_1^C, k_1^R, \theta, \overline{\beta}, (1+\nu)\mu, p]$ using a minimum distance estimation procedure.[22] The parameters are identified by minimizing the distance between 14 empirical and simulated moments. These moments are the median and the 75^{th} percentile of the distribution of quality of all authors and the share of censored ideas.[23] We consider these variables for each period $t = \{2, ..5\}$, and we also included the overall knowledge quality in 1.[24]

The identification comes from the initial share of censored authors, initial authors' quality (which identifies parameters k_1^C, k_1^R), and their dynamics. Specifically, $(1+\nu)\mu$ is identified by the growth rate of overall quality. Parameter $\overline{\beta}$ is intuitively identified by the share of censored authors. Parameter p is identified by the dynamics of the shared of censored authors and scholars quality. Parameter θ governs the shape of the Frechet distribution of knowledge quality and is identified by the 75^{th} percentile of the quality distribution.

[22] These are actually 7 parameters, but we cannot distinctly identify the number books read by authors, μ, and the parameter governing the relative productivity of new ideas, ν. They both matter only for growth in quality, which is governed by $(1+\nu)\mu$.

[23] We target the median instead of the mean because it is less sensitive to outliers.

[24] We do not consider the share of censored authors in $t = 1$ as in the model censorship in t affects books written in $t-1$, and censorship starts in $t = 3$.

The objective function $\Omega(\vartheta)$ to minimize is given by

$$\Omega(\vartheta) = (\mathbf{m} - \mathbf{m}_\vartheta)'\mathbf{W}(\mathbf{m} - \mathbf{m}_\vartheta), \tag{2.13}$$

where ϑ is a vector of parameters, $\mathbf{m}$ is the vector of data moments, and $\mathbf{m}_\vartheta$ is the vector of moments obtained simulating the model with parameters ϑ. $\mathbf{W}$ is a diagonal matrix with $1/\mathbf{m}^2$ as elements. The objective function is minimized using the genetic algorithm package in R developed by Scrucca et al. (2013), which allows for global optimization. We computed bootstrapped standard errors of the parameters by drawing 500 random samples with replacement from the original data. For each bootstrap sample, we computed the 14 moments and estimate the corresponding parameters. We then use these boot-strapped estimates to compute the standard errors. The model's simulation is straightforward since there is no uncertainty, and the parameters define both the initial conditions and govern model dynamics. Note that we run simulations assuming that censorship starts in $t = 3$. The timing of censorship depends on the fixed cost of censorship ψ, whose estimation is discussed below.

Parameter set a posteriori. We are left with parameter ψ, namely the fixed cost to set up the censorship apparatus. We set it to rationalize that censorship starts in $t = 3$, given the value of all the other parameters. This parameter is *set identified*: there is a range of values that can rationalize the timing of censorship. The bounds of ψ, namely ψ_L and ψ_R, are set as follows. The lower bound ψ_L is the limit value of ψ for which starting censorship in $t = 3$ gives a larger utility for the Church than starting it in $t = 2$. The higher bound ψ_R is the limit value of ψ for which starting censorship in $t = 3$ gives a larger utility for the Church than waiting and starting it in $t = 4$.[25] Note that we set ψ assuming a linear time utility function $u(1-m)$. If we chose a different shape that respects the assumptions about $u()$, the value of ψ would have changed, but the timing of censorship and the dynamics would have stayed the same. Note that in Table 2.3 we report a scaled value of the fixed cost, defined as $\hat{\psi} = \psi/[V^C(1/(2-\overline{\beta})) - V^N(1/(2-\overline{\beta}))]$.

[25]Starting censorship in previous periods (2,1,0,-1..) would have given the Church a lower utility than waiting for $t = 3$.

2.4.2 Estimation Results

We list the identified parameters and their standard errors in table 2.3. The estimation delivers $k_1^R > k_1^C$: this implies that the quality of censored authors is larger than non-censored authors, which is consistent with data even if the relative quality by sector is not among the targeted moments. The productivity of books θ equals 0.36: this is slightly lower than the value (0.5) used by Lucas (2009). Our estimate is lower because the dispersion in log publications is lower than the one in earnings observed in modern U.S. data, which is the target of Lucas (2009).[26] The relative price of revolutionary books p equals 0.49. This assures that the initial share of revolutionary authors is not too large, even if they have a much larger quality than compliant scholars. For example, if p was equal to 1, the share of revolutionary authors would converge to 1 very fast: as a result, the share of censored authors would converge to $\overline{\beta}$ and stay constant, unlike in the data. The combinations of parameters $(1+\nu)\mu$ assure that knowledge quality would have kept growing if censorship was never introduced. The most interesting parameter is the rate of censorship $\overline{\beta}$ that the Church imposes, which equals 18%.

Table 2.3: Identification of Parameters

Calibrated Parameters		Value	Standard Errors	Target
Discount Factor	δ	0.06	-	RBC literature
Fixed Cost of Censorship	$\hat{\psi}$	(1.029 - 1.031)	-	Index set-up
Estimated Parameters		**Value**	**Standard Errors**	**Target**
Compliant knowledge in 1	k_1^C	13.7	0.93	$\Omega(\vartheta)$
Rev. knowledge in 1	k_1^R	106.7	9.46	$\Omega(\vartheta)$
Productivity of books	θ	0.36	0.016	$\Omega(\vartheta)$
Max Censorship	$\overline{\beta}$	0.18	0.018	$\Omega(\vartheta)$
Knowledge Growth	$(1+\nu)\mu$	2	0.07	$\Omega(\vartheta)$
Price of rev. books	p	0.49	0.017	$\Omega(\vartheta)$

The model fit is reported in figure 2.5, panels a and b. The simulated variables rarely lie outside the 95% confidence interval of the data moments.[27] An exception is the 75^{th} percentile of the overall knowledge quality. This reflects that the underlying empirical distribution does not follow exactly a Fréchet distribution like in the model.

As a test of the theory, we compare our results to empirical observations that were not used

[26] The gini index of log publications is 0.34.

[27] The confidence intervals are computed drawing 500 random samples with replacement and then using the 2.5^{th} and 97.5^{th} percentile from the distribution of the variable of interest.

to identify the parameters. Looking at the dynamics of censored and non-censored authors (figure 2.5, panels c and d) is particularly interesting as it allows us to test whether printers choose their sector according to its (relative) quality. This mechanism is summarized by Equation 2.11: the share of revolutionary authors can assume one and only one variable given the ratio of the quality in the two sectors. This ratio can be proxied by the ratio of censored to non-censored authors' quality, which we can measure in the data. Since the model fit well the dynamics of censored and non-censored authors, we can claim that Equation 2.11 is likely to hold in the data too. The model also predicts that the share of revolutionary ideas was increasing before $t = 3$. This is consistent with the fact that the share of censored books was larger in the period 1470-1540 than 1400-1470. Moreover, the average difference in quality between censored and non-censored authors decreases over time. It is 3.83 in 1470-1540, and it drops to 1.57 in 1680-1750.

2.4.3 The Role of Censorship for Knowledge Formation

What is the role of the Catholic Church for the demise in knowledge production in early modern Italy? How much of this effect is driven by selection into the revolutionary/compliant sectors? In this section we answer these questions by comparing model simulations with and without censorship. This is done by using the parameters identified in section 2.4, with the exception of the rate of censorship $\overline{\beta}$, which is set to 0 in the no-censorship scenario. Figure 2.6 illustrates the outcomes of the experiments.

Without censorship, the share or revolutionary authors m_t would have kept increasing. It would have reached 55% in $t = 5$, instead of decreasing to 25% in $t = 5$. This fact demonstrates the effectiveness of censorship, which can change the dynamics of revolutionary ideas drastically. Moreover, censorship has the unintended effect of reducing the overall quality of scholars, which would have been 28% lower under the baseline than in the $\overline{\beta} = 0$ scenario. In table 2.4 we present a summary of the results for our baseline estimation and from a series of sensibility checks using slightly different versions of the model (e.g., accounting for self-censorship and imperfect enforcement of censorship), different samples, and a different measure of quality. The results are robust to all the sensitivity checks. Appendix 2.D contains further information about each sensibility check.

The loss in the overall quality is both driven by a reduction of the stock of knowledge *within each sector* and by self-selection *across sectors*. This results come from the following

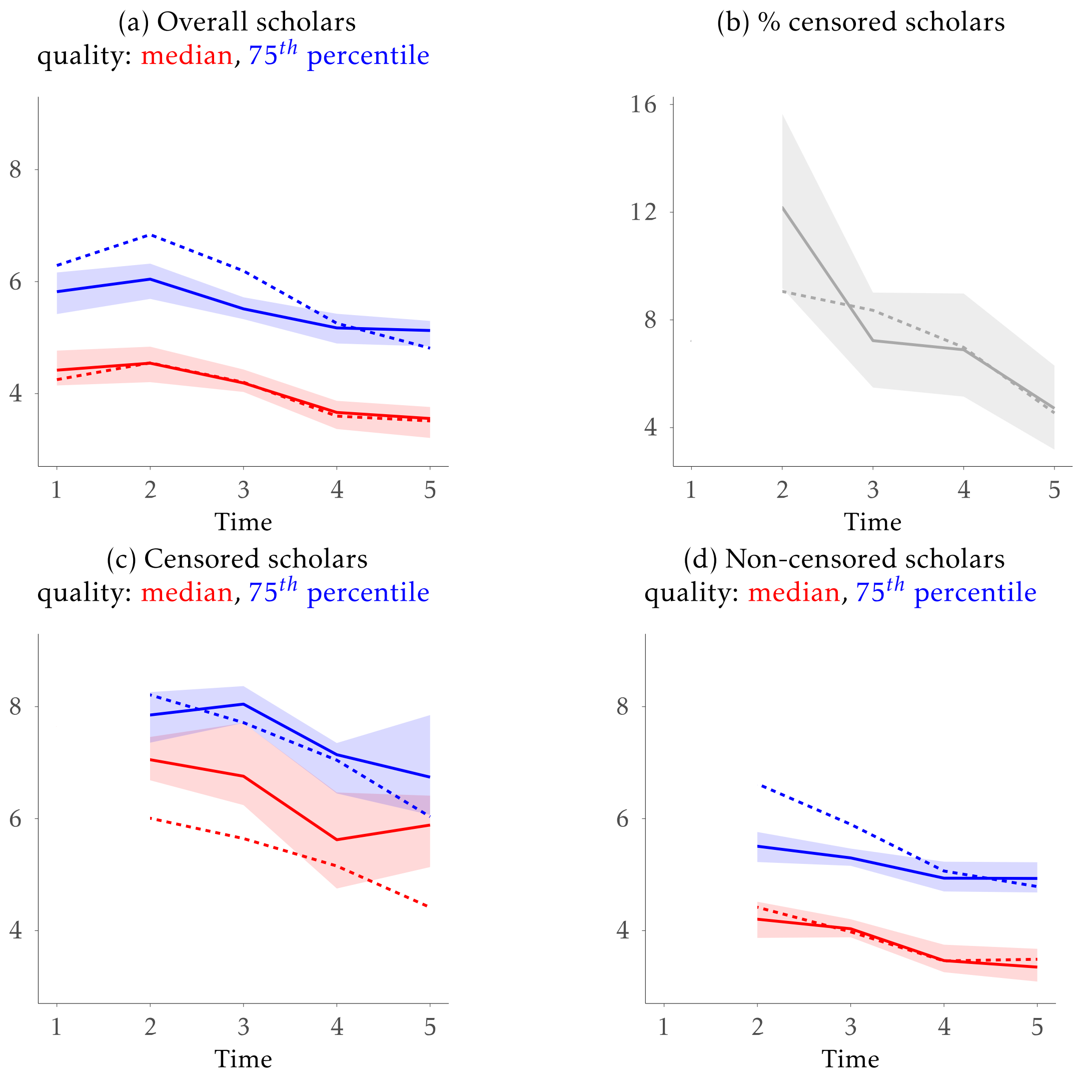


Figure 2.5: Model fit (panels a,b), over-identification checks (panels c,d). Data (solid) and simulations (dashed).

NOTES. periods: 1:1400-70, 2:1470-1540, 3:1540-1610, 4:1610-80, 5:1680-1750.

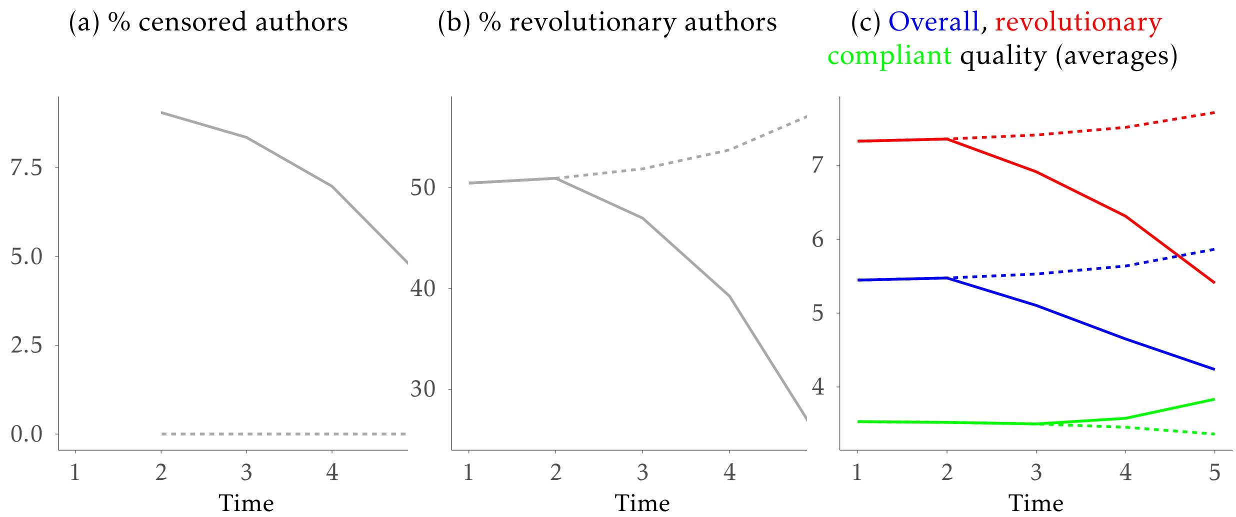


Figure 2.6: Baseline simulations (solid), simulations without cenosorship (dashed)

NOTES. periods: 1:1400-70, 2:1470-1540, 3:1540-1610, 4:1610-80, 5:1680-1750.

Table 2.4: Sensibility checks

	Impact of censorship on: scholars' quality	Impact of censorship on: % heretic scholars	Rate of censorship	% heretic scholars
Symbol	$(q_5 - \hat{q}_5)/\hat{q}_5$	$(m_5 - \hat{m}_5)/\hat{m}_5$	$\overline{\beta}$	m_5
Benchmark	-28%	-56%	18%	26%
Only Italian scholars	-29%	-54%	17%	27%
Only southern Italian scholars	-34%	-53%	19%	32%
Only northern Italian scholars	-23%	-50%	15%	26%
Only $t \leq 4$	-16%	-60%	19%	17%
No weak links to institution	-28%	-54%	17%	26%
All Publications	-29%	-56%	18%	25%
Lenght wikipedia page	-34%	-56%	18%	25%
Imperfect censorship	-24%	-56%	18%	23%
Self censorship	-26%	-51%	16%	27%
Universities only	-28%	-56%	18%	26%

Notes: variables denoted by the hat relate to simulations under a no-censorship scenario, while all the other variables relate to simulations with censorship. Subscript 5 correspond to the period 1680 - 1750. Appendix 2.D contains further information about each sensibility check.

decomposition, where all the variables with a hat relate to the experiment where $\overline{\beta} = 0$:

$$\begin{aligned}\underbrace{q_5 - \hat{q}_5}_{=\text{-1.62 (100\%)}} \quad &= \underbrace{\hat{m}_5[q_5^R - \hat{q}_5^R] + (1 - \hat{m}_5)[q_5^C - \hat{q}_5^C]}_{=\text{-1.39 (69\%); }(a)} + \\ &\underbrace{[m_5 - \hat{m}_5]\hat{q}_5^R + [(1 - m_5) - (1 - \hat{m}_5)]\hat{q}_t^C}_{=\text{-1.12 (85\%); }(b)} \\ &+ \underbrace{(m_5 - \hat{m}_5)[(q_5^R - q_5^C) - (\hat{q}_5^R - \hat{q}_5^C)]}_{=\text{0.88 (-54\%); }(c)}.\end{aligned} \tag{2.14}$$

The effect of censorship due to changes in quality within sectors is captured by (a) in Equation 2.14 and accounts for 69% of the overall drop. The self-selection effect (b) accounts for 85% of the overall drop. This shows that censorship is important as it pushes printers to select compliant knowledge, which has a lower quality. Finally, (c) captures the interaction between effects (a) and (b) and accounts for -54% of the total effect.

To sum up, the effect of censorship on knowledge accumulation is not entirely due to the decrease in quality within sectors. The drop in the revolutionary sector is partially compensated by the increased quality within the compliant sector. Half of the effect of censorship on knowledge growth is due to its ability to make compliant ideas relatively more available. Not only compliant ideas have a lower quality than revolutionary ones, but they would have displayed no growth in quality if there was no censorship.

2.5 Conclusion

Censorship has a direct effect on knowledge accumulation by making censored material less available to scholars. It also discourages writers from engaging in non-compliant work, and hence modifies the allocation of talents across different types of activities. In this paper, we developed a new method that considers these two channels. Then, we applied it to the Catholic Church's censorship from the Counter-Reformation until the age of Enlightenment. We investigated whether censorship was responsible for the demise of Italian science and evaluated the relative importance of the direct channel vs. the activity choice channel.

The analysis proceeded in three steps. First, we collected data on members of universities and academies, identifying the scholars whose books were either free to be printed and sold, or put in the Index Librorum Prohibitorum, i.e. censored. Second, we built a theoretical model of knowledge accumulation through book production and censorship, distinguishing non-compliant knowledge (susceptible of being censored) from compliant knowledge. Third, we estimated the structural parameters of the model using facts collected from the dataset. We used the quantitative model to answer our questions by simulating a counterfactual path of knowledge dynamics characterized by the absence of censorship.

We concluded that censorship reduced by 28% the average log publication per scholar in Italy from 1470-1550 to 1680-1750. Renaissance Italy has been regarded as the cradle of culture and science. Yet, Italy found itself in a scientific backwater during the seventeenth and eighteenth centuries, being overtaken by non-Catholic countries such as Great Britain and The Netherlands. The sizeable effect that we estimated allows claiming that Church's censorship is one of the main drivers of Italy's decline.

Half of this drop stems from the induced reallocation of talents towards compliant activities, while the other half arises from the direct effect of censorship on book availability. This result stresses the importance of selection effects when analyzing the impact of censorship on output. The top scholars at the time of the Counter-Reformation were all censored (Bruno, Galilei, Copernicus), and their successors might have decided to become mediocre poets or theologians rather than dangerous mathematicians.

Finally, one may wonder whether the Church's censorship also had a role in the *economic* decline of Italy. This possibility does not seem implausible, given that recent research highlighted the role of upper-tail human capital production for pre-industrial Europe's take-off (Squicciarini and Voigtländer, 2015; Cantoni and Yuchtman, 2014; Mokyr, 2011, 2016). Our analysis sets the stage for future research on this topic by directly linking the Church's censorship to upper-tail human capital production.

Appendix

2.A Additional data

2.A.1 How representative are university professors and academicians?

The paper is based on publications by university professors and members of academies. One may wonder how representative are those publications of the total production of knowledge in early modern times. To answer that question, one needs to define a new universe of persons from which we can extract the sample of university professors and compute their share. Looking at scientific domains, let us consider the scientists who have given their name to a crater on the moon. Those names were given by the Commission on Lunar Nomenclature of the International Astronomical Union from 1935 onward (Richardson, 1945). Among these persons there are 54 Italians born before 1770. Figure 2.A.1 represents their occupation breakdown. A large majority of them were either university professor or member of an academy or both. This comforts the idea that our sample of scholars is a good representation of people working in sciences.

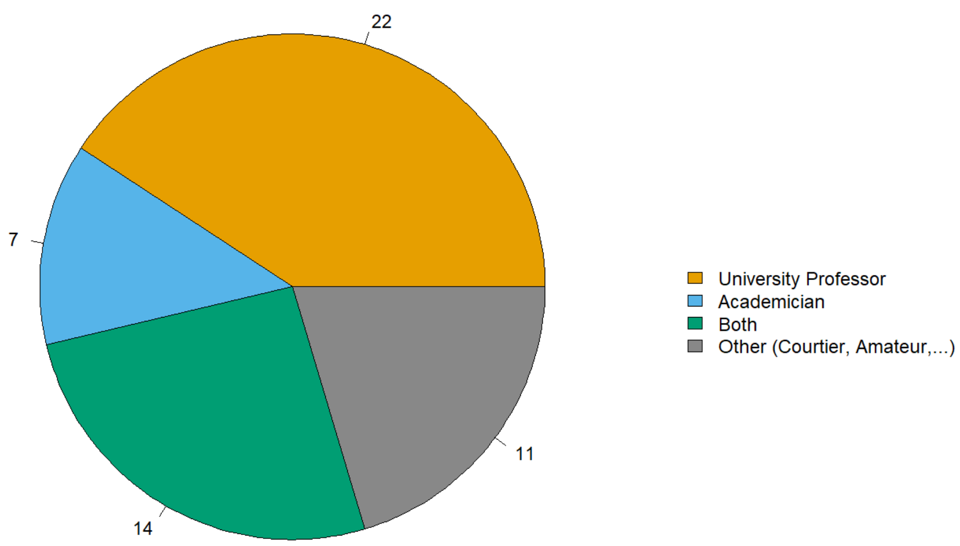


Figure 2.A.1: Occupations of Italians having given their name to a crater on the moon

2.A.2 How much of the Italian University/Academy population is covered?

Another question is how much of the Italian University/Academy population is covered. For universities, we can provide a precise answer. We believe we have a comprehensive coverage for the following universities: Ubologna-1088, Umacerata-1540, Umodena-1175, Upadua-1222, Upavia-1361, Upisa-1343, UromaGregoriana-1556. Thanks to very detailed secondary sources, we almost have all professors having taught there.

We have a broad coverage for the following universities StudFlorence-1321, Ucagliari-1606, Ucatania-1444, Umantua-1625, Umessina-1548, Unapoli-1224, Upalermo-1578, Uperugia-1308, Uroma-1303, Usalerno-1231, Usassari-1617. Thanks to detailed secondary sources, we have a large number of the professors having taught there, and we probably have all those who published something, which is the relevant dimension for this paper.

For the following list, we have only a partial coverage. Many of those universities are quite small, or specialized, or detached from bigger universities (Milano & Venice). We will be able to complete Ferrara and Parma soon. Ualtamura-1748, Uancona-1562, Ucamerino-1727, Ufermo-1585, Uferrara-1391, Ugenoa-1773, Ulucca-1369, Umilano-1556, Umondovi-1560, Uparma-1412, Upau-1722, Usiena-1246, Utorino-1404, Uurbino-1671, Uvenice-1470, Uvicenza-1204.

For academies, assessing our coverage is more complicated, as the number of academies is potentially very large. Each city had one or more small academies, sometimes very temporarily, gathering the curious minds of the moment. As we explained in the text, our more important source comes from the data compiled by the British library based on all the books in their possession related in one way or in another to an Italian academy. To this list, we added important academies for which there is a complete coverage based on a biographical dictionary of their members: the Crusca, the Ricovrati, and the Gelati.

2.A.3 How is the distribution of the scholars'fields changing over time?

Europe overtook Italy in terms of scholars quality. In principle, this could be driven by the mere fact that a field with low average publications became relatively more common in Italy than in Europe. To answer this question, in table 2.A.1 we show the dynamics of scholars quality in Italy and Europe by field.[28] We observe that Europe overtakes Italy in each field at the time censorship was introduced. Moreover, the share of publications in science and medicine rises much faster in Europe than in Italy. This analysis is consistent with the effect of censorship in the model, according to which the Index i) decreases quality within sectors/fields; ii) affect the occupational choice of scholars.

Table 2.A.1: Distribution & publications by period and field

	Distribution (%) of the scholars'fields for each period					Average publications per person				
Period	1	2	3	4	5	1	2	3	4	5
						Italy				
Theology	7	7	9	8	11	329	254	406	187	215
Law	41	27	20	14	16	206	220	167	114	70
Humanities	27	40	44	47	35	641	608	452	293	207
Medicine	13	13	14	11	11	119	360	248	143	135
Sciences	11	10	8	13	16	72	456	724	238	88
Others	1	3	5	8	11	6	3539	79	40	180
						Europe (excluding Italy)				
Theology	34	26	26	27	22	173	595	945	350	278
Law	25	18	19	16	13	173	123	252	261	379
Humanities	30	41	35	33	32	313	346	470	442	496
Medicine	6	8	12	11	14	62	196	322	208	291
Sciences	3	5	8	11	14	62	254	417	313	384
Others	2	1	1	2	6	22	48	1012	364	736

Note: periods: 1:1400-70, 2:1470-1540, 3:1540-1610, 4:1610-80, 5:1680-1750.

- **Theology**: Theology, scriptures
- **Law**: Canonic law, Roman law, French law, juriconsulte
- **Humanities**: History, Literature, Philosophy, Ethics, Rethoric, Greek, Poetry
- **Medicine**: Medicine, Anatomy, Surgery, Veterinary, Pharmacy, Botany
- **Sciences**: Mathematics, Logic, Physics, Chemistry, Biology, Astronomy, Earth Science, Geography
- **Others**: Applied Sciences (Engineering, Architecture, Agronomy), Social Sciences (Economics, Political Science)

[28]In case the scholar is associated with more than one field, we expand the observation according to the number of her/his fields. Details about each disciplie can be found below table 2.A.1.

2.A.4 Additional Figures and Tables

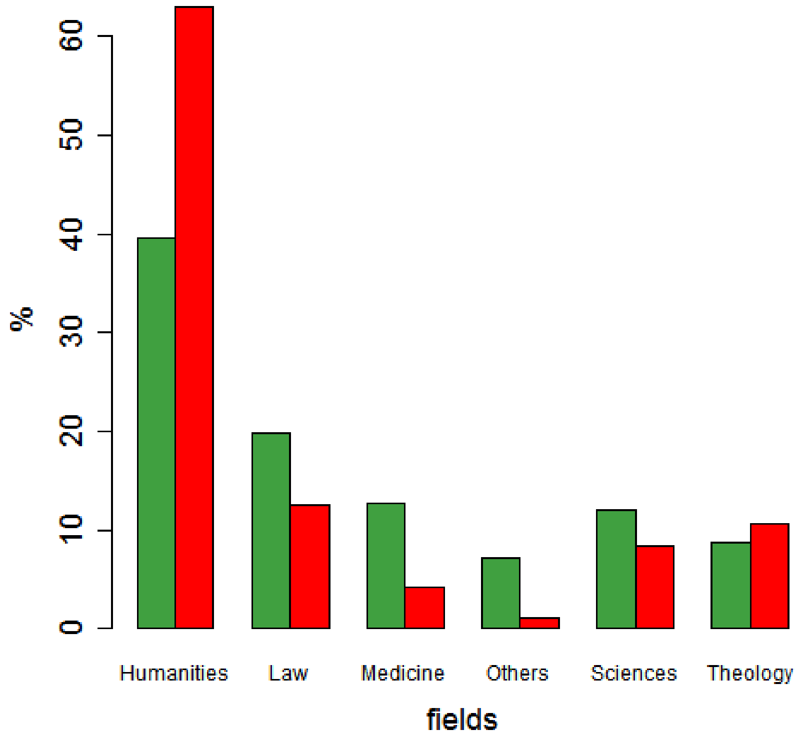


Figure 2.A.2: Distribution of the fields of scholars. Red: censored. Green: non-censored.

Table 2.A.2: Total number of scholars & publications by period

	Number of published scholars					Mean publications per person				
Period	1	2	3	4	5	1	2	3	4	5
Europe	107	270	599	577	698	536	936	1048	736	915
France	15	27	116	124	169	946	599	894	947	1231
Germany & Austria	12	67	110	37	152	602	913	1933	965	862
Great Britain & Ireland	5	17	43	89	184	14	1845	616	960	1055
Denmark & Sweden		2	9	16	31		1084	553	495	829
Spain & Portugal	6	25	58	25	11	755	371	454	514	571
Italy	60	105	174	215	85	485	875	773	462	510
Ubologna-1088	4	16	13	14	5	139	360	546	140	150
Unapoli-1224	1	5	2	1	2	1840	116	105	96	2142
Upadua-1222	9	16	24	9	10	140	683	622	716	141
Upavia-1361	6	10	6			410	520	701		
Uroma-1303	17	16	13	9	4	888	1032	1535	428	234
Upisa-1343		6	6	11	4		232	313	285	152
UromaGregoriana-1556			7	6				1211	557	
Umacerata-1540			1		1			34		218
AcadRicovrati-1599			2	15	25			171	518	706
AcadCrusca-1583			4	23	10			397	331	1052
AcadGelati-1588			1	11	3			887	328	150

Note: periods: 1:1400-70, 2:1470-1540, 3:1540-1610, 4:1610-80, 5:1680-1750

2.B Bibliographies

John Barclay (Pont–Mousson 1582 - Roma 1621, censored in 1608) was born in Pont-à-Mousson, Lorraine, France, from a Scottish-born father. In 1605 John Barclay presented the first part of his *Euphormionis Lusinini Satyricon*. This humanist novel is a very original piece of work (Correard, 2017), including a satirical description of the Jesuit schools (in which he was raised). This book was put on the Index on 13 December 1608 (De Bujanda and Richter, 2002). Upon invitation by the Pope himself, he went to Rome in 1616 and resided there until he died in 1621. Moving to Rome was a way to testify that he was a good Catholic. John Barclay was a member of several Italian academies, including the Accademia degli Umoristi and the Accademia dei Lincei.

Giordano Bruno (Nola 1548 - Roma 1600, censored in 1600) was an Italian friar, member of the Dominicans. His contributions span from philosophy to mathematics and cosmology. He is best known for being persecuted by the Catholic Church and was later regarded as a martyr for science. The Inquisition found him guilty of heresy for several of his views, among which his positions on cosmology: he theorized an infinite universe and a plurality of words. All his works entered the Index of forbidden books, and he was burned at stake in Rome's square Campo de' Fiori.

Bernardino Ciaffoni (Porto Sant'Elpidio 1615/1620 - Marches 1684, censored in 1701) was a theologian and belonged to the order of the Franciscans. He also used to be a rector of the well-known college San Bonaventura, located in Rome. His *Apologia*, published posthumously, defends the rigorist doctrine and fights the probabilism, supported by Jesuits. This piece of work was introduced in the Index because of its insulting claims against Jesuits.

Nicolaus Copernicus (Thorn 1474 - Frauenburg 1543, censored in 1616) was a Prussian mathematician and astronomer. In his book *De revolutionibus orbium coelestium*, he theorized the cosmos as having the Sun at the center of the solar system, where the Earth rotated around it. This theory is in deep contrast with the Ptolemaic model, where the Earth is stationary at the center of the universe. Several other scientists, among which Galilei, contributed to his theory by bringing evidence in its support. While his theories were welcomed positively by the Church at first, his *De revolutionibus* was censored in 1616, after that the Church's conservative revolution.

Achille Gagliardi (Padova 1537 – Modena 1607, censored in 1703) was a Jesuit theologian and spiritual writer. He taught philosophy at the Roman College, then theology in Padua and Milan. He was a collaborator of the Archbishop of Milan Carlo Borromeo, who requested him to write a handbook of religion, the popular *Catechismo della fede cattolica*. His *Breve compendio* was censored because of his thoughts about the annihilation of the will during mystical states. These ideas are not compatible with free will, which is a cornerstone of catholic theology.

Galileo Galilei (Pisa 1564 - Arcetri 1642, censored in 1634) was an Italian astronomer and physicist. Also Professor in Padova and member of the prestigious Accademia dei Lincei, arguably he was the most notable and influential scientist of his times. He is also known as the father of modern science because of his work on the scientific method. His books were censored because of its support to atomism, heliocentrism, and Copernicanism. The Inquisition condemned him, and he was forced to abjure his thesis and spent the last part of his life under house arrest.

Serry Jacobus Hyacinthus (Toulon 1659 – Padua 1738, censored in 1722) was a theologian and belonged to the order of the Dominicans. Also consultor of the Congregation of the Index, he taught theology at the university of Padua from 1698. His *Historiae*, written under the pseudonym Augustinus Leblanc, deals with the Jesuit-Dominican controversy on grace and was prohibited by the Inquisition.

2.C Proofs of Propositions

2.C.1 The Fréchet Cheat Sheet

Since the irrelevance of books of type j is exponentially distributed with scale parameter k_t^j and given Equation (2.1), the distribution of book quality follows a Fréchet distribution with scale parameter ${k^j}^\theta$ and shape parameter $1/\theta$. This allows us to write the average book quality q^j by sector as:

$$E(q_i^j) = \int_0^\infty h_i^{-\theta}(k^j e^{-k^j h_i})dh_i \quad \text{with } j \in \{C, R\} \text{ and } i \in \{1,..,N\},$$

Now we can multiply the RHS by $(k^j)^{1+\theta}/(k^j)^{1+\theta}$ to obtain:

$$E(q_i^j) = (k^j)^{1+\theta} \int_0^\infty (k^j h_i)^{-\theta} (e^{-k^j h_i}) dh_i.$$

Now, using a change of variable $y = k^j h_i$ we have that

$$E(q_i^j) = (k^j)^{1+\theta} \int_0^\infty (y)^{-\theta} (e^{-y})(1/k^j) dy.$$

We can finally show that

$$E(q_i^j) = (k^j)^{\theta} \Gamma(1-\theta) \quad \text{with } j \in \{C, R\} \text{ and } i \in \{1,..,N\},$$

where $\Gamma(x) = \int_0^\infty s^{x-1} e^{-s} ds$ is the Euler gamma function.

2.C.2 Proof of Proposition 1

Using the variable z_t, Equation (2.12) can be rewritten as

$$z_{t+1} = \frac{1-\beta}{\hat{p}} (z_t)^2.$$

This recurrence Equation admits an explicit solution:

$$z_t = \frac{\hat{p}}{1-\beta} \left(\frac{z_0(1-\beta)}{\hat{p}} \right)^{2^t}. \tag{2.15}$$

Equation (2.11) implies that once we know the dynamics of z_t, we also know the dynamics of m_t. Given this change of variable, we use Equation (2.15) to study the limit of z_t and obtain

a) $z_0 < \hat{p}/(1-\beta) \Rightarrow \lim_{t\to\infty} z_t = 0$. Note also that $m_0 < 1/(2-\beta) \Leftrightarrow z_0 < \hat{p}/(1-\beta)$.

b) $z_0 > \hat{p}/(1-\beta) \implies \lim_{t\to\infty} m_t = 1$. Note also that $m_0 < 1/(2-\beta) \Leftrightarrow z_0 < \hat{p}/(1-\beta)$.

c) $z_0 = \hat{p}/(1-\beta) \implies z_t = \hat{p}/(1-\beta) \forall t$. Note $m_t = 1/(2-\beta) \forall t \Leftrightarrow z_t = \hat{p}/(1-\beta) \forall t$

From a) and Equation 2.11, i) follows. From b) and Equation 2.11, ii) follows. From c) and Equation 2.11, iii) follows.

Note that we excluded $m_0 = 1$ from the proposition. In that case, no compliant books are left in the economy and imposing $\beta = 1$ would shut down the whole production of knowledge.

2.C.3 The Dynamics under a Rule of Thumb Church's Behavior

In Section 2.3.4 we described the dynamics under a constant rate of censorship β_t. Here we endogenize the introduction of censorship by assuming that the Church chooses the lowest censorship rate that allows to converge to a world with no revolutionary ideas. This is equivalent to assume that the Church has lexicographic preferences, caring firstly to have $\lim_{t\to\infty} m_t = 0$, and secondly to minimize β_t. Given our assumptions, we can describe the dynamics of the share of revolutionary ideas in Proposition 4.

Proposition 4. *For a given share of revolutionary ideas* $m_t \in [0,1)$*, the Church will choose a level of censorship* β_t *such that* $\beta_t = \max\{2 - 1/m_t + \epsilon, 0\}$*, where* ϵ *is arbitrarily small.*

Proof. Notice that Proposition 1 states that $\lim_{t\to\infty} m_t = 0$ when $m_t < 1/(2-\beta_t)$, from which it trivially follows that $\beta_t = \max\{2 - 1/m_t + \epsilon, 0\}$. □

Note that for any initial $m_0 \in [0,1)$, we will have $\lim_{t\to\infty} m_t = 0$, but the convergence will be slow due to the fact that in any period m_t would be set very close to the unstable steady state $1/(2-\beta_t)$. It is worth noting that Proposition 4 implies that the Church will impose no censorship if $m_t < 1/2$.

2.C.4 Proof of Proposition 2

Note that imposing censorship when $m = 0$ is not convenient:

$$\frac{u(0)}{1-\delta} = V^N(0) > V^C(0) - \psi = \frac{u(0)}{1-\delta} - \psi.$$

Note also that imposing censorship when $m = 1$ is not convenient.

$$\frac{u(1)}{1-\delta} = V^N(1) > V^C(1) - \psi = \frac{u(1)}{1-\delta} - \psi.$$

Note also that $V^M(m)$ and $V^C(m)$ are continuous functions in $m \in [0,1]$: see Norets (2010) for a formal proof of continuity of discrete choice dynamic value functions under a set of assumption that are satisfied in our case.

Then, it follows that there exists $\tilde{m}$ and $\check{m}$, respectively in a neighborhood of 0 and 1, such that for each $m \in [0,\tilde{m}]$ and also for each $m \in [\check{m},1]$, $V^N(m) > V^C(m) - \psi$ holds. According to proposition 1, if censorship is not imposed, $\tilde{m}$ converges to 0, while $\check{m}$ will coverge to 1.

Since censorship do not happen foreach $m \in [0, \tilde{m}]$ and for each $m \in [\check{m}, 1]$, proposition 2 is proved.

2.C.5 Proof of Proposition 3

We take $\overline{\psi}$ such that for some m^* we have $V^C(m^*) - \overline{\psi} > V^N(m^*)$, then for each $\psi < \overline{\psi}$ it holds $V^C(m^*) - \psi > V^N(m^*)$. Now define $\mathcal{D}(m) = V^C(m) - \psi - V^N(m)$: since this function is continuous, for an arbitrarily small ϵ we have that $\mathcal{D}(m^* - \epsilon) > 0$ and $\mathcal{D}(m^* + \epsilon) > 0$. Using again continuity we can claim that $\mathcal{D}(m) > 0$ for each $m \in [m^* - \epsilon, m^* + \epsilon]$, which implies that the Church will immediately impose censorship if m_0 belongs to this set.

2.D Robustness Checks

This section describes the sensitivity checks whose results are reported in table 2.4.

Only Italian scholars. Some scholars might have spent only a period of their time in Italy. Living outside Italy could have allowed them to access forbidden books without consequences. To limit this problem, we estimate the model using a sample of Italian scholars only. Table 2.4 shows that this sensitivity check's results differ only slightly with respect to baseline results.

Only southern/northern Italian scholars. The model used for the baseline estimation assumes that the rate of censorship that the Church can enforce does not depend on scholars' location in Italy. This assumption is problematic if the actual rate of censorship differed drastically across Italian regions. To understand whether this is the case, we estimate the model separately for Italian scholars born in northern and southern Italy. A scholar is defined as northern Italian if he is born in a city whose latitude is larger than 43.8, which corresponds to cities northern than Florence. The results reported in table 2.4 indicate that the effect of censorship on knowledge growth is very similar for northern and southern Italian scholars. The effect is slightly stronger for southern Italians because the rate of censorship there is slightly larger. This result is consistent with the stronger capacity that the Church had in the Papal state.

Only $t \leq 4$. In the baseline model, we assume that the Church could enforce censorship until 1750, the end of period $t = 5$. In this sensitivity check, we re-estimate the model assuming that the Church can enforce censorship until the end of $t = 4$ only, or 1680. In the last period $t = 5$, the Church keeps censoring authors, but anyone can read revolutionary books. The Church's ability to enforce censorship likely decreased over time. It is also likely that her ability to censor did not disappear completely. Hence, we think that this robustness provides a lower bound to the effect of censorship on knowledge growth. Despite the conservative assumption, the results in table 2.4 show that the effect of censorship is still large, even though slightly lower than in the baseline case. This is because once the decline of revolutionary ideas started, its decline becomes unstoppable because of inertia.

No weak links. In our baseline sample, we included scholars who have a weak link to a university or academy. These include foreign and corresponding members to academies. One example is Leonhard Euler at Accademia Ricovrati. While all these scholars decided to do some work with the institution, they might not have been there physically. Scholars with weak links might be less constrained by the Church's censorship: for example, because they lived elsewhere in Europe. Hence, we propose a sensitivity check where we exclude them from the sample and then re-estimate the model. Table 2.4 reports the results, that differ only slightly from the baseline estimation. One reason why excluding weak links have a low effect on the results is that they represent less than 2% of the original sample.

All publications. In the baseline sample, we measure the author's quality by the number of publications written *by* them. It is possible to argue that quality is better measured if also publications *about* the author are included. These capture the impact that these authors had on future generations. Table 2.4 reports the results where quality is measured by considering both publication *by* and *about* the author. The role of knowledge accumulation is very similar to the baseline, which indicates that results are robust to different quality measures.

Lenght Wikipedia pages One problem with our measure of authors' quality is that it may be biased because older works have more editions. To limit this problem, we consider a different measure of author's quality, based on the number of characters of the author's longest Wikipedia page. Table 2.4 shows that our results are robust to this different measure of quality. Note that for building this measure of authors' quality we followed de2020academic by assuming that having no Wikipedia page is similar to having one page with a length of 60 characters.

Imperfect censorship. In the model, we assumed that none could access the knowledge embodied in forbidden books. As historians (Grendler, 1975) documented the smuggling of forbidden books, this assumption is not always met. This sensibility check consists in assuming that the Church was able to enforce censorship only in χ% of total cases. Hence, even if $m_t\overline{\beta}$ authors have been censored, only $m_t\overline{\beta}\chi$ are not available to the next generation. One important question is how to set the value of χ. Our strategy is the following: we set χ the lowest possible, under the constraint that, after the re-estimation of the model given χ, it exists one value of ψ rationalizing the timing of censorship (introduced in $t = 3$).[29] In this way, we give the best shot to this sensibility check to undermine the baseline results. We find $\chi = 0.91$. When χ is too small, the model's estimation implies that the share of revolutionary ideas would be declining already in $t = 1$.[30] When this happens, censorship is set immediately, and the timing of censorship cannot be rationalized. Also, low levels of χ implying $m_1 < m_2$ are at odds with the fact that the share of revolutionary authors is larger in 1470-1540 than in 1400-1470. The results reported in table 2.4 indicate that imperfect censorship has only a minor effect on the baseline results. In particular, the impact of censorship on knowledge growth stays large and negative.

Model with self-censorship In section 2.2.2 we claimed that, in some cases, the persecution by the Church was not confined to censorship, but the author lost his life or was imprisoned. Hence, it is reasonable to think that some authors might have decided to self censor to avoid risking their life. Others might have migrated elsewhere in Europe, where the Church could not reach them.[31] These two mechanisms can decrease the share of revolutionary authors for a reason other than direct censorship. In subsection 2.3.3 we proposed to take these mechanisms into account by discounting the quality of revolutionary books by γ. Here we re-estimate the model enriched by this featured. Parameter γ is mostly identify by $\overline{\beta}m_2$, which is too low in the simulations when the baseline model is used. Parameter γ helps to make the demise of revolutionary ideas faster, thus allowing for an initial larger level of revolutionary ideas. The estimation implies that γ=0.95 and $\overline{\beta}$=0.16, which is very close to the baseline. Then, we asses the role of direct censorship by comparing simulations with the estimated $\overline{\beta}$ and setting $\overline{\beta} = 0$, where γ is always set to its estimated value. If the baseline model was misspecified, the version with self censorship should give a different effect of direct censorship on knowledge growth. This is not the case: table 2.4 shows that

[29]The procedure to set ψ is the one described in section 2.4. In particular, the timing of Censorship by the Church can be identified if ψ_L¡ψ_R.

[30]When censorship is not effective, the decrease in quality can be matched only when this happens naturally, without any intervention by the Church.

[31]de2020academic show that an European academic market existed in early modern times.

the results differ only slightly from the baseline. To understand the joint role of direct and self censorship, we perform a counterfactual simulations where $\overline{\beta} = 0$ and $\gamma = 1$. The joint effect reduces knowledge quality by 41%. Since the effect of direct censorship was 29%, this means that also self-censorship has an effect on knowledge quality, even if including it in the model do not alter the baseline results about the effects of direct censorship.

Universities only In the baseline estimation we consider both University professors and members of academies. In appendix 2.A.2 we show that while the coverage of University professor is very good, we probably miss many member of academies. Hence, we provide a robustness check where we exclude those scholars who were not professors. Table 2.4 shows that the result of the baseline and this alternative estimation are very similar: censorship reduced log publications by 26% in the first case and 24% in the second case.

Chapter 3

(Changing) Marriage and Cohabitation Patterns in the US: do Divorce Laws Matter?[1]

[1]This chapter is coauthored with Egor Kozlov (Northwetstern University, Evanston, Illinois, United States). The authors wish to thank Edoardo Ciscato, David de la Croix, Matthias Doepke, Jesús Fernández-Villaverde, Philipp Kircher, Fabio Mariani, Pauline Morault, Luca Pensieroso, Alessandra Voena, as well as seminar and conference participants at Northwestern University, KU Leuven, UCLouvain, the University of Cergy-Pontoise, and the University of Michigan. Computational resources were provided by the supercomputing facilities of the Université catholique de Louvain (CISM/UCL) and Northwestern University. Blasutto acknowledges financial support from the French speaking community of Belgium (ARC project 15/19-063 on *family transformations* and *mandat d' aspirant* FC 23613).

3.1 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 partnership 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.[2]

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

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

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.[3] 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 cooperates while married and the allocation of resources corresponds to the Pareto-efficient inter-temporal allocation. When just one spouse (cohabitant) can decide to terminate the relationship this causes a lack of commitment,[4] 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.[5]

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

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

[4] 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.

[5] 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).

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 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.[6] 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

[6]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.

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. 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 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.[7] 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

[7] 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.

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.[8]

Another subset of these papers studies the effect of changes in cohabitants' rights on partnership 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 (Fisher, 2012; Goussé and Leturcq, 2018; Chigavazira et al., 2019). 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.[9] 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 the 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 (Fernández et al., 2004; Greenwood et al., 2005; Albanesi and Olivetti, 2016; Greenwood et al., 2016), the changes in household formation and dissolution (Greenwood et al., 2016; Ciscato, 2019), the rise in positive assortative mating (Fernandez et al., 2005; Greenwood et al., 2016; Ciscato, 2019) and the increase in the age at marriage

[8] 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.

[9] 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.

(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.

3.2 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.[10] Both spouses' agreement was needed to obtain a divorce for mundane reasons (i.e., without misconduct 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 3.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. 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

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

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: *i*) cohabitation is usually very short and not committed; *ii*) cohabitants are younger and poorer than marrieds; *iii*) 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.[11] The lower costs of a breakup are

[11] 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. Under this legal framework, 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

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 3.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.

3.3 Data and Empirical Evidence

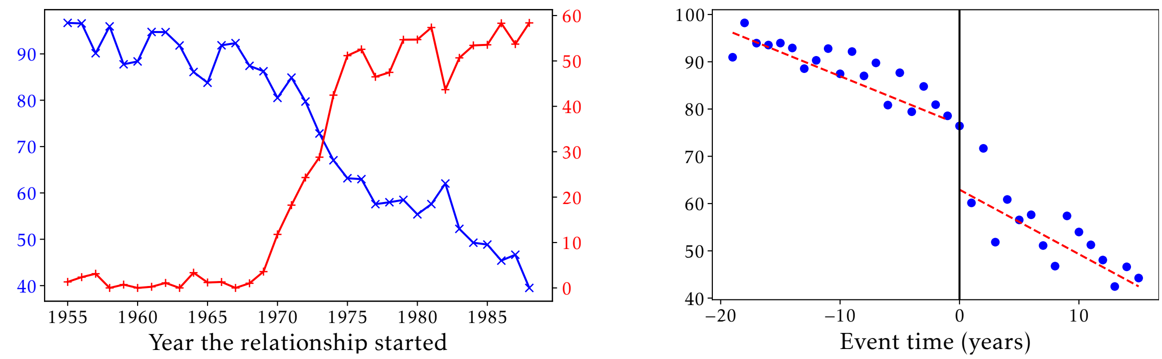
Is the introduction of unilateral divorce related to the rise of cohabitation? Figure 3.1 suggests that the link between these two events merits investigation. The left panel shows a negative correlation between the share of couples that decide to marry as opposed to cohabit, and the share of couples 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.

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 3.1:
Newly formed relationships (either married or cohabiting) and unilateral divorce (U.D.)

(a) % new couples choosing marriage instead of cohabitation, % relationships started 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.

3.3.1 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.[12] While the NSFH I is the first of three longitudinal waves, NSFG is made of several repeated cross sectional samples.[13] 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

[12]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.

[13]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.

NSFG.[14] Since we also know 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*, which 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.[15] One first relationship is defined observing the first time (if ever) a certain person started cohabiting or married. This observation is associated with the date at which the relationship starts, the characteristics of the respondent member of the formed couple, and with a *type*, which can be 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 or two. We did not consider third or higher order relationships 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 3.1 we report the descriptive statistics of this sample.

[14]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.

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

Table 3.1:
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 3.2 we report the descriptive statistics of this sample.

Table 3.2:
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

3.3.2 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.[16] 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.[17] 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 (3.1), where i is the newly formed couples, t is the calendar time, and s is the state:

$$\text{married}_{i,t,s} = \beta_0 + \beta_1 \cdot \text{Unilateral}_{t,s} + \gamma' \mathbf{X}_i + \delta_s + \nu_t + \epsilon_{i,t,s}. \tag{3.1}$$

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 3.3 for different samples. Column (1) reports the results for the full sample described in section 3.3.1, 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. The results reported in table 3.3 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 3.F.1), to the use of a multinomial logit that takes into account the triple choice between staying single, cohabiting and marrying (table 3.F.5), and to dropping California from the sample (table 3.F.3). Moreover, table 3.F.6 shows that the shift towards cohabitation holds both for households

[16] See table 3.A.1 for the timing of adoption of unilateral divorce.
[17] Among the others, see Wolfers (2006), Stevenson (2008), Voena (2015), Reynoso (2020) and Ciacci (2017).

Table 3.3:
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%.

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 3.F.1, show that the effect stays significant and is even slightly larger.

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 (3.2)

$$\begin{aligned} \text{married}_{i,t,s} = \beta_0 + &\beta_1 \cdot \text{Unilateral}_{t,s} \cdot \text{No Title Based}_{t,s} \\ &+ \beta_2 \cdot \text{Unilateral}_{t,s} \cdot \text{Title Based}_{t,s} + \\ &\quad \beta_3 \cdot \text{Title Based}_{t,s} + \gamma' \mathbf{X_i} + \delta_s + \nu_t + \epsilon_{i,t,s}, \end{aligned} \tag{3.2}$$

whose indexes and controls are the same as in equation (3.1), 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 3.4 we report the results of the estimation of equation 3.2. Similarly to table 3.3, column (1) reports the results for the full sample described in section 3.3.1, 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 3.4:
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 couple chooses marriage over cohabitation in non-title-based states is significant with a magnitude between −7% and −9% depending on specification. At the same time, 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 (3.F.2), to the use of a multinomial logit that takes into account the triple choice between staying single, cohabiting and marrying (3.F.5), and to dropping California from the sample (3.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:

$$\begin{aligned}
Y^{\text{Marry}}_{i,s,t,d} &= \beta^{\text{Marry}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Marry}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon^{\text{Marry}}_{i,s,t,d}, \\
Y^{\text{Cohabit}}_{i,s,t,d} &= \beta^{\text{Cohabit}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Cohabit}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon^{\text{Cohabit}}_{i,s,t,d}, \\
Y^{\text{Breakup}}_{i,s,t,d} &= \beta^{\text{Breakup}} \cdot \text{Unilateral}_{s,t} + \gamma^{\text{Breakup}'}\mathbf{X_i} + \alpha_d + \delta_s + \nu_t + \epsilon^{\text{Breakup}}_{i,s,t,d},
\end{aligned} \tag{3.3}$$

where

$$\epsilon^{\text{Marry}}_{i,s,t,d}\epsilon^{\text{Cohabit}}_{i,s,t,d}\epsilon^{\text{Breakup}}_{i,s,t,d} \sim \mathcal{N}(\mathbf{0}, \Sigma), \tag{3.4}$$

and

$$Y_{i,s,t,d} = \begin{cases} \text{Marry} & \text{if } \; Y^{\text{Marry}}_{i,s,t,d} > Y^{\text{Cohabit}}_{i,s,t,d} \text{ and } \; Y^{\text{Marry}}_{i,s,t,d} > Y^{\text{Breakup}}_{i,s,t,d} \\ \text{Cohabit} & \text{if } \; Y^{\text{Cohabit}}_{i,s,t,d} > Y^{\text{Marry}}_{i,s,t,d} \text{ and } \; Y^{\text{Cohabit}}_{i,s,t,d} > Y^{\text{Breakup}}_{i,s,t,d} \\ \text{Breakup} & \text{otherwise.} \end{cases} \tag{3.5}$$

Note that $\text{Unilateral}_{s,t}$ does not have the subscript d since this variable takes value 1 if cohabitation started under a unilateral divorce regime and 0 otherwise. 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 3.5 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 3.5 reports the parameters of the multinomial probit and the average relative risk that the event of interest (marriage of breakup) is realized.[18] The results show that unilateral divorce caused an increase in cohabitation duration, 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,[19] unilateral divorce drives down the risk of breakup because of a selection effect. In appendix 3.G we provide a robustness check for the duration analysis of cohabitation. More specifically, we estimate two linear models: in the first one the dependent variable is a dummy taking value 1 if in that month the couple married (table 3.G.1), while in the second one the dependent variable is a dummy taking value 1 if that month the couple married (table 3.G.2). The samples and the controls are the same used for the multinomial probit reported in table . This additional analysis confirms that cohabiting couples formed after the introduction of unilateral divorce last longer because of a lower risk of breaking up and marrying.

[18]These risks are computed relatively to the probability to continue cohabiting.
[19]This seems plausible because the risk of divorce is much lower than the risk of breakup.

Table 3.5:
Multinomial Probit. Observation: person-month of cohabitation

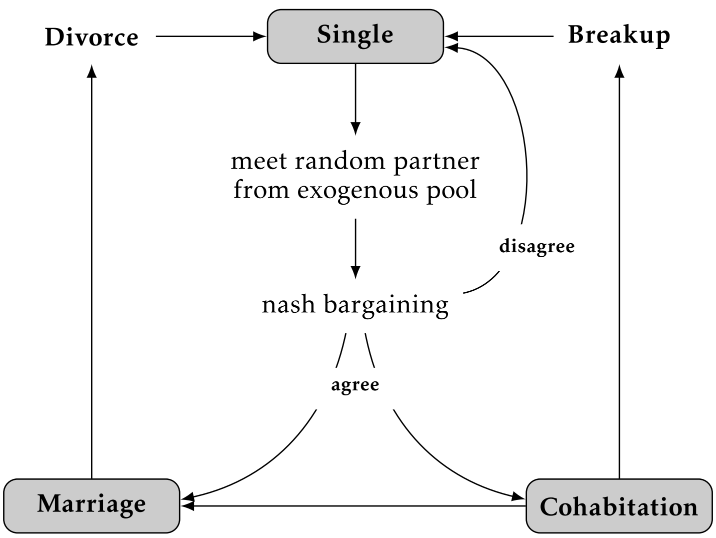
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
	Risk of Marriage relative to Cohabitation			
Unilateral Divorce	-0.24^{***}	-0.25^{***}	-0.28^{***}	-0.28^{***}
	(0.06)	(0.08)	(0.09)	(0.09)
Average Relative Risk	0.64	0.63	0.59	0.6
	Risk of Breakup relative to Cohabitation			
Unilateral Divorce	-0.19^{***}	-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%.

3.4 Theory

To identify the channels through which unilateral divorce impacts 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 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 3.2



3.4.1 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 $\sigma_{\psi,I}^2$. Note that if the couple is cohabiting, the utility of the two partners is decreased by γ. This parameter 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.[20]

3.4.2 Wages

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

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

[20] 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 marriage and cohabitation choices closely using a theoretical framework close to ours, without needing to introduce a stigma towards cohabitation. This result 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 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:

$$z_t^s = z_{t-1}^s - (1 - P_t^s)\mu + \zeta_t^s, \text{ where } \zeta_t^s \overset{\text{i.i.d.}}{\sim} \mathcal{N}(0, \sigma_{\zeta}^{2_s}), \text{ and } \zeta_1^s = z_1^s. \tag{3.6}$$

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 $P_t^m = 1$.[21] 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.

3.4.3 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:

$$Q_t = [(\underbrace{d_t}_{\text{money}})^{\nu} + \kappa(\underbrace{\phi}_{\text{men's time}} + \underbrace{(\phi + (1 - P_t^f)(1 - \phi))}_{\text{women's time}})^{\nu}]^{\frac{1}{\nu}}, \text{ where } 0 < \nu < 1, \tag{3.7}$$

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

$$Q_t^s = [(\underbrace{d_t^s}_{\text{money}})^{\nu} + \kappa(\underbrace{\phi}_{\text{time}})^{\nu}]^{\frac{1}{\nu}}. \tag{3.8}$$

The parameter ν captures the degree of substitutability between women's time and the use of market goods in the production of home goods.[22] This structure implies that when the

[21] 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.

[22] Market goods d include both durables goods, such as the washing machine, and services that can be bought on the market, such as a nanny that takes care of children.

relative price of d_t decreases and when wages go up,[23] women spend less time producing household goods and their employment outside the home increases.

3.4.4 Budget Constraints

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

$$a^s_{t+1} = Ra^s_t + w^s_t(1-\phi) - c^s_t - d^s_t, \text{ with } a^s_{t+1} \geq 0, \tag{3.9}$$

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:

$$a^f_{t+1} + a^m_{t+1} = Ra_t + w^m_t(1-\phi) + P^f_t w^f_t(1-\phi) - c^f_t - c^m_t - d_t, \text{ with } a_{t+1} \geq 0, \tag{3.10}$$

When a couple divorces in t, we assume

$$a^m_t + a^f_t = \delta a_t,$$

where δ is the fraction of total assets a_t left after divorce. We assume $\delta = 1$ for breakup.[24] 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.

3.4.5 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. The state variable of a single is

[23] The relative price of d_t is normalized to 1 in equations (3.7) and (3.8)

[24] The assumption that divorce erodes a fraction of wealth is common to Cubeddu and Ríos-Rull (2003). In appendix 3.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 evidence of 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 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. When we tried estimating the model allowing for a positive cost of breakup, this parameter was not identified.

$\omega_t^i = \{a_t^i, z_t^i\}$, 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 as

$$
\begin{aligned}
& V_t^{iS}(\omega_t^i) = \max_{\mathbf{q}_t^i} u(c_t^i, Q_t^i) + \beta E_t \Bigg\{ && (3.11) \\
& (1-\lambda_{t+1}) V_{t+1}^{iS}(\omega_{t+1}^i) + && \text{(no meeting)} \\
& \lambda_{t+1}(1-M_{t+1})(1-C_{t+1}) V_{t+1}^{i,S}(\omega_{t+1}) + && \text{(meeting, stay single)} \\
& \lambda_{t+1} M_{t+1} V_{t+1}^{i,M}(\Omega_{t+1}) + && \text{(meeting, marry)} \\
& \lambda_{t+1} C_{t+1} V_{t+1}^{i,C}(\Omega_{t+1}) \Bigg\}, && \text{(meeting, cohabit)} \\
& \text{s.t. budget constraint of singles (3.9),}
\end{aligned}
$$

where $V^{i,M}$ and $V^{i,C}$ are the individual values of being married and cohabiting. Note that if the couple 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. Variables M_{t+1} and C_{t+1} are not solely choices of the individual, but of the (potential) couple: I explain in subsection 3.4.7 how these are set.

3.4.6 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 or 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 $1-\theta^f$.[25] 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:

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

$$V_t^{\hat{M}}(\Omega_t^{\hat{M}}) = \max_{\mathbf{q}_t^M} \Bigg\{ \tag{3.12}$$

$$(1-D_t)\{\theta^f u(c_t^f, Q_t) + (1-\theta^f)u(c_t^m, Q_t) + \psi_t + \beta E_t V_{t+1}^{\hat{M}}(\Omega_{t+1}^{\hat{M}})\} + \qquad \text{(stay married, } D_t = 0\text{)}$$

$$D_t\{\theta^f V_t^{fS}(\omega_t^f) + (1-\theta^f)V_t^{mS}(\omega_t^m)\}\Bigg\} \qquad \text{(divorce, } D_t = 1\text{)}$$

if $D_t = 0$: s.t. budget constraint of couples (3.10)

if $D_t = 1$: s.t. budget constraint of singles (3.9) for $s \in \{f, m\}$,

$$a_t^m + a_t^f = \delta a_t, \qquad \text{(loss of assets)}$$

$$V_t^{fS}(\omega_t^f) \geq W_t^{f\hat{M}}(\Omega_t^{\hat{M}}) \text{ for } s \in \{f, m\}, \qquad \text{(divorce if both agree)}$$

where the last line of the above problem means that divorce can happen only if both spouses agree. 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

$$W_t^{i\hat{M}} = u(\tilde{c}_t^i, \tilde{Q}_t) + \psi_t + \beta E_t V_{t+1}^{i\hat{M}}(\Omega_{t+1}^{\hat{M}}), \tag{3.13}$$

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 (3.12) 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 (3.11), 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, suppose only one spouse wishes to divorce under the divorce allocation dictated by the law where assets are split equally. In that case, they will "bribe" the other by offering a larger share of assets to make them indifferent between staying married and divorcing.[26]

[26]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.

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 the vector $\mathbf{q}_t^{\overline{M}} = \{\tilde{a}_{t+1}, \tilde{d}_t, \tilde{c}_t^m, \tilde{c}_t^f, \tilde{P}_t^f\}$.[27] The formal problem of a couple entering t as married is:

$$V_t^{\overline{M}}(\Omega_t^{\overline{M}}) = \max_{\mathbf{q}_t^M} \Big\{ \tag{3.14}$$

$$(1-D_t)\{\theta_t^f u(c_t^f, Q_t) + \theta_t^m u(c_t^m, Q_t) + \psi_t + \beta E_t V_{t+1}^{\overline{M}}(\Omega_{t+1}^{\overline{M}})\} + \qquad \text{(stay married, } D_t = 0)$$

$$D_t\{\theta_t^f V_t^{fS}(\omega_{t+1}^f) + \theta_t^m V_t^{mS}(\omega_t^m)\}\Big\} \qquad \text{(divorce, } D_t = 1)$$

if $D_t = 1$: s.t. budget constraint of single (3.9) for $i \in \{f, m\}$,

$$a_t^m + a_t^f = \delta a_t, \qquad \text{(loss of assets)}$$

$$a_t^m = a_t^f, \qquad \text{(community property regime)}$$

if $D_t = 0$: s.t. budget constrain of the couple (3.10),

$$W_t^{s\overline{M}}(\Omega_t^{\overline{M}}) \geq V_t^{sS}(\omega_t^s) \text{ for } s \in \{f, m\}, \qquad \text{(participation constraints)}$$

where, whenever $D_t = 0$ and the participation constraint of person s binds, the new Pareto weight is: $\theta_{t+1}^s = \text{argmin}_{\theta^*}|\theta_t^s - \theta^*|$ such that participation constraints hold. When both participation constraints are satisfied the Pareto weights stay constant. Divorce happens when at leas one of the participation constraints cannot be met for any change in the bargaining power. 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 func-

[27] Since we analyze community property states, the assets of husband and wife are jointly owned and hence a_t^m, a_{t+1}^m and a_t^f, a_{t+1}^m do not appear separately as state and choice variables. Under the mutual consent regime we included these as the state and choice varibles because they had a role during divorce (only), since one spouse could bribe the other to divorce.

tional 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.

Note that the choice to use divorce as an outside option is not casual. It reflects the evidence that changes in the right divorce modified the behavior of married couples Voena (2015).

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. Variable χ_t is the share of assets going to the woman in case of breakup.[28] Third, the time utility of the cohabiting couple is decreased by γ. The problem of the cohabiting couple can be found in section 3.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.[29] 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

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

[29]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.

rule decided by the couple in the previous period.[30] They show that this regime is always preferred to community property if outside options are invariant to property right regimes. This result implies that if i) the costs of breaking up and divorcing were the same and ii) 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 allow us to generate a positive number of marriages and match the data.

3.4.7 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:

$$\ln(a_t^p) = \ln(a_t^s) + \overline{a}^g + \epsilon^a, \tag{3.15}$$

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

$$\ln(z_t^p) = \alpha(\ln(\overline{z}_t^{s^*,i^*}) + \epsilon_t^z) + (1-\alpha)\ln(z_t^s), \tag{3.16}$$

where $\overline{z}_t^{s^*,i^*}$ represents the average productivity of singles of gender s^*, z_t^s 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 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.[31]

[30] 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.

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

Formally, for relationship $J \in \{M, C\}$ the set is

$$\Theta_t^J(\Omega_t^J, \omega_t^f, \omega_t^m) = \left\{\theta_t^f : V_t^{fJ}(\Omega_t^J) \geq V_t^{fS}(\omega_t^f), V_t^{mJ}(\Omega_t^J) \geq V_t^{mS}(\omega_t^m)\right\}. \tag{3.17}$$

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.[32] Formally $\theta_t^{J,f}$ is set to :

$$\theta_t^{J,f} = \arg\max_{\theta_t^f \in \Theta_t^J} \Upsilon^J(\theta_t^f, \Omega_t^{J-1}, \omega_t^f, \omega_t^m), \tag{3.18}$$

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

$$\Upsilon^J(\theta_t^f, \Omega_t^{J-1}, \omega_t^f, \omega_t^m) = \left[V_t^{fJ}(\Omega_t^{J-1}, \theta_t^f) - V_t^{fS}(\omega_t^f)\right] \times \left[V_t^{mJ}(\Omega_t^{J-1}, 1-\theta_t^f) - V_t^{mS}(\omega_t^m)\right]. \tag{3.19}$$

3. Four possible situations can arise:
   - $\Theta_t^M = \emptyset$ and $\Theta_t^C \emptyset \Rightarrow$ stay single.
   - $\Theta_t^M \neq \emptyset$ and $\Theta_t^C \emptyset \Rightarrow$ marry.
   - $\Theta_t^M = \emptyset$ and $\Theta_t^C \neq \emptyset \Rightarrow$ cohabit.
   - $\Theta_t^M \neq \emptyset$ and $\Theta_t^C \neq \emptyset \Rightarrow$ 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.

Note that the algorithm presented above determines the value of C_t and M_t, which I introduced while presenting the value function of singles (3.11).

3.5 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

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

reduces the burden on structural estimation.[33] 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.

3.5.1 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:

$$\ln(w^m_{i,t,s,sur}) = \iota^m_0 + \iota^m_1 \cdot t + \iota^m_2 \cdot t^2 + \delta_s + \nu_{sur} + u^m_{i,t,s,sur}, \tag{3.20}$$

where i stands for individual, t for age, s for state and sur for survey year. Moreover, $u^m_{i,t,s,sur} = z^m_t + e^m_{i,t,s,sur}$, where z^m_t follows equation 3.6, while $e^m_{i,t,s,sur}$ 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 3.E.1. Then, using the residuals $\hat{u}^m_t$, we estimate through GMM 1) the variance of the permanent component of income $\sigma^{2^m}_\zeta$, 2) the variance of the measurement error $\sigma^{2^m}_e$ using the following conditions:

$$\begin{aligned} E((\Delta\hat{u}^m_t)^2) &= \sigma^{2^m}_\zeta + 2\sigma^{2^m}_e \\ E(\Delta\hat{u}^m_t \Delta\hat{u}^m_{t-1}) &= -\sigma^{2^m}_e \end{aligned} \tag{3.21}$$

Results are reported in table 3.6.

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-

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

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 (3.20) 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.[34] Women participate in the labor market if

$$\gamma'\mathbf{Z}_{i,t,s,sur} + \pi_{i,t,s,sur} > 0, \tag{3.22}$$

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:

$$\ln(w^f_{i,t,s,sur}) = \iota^f_0 + \iota^f_1 \cdot t + \iota^f_2 \cdot t^2 + \delta_s + \nu_{sur} + \varphi_{i,t,s,sur} + u^f_{i,t,s,sur}, \tag{3.23}$$

where i stands for individual, t for age, s for state and sur for survey year. Moreover, $u^f_{i,t,s,sur} = z^f_t + e^f_{i,t,s,sur}$. z^f_t follows equation 3.6, while $e^f_{i,t,s,sur}$ 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 3.E.3 and 3.E.2. We then use the regression residuals from the second step $\hat{u}^m_t$ to estimate through GMM 1) the variance of the permanent component of income σ^{2f}_ζ, 2) the variance of the measurement error σ^{2f}_e using the following conditions:[35]

$$\begin{aligned}
E(\Delta\hat{u}^f_t | P^f_t = 1, P^f_{t-1} = 1) &= \sigma^f_\pi \frac{\phi(\tau_t)}{1-\Phi(\tau_t)}, \\
E((\Delta\hat{u}^f_t)^2 | P^f_t = 1, P^f_{t-1} = 1) &= \sigma^{2f}_\zeta + \sigma^{2f}_\pi + 2\sigma^{2f}_e + \tau_t \frac{\phi(\tau_t)}{1-\Phi(\tau_t)}, \\
E(\Delta\hat{u}^f_t \Delta\hat{u}^f_{t-1} | P^f_t = 1, P^f_{t-1} = 1, P^f_{t-2} = 1)) &= -\sigma^{2f}_e.
\end{aligned} \tag{3.24}$$

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 3.6.

[34] 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.

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

Table 3.6:
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	σ_ζ^{2f}	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	σ_ζ^{2m}	0.0417

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

3.5.2 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 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.[36] 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), $\sigma_{\zeta,1}^{2m}(\sigma_{\zeta,1}^{2f})$ is taken directly from the PSID data. The parameters regarding the mating market, contained in equations 3.15

[36]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 use in the structural estimation are based on the behavior of people born in those years.

and 3.16, 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.[37] Since we set these parameters before the structural estimation occurs, we cannot perfectly match the mating market moment we targeted. The correlation in log wages of couples in the PISD is 0.58 versus 0.62 in the simulated sample,[38] while the share of people with 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.

[37]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. We follow the partners' behavior while they are in a relationship with the person in our fictional sample. Figure 3.I.1 shows the mean and variance of productivity and wealth by age, both for agents belonging to the "fictional sample" and their "partners." The variables of interest are similar for the two groups, which means that the two groups are symmetric concerning these variables.

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

Table 3.7:
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 $t = 1$	$\sigma^{2f}_{\zeta,1}$	0.54	PSID
Var. m's productivity $t = 1$	$\sigma^{2m}_{\zeta,1}$	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)

3.5.3 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 3.1,[39] 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),[40] differences in female labor force participation between marriage

[39]Note that the sample used for estimating equation 3.1 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.

[40]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.

and cohabitation (2) and 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 all the others are computed using the PSID.[41] 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 for 30000 fictional individuals to obtain their simulated behavior for the given parametrization. The next step is to perform stratified sampling on the simulated population 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

$$\hat{\vartheta} = \arg\min_{\vartheta} \qquad (\mathbf{m} - \mathbf{m}_{\vartheta})'\mathbf{W}(\mathbf{m} - \mathbf{m}_{\vartheta}), \tag{3.25}$$

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. We perform the minimization of this object function 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 3.C, we describe how the algorithm TikTak works and how we modify it to allow for the possibility of running it in parallel.

3.5.4 Identification

This section describes 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 require women's time. Parameter μ affects the gap in female labor supply for married and cohabiting couples. When

[41]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.

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 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 cannot generate a marriage surplus large enough to match the number of ever married people. 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 it 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 3.1. 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.

3.5.5 Model Fit

Table 3.8 reports the results of the structural estimation. The estimated standard deviation σ_ψ of the transitory match quality shock is 0.76, while the 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 3.9. The model generally matches 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 (3.1) 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.[42] We use two linear models on a sample of cohabitation spells-year to study the role of unilateral divorce (independent variable) on the risk of breaking up and separating. In the first model the dependent variable is a dummy taking value one in case of a marriage, while in the second model it is a dummy taking value 1 in case of a breakup.[43] The results, reported in table 3.9, 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 prop-

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

[43]Since the aim of this exercise is to study how the pool of cohabiting couples changes after the reform, the dependent variable takes value one if the cohabitation spell started under the unilateral divorce regime and zero otherwise. We control for duration and we have age, year the relationship started and state fixed effects.

erty 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 (3.1) is -0.09, while it was -0.16 under community property.[44] 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.

Table 3.8:
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

[44]Note that we do not expect to match exactly the empirical coefficient (3.1) 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 3.9.

Table 3.9:
Model fit and validation

Estimated Moments	Model	Data	95% CI
Hazards over Time	fig. 3.H.1	fig. 3.H.1	fig. 3.H.1
Share Ever Cohabited and Married	fig. 3.H.2	fig. 3.H.2	fig. 3.H.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 (3.1)	-0.16	-0.11	[-0.21,-0.02]
External Moments	**Model**	**Data**	
Unilateral Divorce and linkelihood of marrying	-0.04	-0.08	[-0.20,0.04]
Unilateral Divorce and linkelihood of braking up	-0.06	-0.12	[-0.22,-0.02]
Women wages over their life-cycle	fig. 3.H.3	fig. 3.H.3	fig. 3.H.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]

3.6 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

$$\text{Variable of Interest}_{i,a,t} = \sum_{j=-5}^{5} \beta_j^{Uni} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_a + \epsilon_{i,t} \tag{3.26}$$

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 3.3 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.[45] 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.[46] 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.[47] 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 within the household.

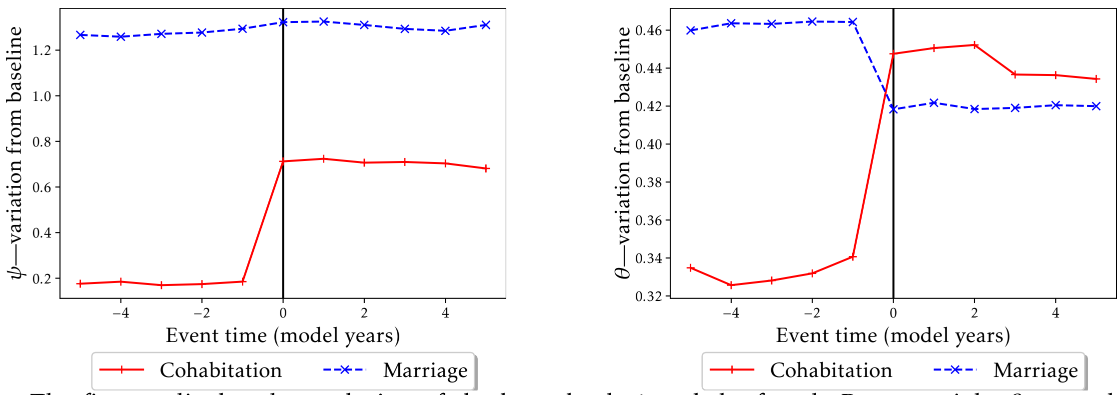
[45] A more in depth analysis reveals that the distribution of match quality at meeting of cohabiting couples dominates that of married couples. See figure 3.I.2.

[46] 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.

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

Figure 3.3:
Event Studies Around the Introduction of Unilateral Divorce–Simulated Data

(a) Love Shock ψ at meeting (b) Women Pareto Weight θ at meeting



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. Note that the sample used to obtain these figures is created simulating 300000 agents behaving following the policy function implied by the model estimated with the parameters of table 3.8.

Risk-sharing and consumption insurance. The veto over divorce in the mutual consent regime and the high cost of divorce are commitment technologies that enforce cooperation within marriage. What is the quantitative relevance of these mechanisms on married and cohabiting couples' ability to share risk? To answer this question, we study how income shocks affect the changes in consumption over time for men in a couple.[48] 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.[49] 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

[48]Results for women in a couple are displayed in table 3.I.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 3.I.1 are largely driven by the different degree of female labor force participation under different divorce regimes and partnership types.

[49]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. The first relationship of a man that cohabited before marrying 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.

fixed effects. Coefficient μ is informative about how much of the change in income translates into changes in consumption. Consequently, a low value of μ corresponds to a high degree of consumption insurance. The results displayed in table 3.10 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.[50] 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 3.10 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 3.10 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 3.I.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.

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

Table 3.10:
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

$$\log(c)_{i,a,t} = \sum_{j=-3}^{3} \beta_j^{Split} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_a + \epsilon_{i,a,t}, \tag{3.27}$$

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 3.I.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.

3.7 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 3.11 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 in the changes in well-being, we repeat the welfare analysis assuming that cohabitation is no longer a choice.[51] 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 3.11 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 men's welfare losses 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.

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

Table 3.11:
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			
-102.2	-105.73	-90.87	-92.19
Welfare Losses with Unilateral Divorce			
			
13990.74 $		6662.11 $	

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.

3.8 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 3.12 reports the share of people that cohabited at 39 and the average years spent cohabiting under the baseline scenario and the counterfactual.[52] The results show that under the counterfactual only 29% of people would have ever 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 3.12 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%:[53] 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 splitting cost. 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 3.12 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 improved 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

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

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

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 from 2.19 to 2.27. Similar to a reduction in the gender wage gap, improvements in home production technology 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 caused a decline of marriage with respect to singleness, as Greenwood et al. (2016) claim, and a change in the relative convenience of partnership contracts.

No stigma on cohabitation. Table 3.12 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 play an important role in 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 can generate a positive surplus of marriage with respect to cohabitation for specific individuals.

Table 3.12:
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.

3.9 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 and 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, 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 and 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 have 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. This is becuase 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.

Appendix

3.A History of US divorce Laws

Table 3.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)

3.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 moves 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.[54] 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

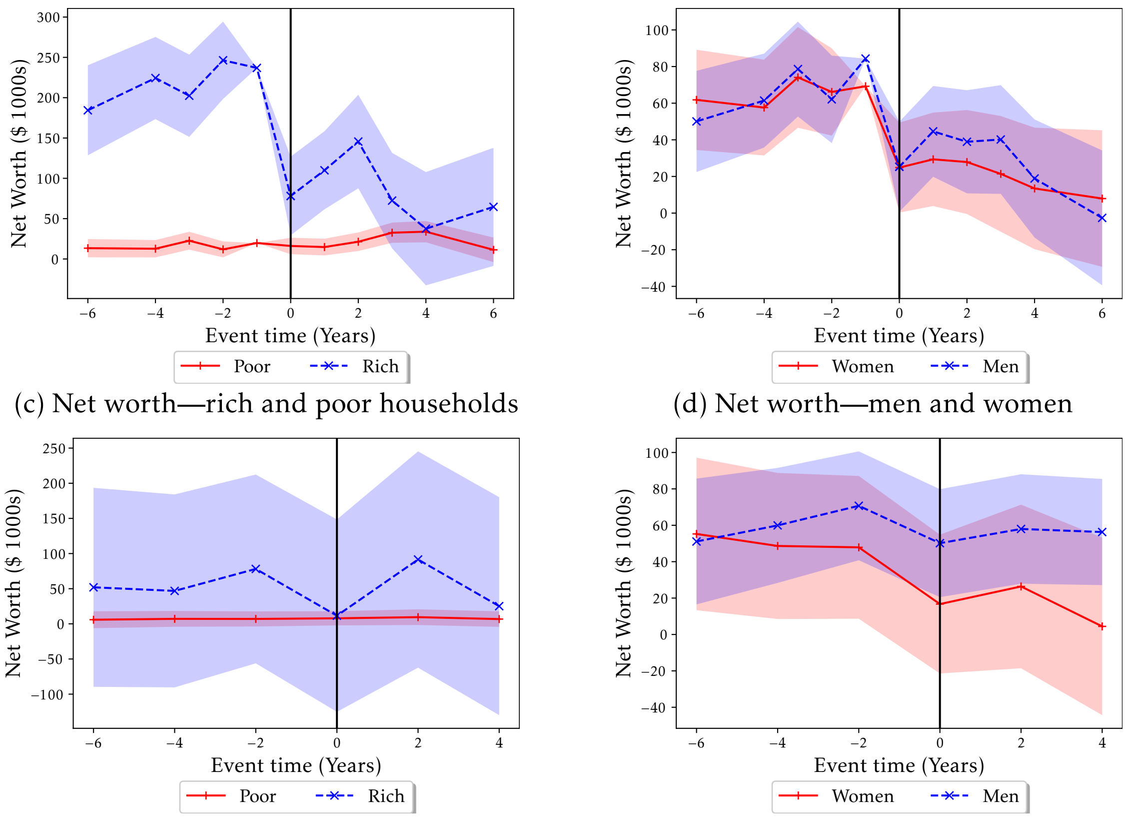
$$\text{Net worth}_{i,a,t,y,ma} = \sum_{j=-6}^{4} \beta_j^{Split} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_a + \alpha_y + \alpha_{ma} + \epsilon_{i,t}, \tag{3.28}$$

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.[55] Figure 3.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 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.

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

[55]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 3.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.)

3.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 (3.25). 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,[56] 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

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

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.

3.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 otherwise.[57] The formal problem that a cohabiting couple at t solves is:

[56] 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.

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

$$V_t^C(\Omega_t^C) = \max_{\mathbf{q}_t^C} \Bigg\{ \tag{3.29}$$

$$(1-S_t)\{\theta_t^f u(c_t^f, Q_t) + \theta_t^m u(c_t^m, Q_t) + \psi_t - \gamma + \beta E_t V_{t+1}^C(\Omega_{t+1}^C)\} + \quad \text{(keep cohabiting)}$$

$$M_t\{\theta_{t+1}^f u(c_t^f, Q_t) + \theta_{t+1}^m u(c_t^m, Q_t) + \psi_t + \beta E_t V_{t+1}^M(\Omega_{t+1}^M)\} + \quad \text{(marry)}$$

$$S_t\{\theta_t^f V_t^{fS}(\omega_{t+1}^f) + \theta_t^m V_t^{mS}(\omega_t^m)\}\Bigg\} \quad \text{(breakup)}$$

if $S_t = 1$: s.t. budget constraint of the single (3.9) for $i \in \{f, m\}$,

$$a_t^m + a_t^f = a_t, \quad \text{(no asset loss)}$$

$$a_t^f = \chi_t a_t, \quad \text{(asset division)}$$

if $S_t = 0$: s.t. budget constraint of the couple (3.10) ,

$$W_t^{sC}(\Omega_t^C) \geq V_t^{sS}(\omega_t^m) \text{ for } s \in \{f, m\} \quad \text{(participation constraints)}$$

where, whenever $S_t = 0$ and the participation constraint of person s binds, the new Pareto weight is: $\theta_{t+1}^s = \text{argmin}_{\theta^*}|\theta_t^s - \theta^*|$ such that participation constraints hold. When both participation constraints are satisfied the Pareto weights stay constant. Breakup happens when at leas one of the participation constraints cannot be met for any change in the bargaining power. The individual value of cohabitation conditional on $S_t =$ is W_t^{iC} for $i \in \{f, m\}$, and it is defined as

$$W_t^{iC} = u(\tilde{c}_t^i, \tilde{Q}_t^i) + \psi_t - \gamma + \beta E_t V_{t+1}^{iC}(\Omega_{t+1}^C), \tag{3.30}$$

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 (3.29) 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 (3.11). 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.

3.E Estimation of Income Processes

Table 3.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 3.6.

Table 3.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 3.6.

Table 3.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%.

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

Relationship Choice - Linear State Trends

Table 3.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 3.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 3.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 3.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

So far, our empirical analysis 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 complete analysis by studying singles' choice, who can decide every month to stay single, cohabit or 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 in table 3.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 instead of 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 3.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%.

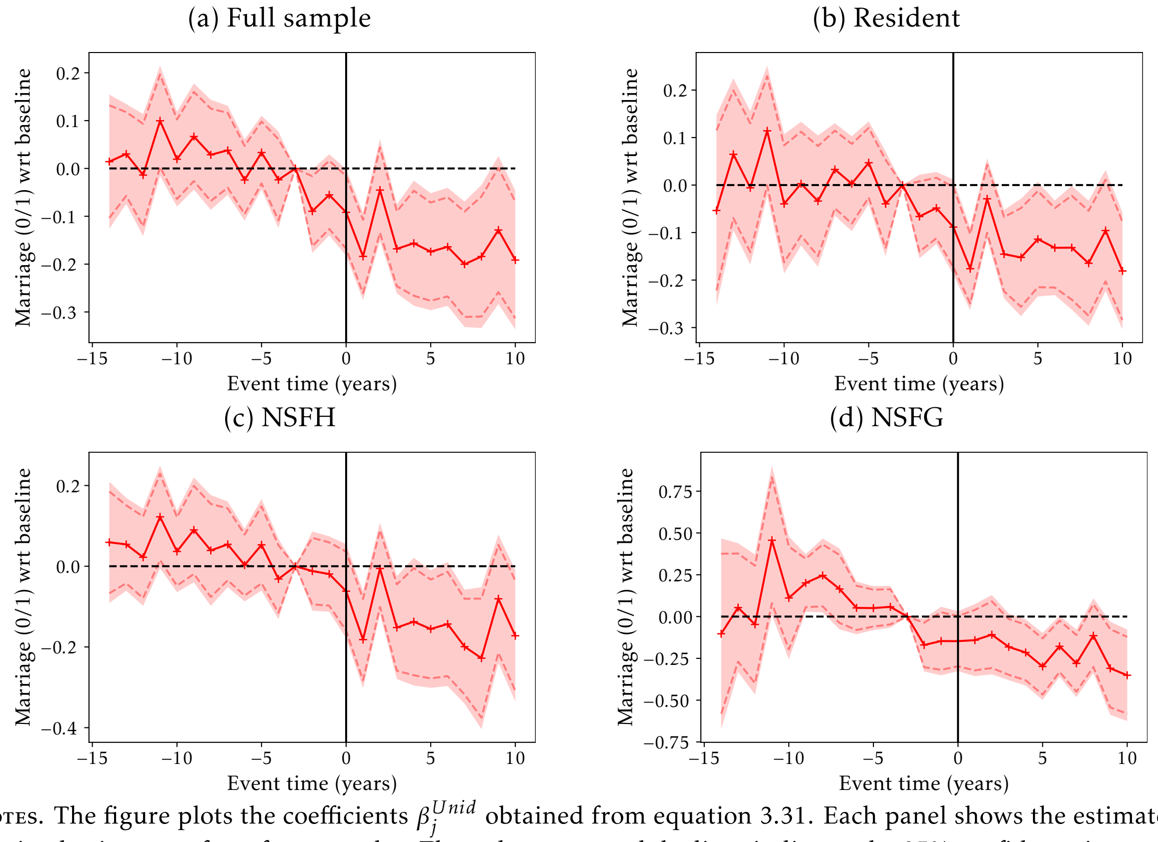
Relationship Choice - Event Study

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

$$\text{married}_{i,s,y,b} = \sum_{j=-32}^{55} \beta_j^{Unid} \cdot \mathcal{I}(t=j) + \alpha_0 + \alpha_b + \alpha_y + \alpha_s + \gamma' \mathbf{Z_i} + \epsilon_{i,s,y,b}, \tag{3.31}$$

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 3.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-way fixed effect estimates over-weights observations that are closer to the policy change, as Goodman-Bacon (2018) notice. This matters for our case as individuals might not recall the exact date at which the couple started living together. This is consistent with the results reported in figure 3.F.1: β_j^{Unid} are negative (even though not significant) 1 or 2 years before the policy changes.

Figure 3.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 3.31. 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 3.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 because 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. Our model captures this heterogeneity. Couples with a high relationship quality are less sensitive to the law changes and are also more likely to produce large quantities of public goods, which include children.

Table 3.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%.

3.G More evidence on the impact of unilateral divorce on cohabitation duration

Table 3.G.1:
Duration model: risk of marriage for cohabiting couples.

	Dependent variable: Married (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
Unilateral Divorce	-0.021^{***}	-0.018^{***}	-0.020^{***}	-0.028^{***}
	(0.004)	(0.004)	(0.004)	(0.010)
State Fixed effects	Yes	Yes	Yes	Yes
Year started Fixed Effect	Yes	Yes	Yes	Yes
Duration Polynomial	Yes	Yes	Yes	Yes
Socio-economic controls	Yes	Yes	Yes	Yes
Observations	137,848	81,756	77,662	60,186
R^2	0.007	0.006	0.008	0.006

Notes: the sample is of the month-cohabitation type. In periods where the couple keeps cohabiting or breaks up, the dependent variable takes value 0. 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%.

Table 3.G.2:
Duration model: risk of breaking up for cohabiting couples.

	Dependent variable: Broke up (0/1)			
	Full Sample	Resident	NSFH	NSFG
	(1)	(2)	(3)	(4)
Unilateral Divorce	-0.008^{***}	-0.008^{***}	-0.006^{**}	-0.015^{**}
	(0.002)	(0.002)	(0.002)	(0.006)
State Fixed effects	Yes	Yes	Yes	Yes
Year started Fixed Effect	Yes	Yes	Yes	Yes
Duration Polynomial	Yes	Yes	Yes	Yes
Socio-economic controls	Yes	Yes	Yes	Yes
Observations	137,848	81,756	77,662	60,186
R^2	0.003	0.003	0.005	0.004

Notes: the sample is of the month-cohabitation type. In periods where the couple keeps cohabiting or gets married, the dependent variable takes value 0. 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%.

3.H Model Fit

Figure 3.H.1:
Hazards by duration of spells: data and simulations

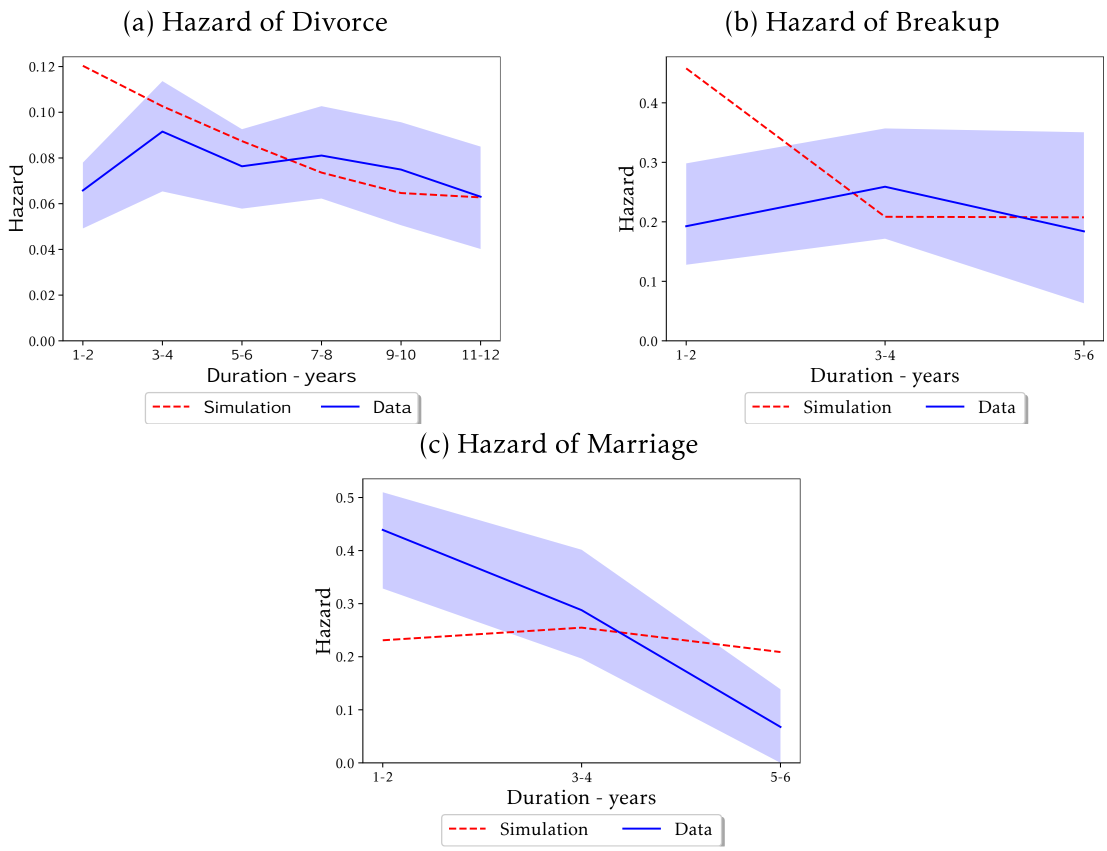


Figure 3.H.2:
Share ever cohabited and married: data and simulations

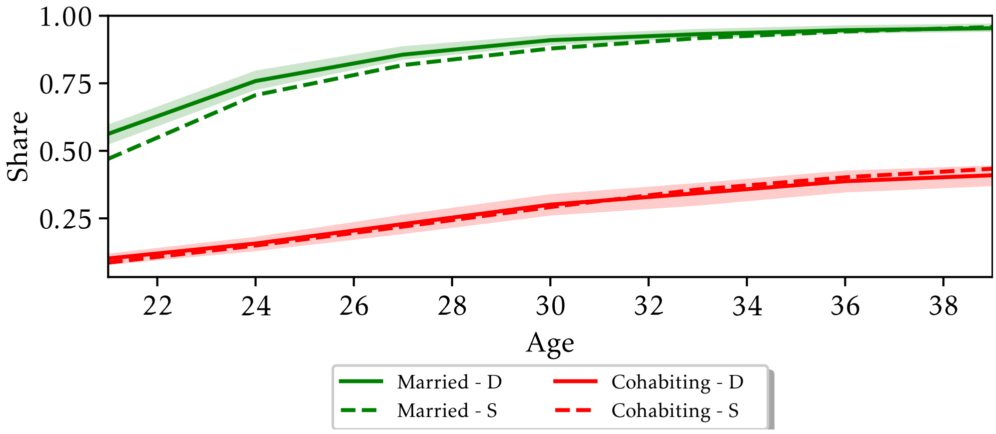
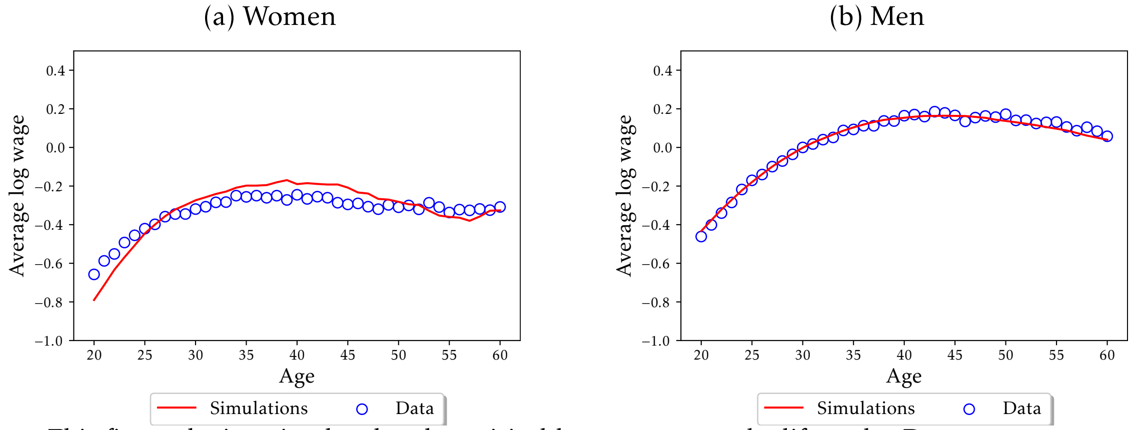


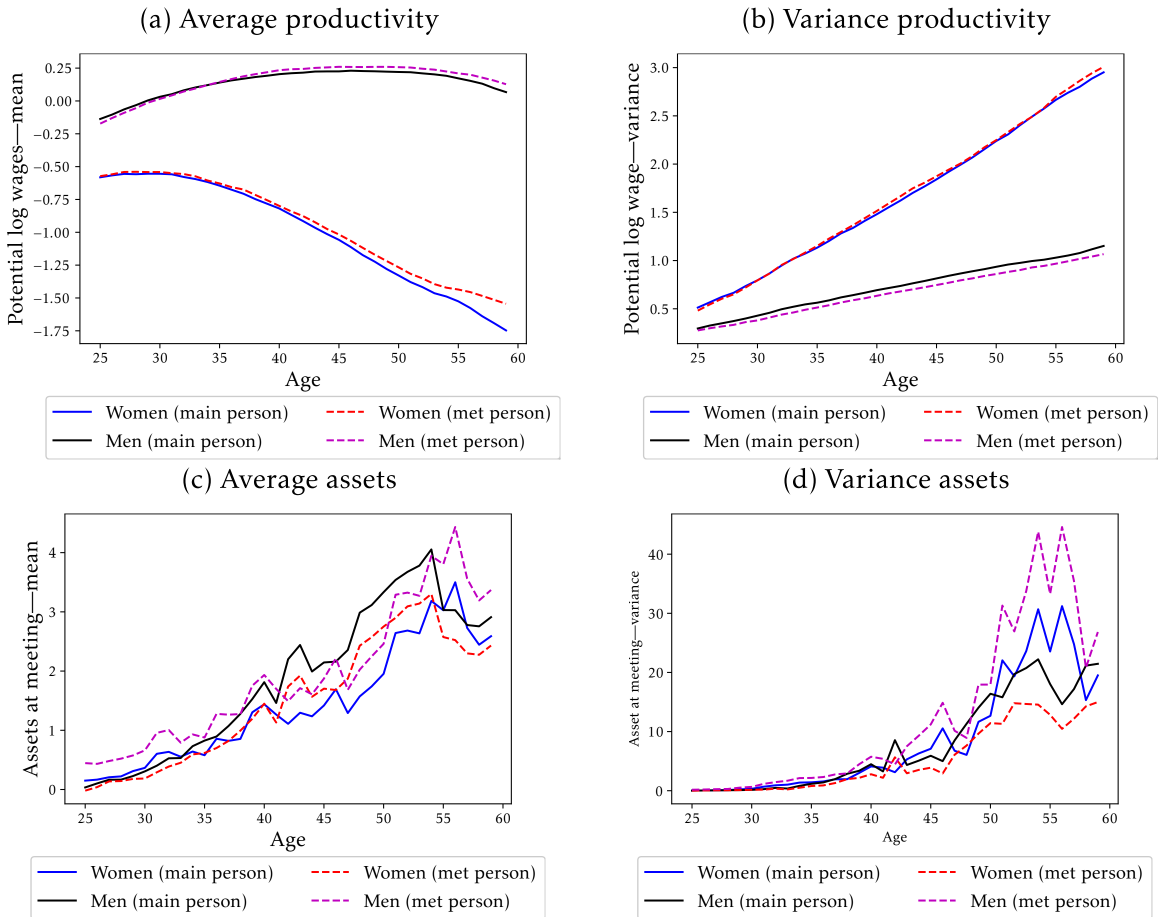
Figure 3.H.3: Low wages over the life cycle: simulations and data



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.

3.I Additional Figures and Tables

Figure 3.I.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 3.I.2:
Cumulative distribution of love shock ψ at meeting

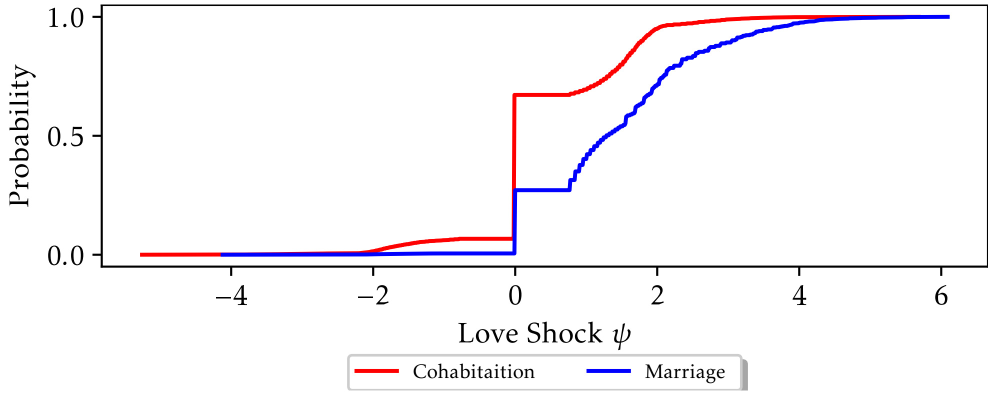
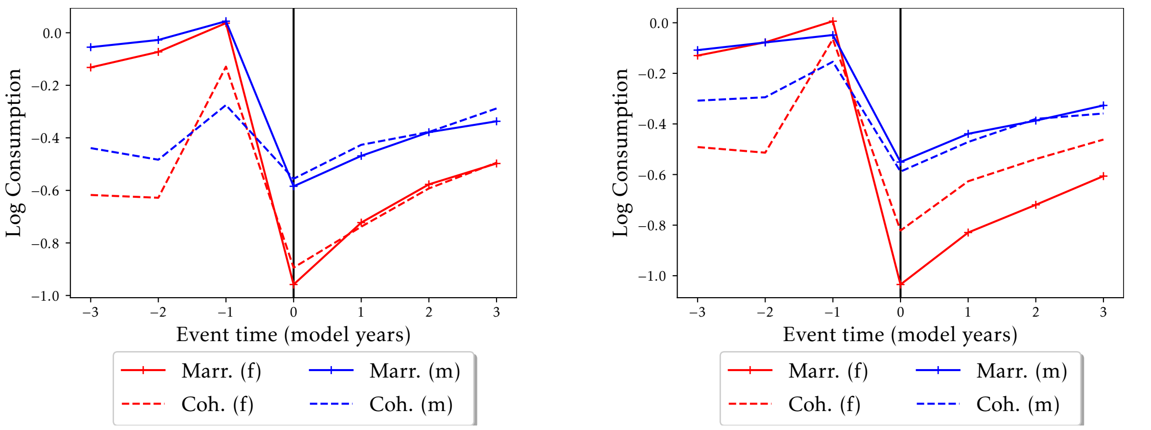


Figure 3.I.3:
Event studies of log consumption around divorce–simulated data

(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.)

Table 3.I.1:
Partnership type and consumption insurance against income shocks

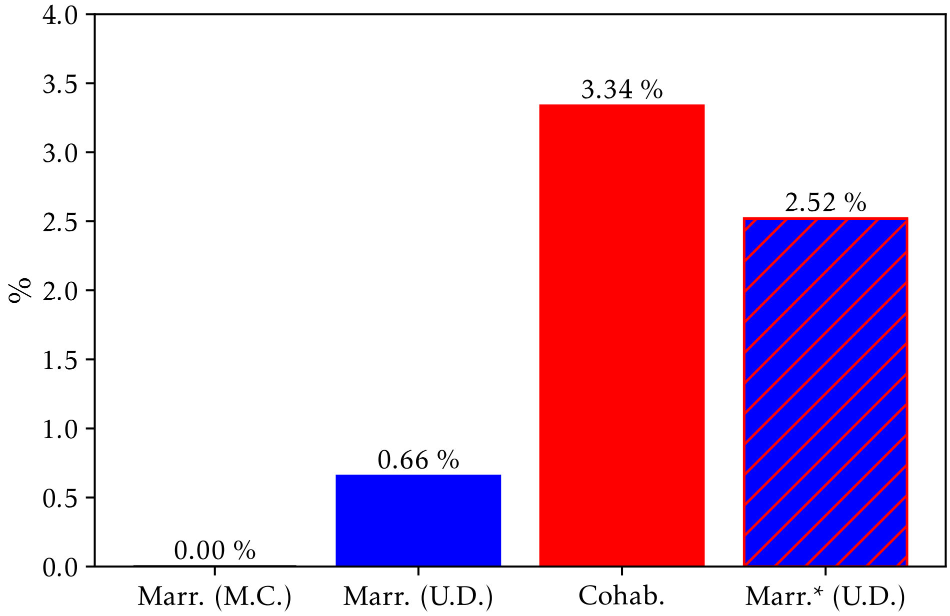
	Married and Cohabiting Women		
	Married		Cohabiting
	M.C.	U.D.	
Baseline	0.201	0.164	0.329
Only marriages preceded by cohabitation	0.097	0.183	-
Only marriages not preceded by cohabitation	0.210	0.160	-
Marriages with cohabitation selection	0.329	0.331	-

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.

Figure 3.I.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.

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