Essays on the Economics of Family Formation, Dissolution, and Bargaining

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To my parents, Miguel and María Esther,
To my brothers, Adrián and Javier,
To my wife, Vanessa.

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Abstract

This thesis sheds light on several aspects of the economics of marital formation, dissolution, and bargaining. The first chapter focuses on the relationship between divorce law and family wellbeing, and shows that lowering the cost of divorce can reduce spousal conflict. The second chapter analyzes the effects of property division laws upon divorce on marital instability and female labor supply. Results suggest that a redistribution of property rights over family assets in case of divorce towards the financially weaker spouse, usually the wife, may increase marital instability and reduce female labor supply. The third chapter examines the role of sex ratios in college in explaining family formation patterns of young adults. Empirical evidence suggests that individuals who are exposed to a larger fraction of opposite-sex school mates are more likely to be married or residing with a partner from the same field of study shortly after finishing school.

Resumen

Esta tesis arroja luz sobre algunos aspectos de la economía de la formación, disolución y negociación familiar. El primer capítulo se centra en la relación entre la regulación sobre el divorcio y el bienestar de la familia, y muestra que una disminución del coste del divorcio puede reducir el nivel de conflicto entre esposos. El segundo capítulo analiza los efectos de las leyes de división de activos en caso de divorcio sobre la inestabilidad matrimonial y la oferta de trabajo de las mujeres. Los resultados sugieren que una redistribución de los derechos de propiedad sobre los activos familiares en caso de divorcio en favor de la parte financieramente más débil, habitualmente la mujer, puede aumentar la inestabilidad matrimonial y reducir la oferta de trabajo de las mujeres. El tercer capítulo examina el papel de la ratio de sexos en la universidad en explicar el patrón de formación familiar de adultos jóvenes. La evidencia empírica sugiere que los individuos que están expuestos a una mayor proporción de compañeros del sexo opuesto tienen más probabilidad de estar casados o residiendo con una pareja del mismo campo de estudios, poco después de finalizar la universidad.

Foreword

This dissertation consists of three self-contained chapters that deal with several aspects of the economics of marital formation, dissolution, and bargaining. The first two chapters are devoted to improve our understanding of the relationship between family policies and household outcomes, where the effects of those policies on incentives and behaviors play a key role. The main point here is that the rules governing the dissolution of marriages affect the value of the spouses' outside option, and then, their relative bargaining position within the marriage. These changes in the intra-household bargaining position, in turn, are shown to have an effect on the level of family conflict, marital instability, and spouses' labor supply. The third chapter provides new insights on the functioning of marriage markets. The study examines the role of sex ratios in college in explaining family formation patterns of young adults.

The first chapter investigates whether lowering the cost of divorce can reduce domestic violence. The cost of divorce influences the bargaining position of spouses, and thus, their behavior within the marriage. This study takes advantage of a large and unexpected reform of the divorce regime in Spain, which allowed for unilateral and no-fault divorce, and eliminated the pre-existing 1-year mandatory separation period, to estimate the causal effects. This reform dramatically reduced the cost of exiting a partnership for married couples, but not for unmarried ones, which favors a difference-in-differences identification strategy. This study analyzes several measures of spousal conflict, ranging from self-reported spousal abuse and technical definitions of spousal violence based on recorded behavior, to more extreme measures of well-being such as partner homicide. Results suggest a decline of 27-36 percent in spousal conflict and around 30 percent in extreme partner violence as a consequence of the reform. Moreover, spousal violence has been found to decrease among couples who remain married after the legal modification, which suggests an important role for changes in bargaining within the marriage when divorce becomes a more credible (cheaper) option. The results are not driven by selection and are robust to a variety of checks.

The second chapter analyzes how the relative bargaining position of spouses affects the incidence of marital dissolution and the labor supply decision of intact couples. The study identifies exogenous variation in bargaining position within the household by exploiting a natural experiment in Spain, where different regions have different rules to divide marital property in case of divorce. This study benefits from two law changes to the separation of property regime in Catalonia, with opposite expected effects on the bargaining position of spouses. Results suggest that a reform that unexpectedly improved the position of the wife within the marriage increased the divorce rate in around 13 percent in the short run, and although this effect seemed to dissipate over time, it remained positive one decade afterwards. For intact couples, results show that the same reform caused a reduction in female labor supply of between 0.6 and 2.5 hours per week, and also a reduction in their probability of employment of 2 percent. Moreover, when the previous improvement in wives' bargaining position was undone by a reform to the scope of marital contracts, female labor supply reacted in the opposite way, with an increase in hours worked and the probability of employment.

The third and final chapter[1] examines the role of sex ratios in college in explaining family formation patterns. Average age at first marriage suggests that the initial search for a spouse often takes place before entering the labor market, especially among more educated individuals. This chapter studies whether the proportion of classmates of the opposite sex affects the family formation patterns of young adults after finishing their college education. The empirical analysis uses Spanish data, as one of the countries with strong field segregation in college and where high quality data are available on sex ratios by year, field and university. The evidence suggests that that sex ratios in college matter. Controlling for labor market

[1] A joint work with Libertad González.

conditions and individual and regional characteristics, and using field fixed-effects to account for the potential endogeneity in the choice of field, it is shown that individuals who are exposed to a larger fraction of opposite-sex school mates are more likely to be married or residing with a partner from the same field of study a few years after finishing school. In addition, the probability of starting a family shortly after school has been shown to increase with the degree of balance in the sex-composition of one's school class.

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1 Domestic Violence and Divorce Law: When Divorce Threats Become Credible

1.1 Introduction

Domestic violence is an important concern for many societies and policymakers worldwide. Statistics available for European countries show that between 20 and 25 percent of women have been victims of physical abuse at least once during their adult lives, and around 10 percent have suffered sexual abuse involving the use of force (CAHVIO, 2011). Estimates for the U.S. from the National Violence Against Women Survey show similar numbers: 1 out of 3 women surveyed reported having been raped or physically assaulted since the age of 18 years (Tjaden and Thoennes, 2000). Moreover, in most of the cases of violence against women, the crime is committed by the intimate partner. In this context, it is natural to ask about the relationship between domestic violence and family policies, and specifically, the rules governing the dissolution of marriages. In recent decades, many countries have adopted reforms aiming at simplifying the dissolution of marriage when one of the spouses wants to end the relationship. Since the early 1970s, many states in the U.S. removed fault as a ground for divorce, and almost all of them allowed one of the spouses to file a petition for divorce without the consent of the other. Many European countries have followed similar paths during the past 50 years.

Making divorce easier can affect the incidence of domestic violence, either by facilitating the dissolution of abusive relationships or by making the threat of leaving more credible, thus improving the situation of the victim within the marriage. Economic theories of household bargaining suggest that policies that affect spouses' well-being outside the marriage may also affect within-household distribution through changes in their relative bargaining position (McElroy and Horney, 1981; Lundberg and Pollak, 1993; Chiappori, 1988, 1992). In spite of the important link between domestic abuse and divorce legislation, the available empirical evidence in the economic literature is scarce and shows conflicting results (Dee, 2003; Stevenson and Wolfers, 2006). The relationship between divorce and domestic abuse has also captured the attention of the sociology and criminology literature. However, although alternative theories have been proposed to explain this relationship, empirical research in these fields has, in general, failed to provide credible causal estimates.

This paper studies how divorce law affects domestic violence. It begins by outlining a simple model of bargaining within the marriage to provide a framework for understanding the mechanisms through which easier divorce influences the incidence of spousal violence. The main prediction of the model is that a reduction in the cost of divorce improves the bargaining position of abused spouses by increasing their threat point (i.e. the minimum utility level required from the marriage to continue married), and this leads to a lower equilibrium level of spousal violence among intact couples.

To identify the causal effects, this paper exploits an unexpected and comprehensive reform of divorce legislation that took place in Spain in 2005. This reform allowed one spouse to file for divorce unilaterally and without the other spouse having committed fault, and eliminated the requirement of mandatory legal separation before divorce, thus reducing the length of time needed to effectively dissolve a marriage. The response of the divorce rate was immediate: In the first year after the reform, the number of divorces grew by 170 percent, and although this increase was partially compensated by the reduction in the number of judicial separations, the evidence points to an important rise in marital dissolution rates, at least in the short run. The empirical strategy takes advantage of the fact that the legal change suddenly and substantially reduced the cost of marital dissolution among the already-married couples, but did not affect the cost of terminating the relationship for unmarried partners, which provides an ideal setting for a difference-in-differences approach. Moreover, the fact that the effective reduction in the cost of divorce varies according to specific characteristics of couples offers additional sources of variation that strengthens the identification of causal effects. In particular, the effective decline in the length of the dissolution process, and consequently, in the cost of divorce, is limited by the presence of young children, in which case, there are decisions regarding custody and maintenance, which require more time.

This study considers a variety of measures of spousal conflict, ranging from self-reported spousal abuse in surveys and technical definitions of spousal violence based on recorded behavior, to more extreme measures of well-being such as partner homicide. The analysis of the impact on non-extreme measures of violence benefits from a large and rich survey on violence against women conducted in Spain, both before and after the legal change. Besides providing different measures of spousal conflict, these data have allowed knowing the respondent's marital status at the time of the legal change, avoiding concerns about selection issues. To study the impact on extreme spousal violence, data on female homicide by intimate partner between 2000 and 2010 have been used.

The main empirical findings point to a significant decline in spousal violence following the introduction of easier divorce. Self-reported abuse from intimate partner has fallen by about 27-36 percent among married couples, with respect to unmarried ones, as a consequence of the legal change. Similarly, technical definitions of intimate partner abuse based on recorded behavior have evidenced a reduction of about 31 percent. Moreover, the incidence of spousal violence has decreased among couples who remain married after the reform, which suggests an important role for changes in bargaining within the marriage when divorce becomes a more credible option. The evidence also suggests that there are important heterogeneous impacts arising from the reform. This study has found that married women without young children gain the most from the reduction in divorce costs, while the level of spousal abuse for mothers of young children has not changed significantly. Having young children seems to prevent women either from leaving an abusive relationship or from credibly threatening to do so. This study has also explored how the effects of the legal change vary with the value of opportunities outside marriage. The theoretical framework suggests that there is a level of the outside option at which a woman would be indifferent between filing for divorce and continuing in an abusive marriage, and that the impact of the reform should be larger around this margin.[1] When using education as an indicator of the outside option, this study has observed larger impacts for women at the center-bottom part of the skill distribution, which indicates that the woman at the margin of indifference has relatively low education.

The results also show a decline in extreme spousal violence, which can be attributed to the legal change. Intimate partner homicides of married women have fallen by around 30 percent after the reduction in the cost of divorce. Moreover, a relevant fraction of this decline is explained by a reduction in violence between spouses who are amid a process of marital dissolution. As other social sciences consider marital dissolution as a key determinant of conflict between separating spouses,[2] and as both the theoretical framework and evidence point to an increase in the share of conflicting divorces (i.e. cases in which one spouse prefers the continuation of the marriage), this result has important implications for the role of the duration of the divorce process. In particular, these findings provide evidence in favor of a negative association between the length of the divorce process and the incidence of ex-spouse victimization.

The literature on the effect of divorce law has focused on a variety of outcomes, such as divorce rates (Peters, 1986; Allen, 1992; Friedberg, 1998; Wolfers, 2006; González and Viitanen, 2009), marriage rates (Rasul, 2006), female labor supply (Gray, 1998; Stevenson, 2008), marriage-specific investments (Stevenson, 2007), fertility decisions (Drewianka, 2008; Alesina and Giuliano, 2006), and children's outcomes (Gruber, 2004). Less attention has

[1]The intuition is straightforward. Women with very poor alternatives outside marriage cannot take advantage of the lower cost of exiting the relationship, while women with very good outside option have a high and credible threat point, independent of the cost of divorce.

[2]See, for instance, Stolzenberg and D'Alessio (2007); Gillis (1996); Campbell (1992); Dugan, Nagin, and Rosenfeld (1999, 2003); Wilson and Daly (1992).

been paid to the effects of unilateral divorce on spousal violence. One exception is the study by Dee (2003), which exploited the variation stemming from the different timing of divorce law reform across states in the U.S. to assess the impact of unilateral divorce on the prevalence of lethal spousal violence. Using state-based panel data from 1968 to 1978, Dee found a small and statistically insignificant effect on the number of wives killed by their husbands, and large and statistically significant positive effects - of around 21 percent - on the number of husbands killed by their wives.[3] These results were revisited by Stevenson and Wolfers (2006), who, using the same data source but with a longer panel (1968-1994), found opposite effects on spousal homicide: No impacts on male homicide and a 10-percent decrease in female homicide. Beyond these discrepancies, other concerns made the findings of Stevenson and Wolfers (2006) less than definitive. One is the timing of the effects. As the authors acknowledge, the decline in female homicide predates the legal change to an extent that may undermine their results.[4] Moreover, those results are not robust to control for the changes in female homicide committed in unmarried partnerships, which should not be directly affected by the law change.[5]

In addition, an identification strategy based on variation across time and states could be problematic if both the legal definitions of divorce regimes and reforms introduced vary from one state to another (Mechoulan, 2005; Allen and Gallagher, 2007; Allen, 2007).[6] For instance, while many states passed unilateral and no-fault divorce law, some of them require a separa-

[3] Dee (2003) noticed that this effect is driven by states where the treatment of marital property favored husbands.

[4] Figure II (p. 285) of their paper clearly shows that the downward trend in female homicide started between 7 and 8 years before the adoption of unilateral divorce law.

[5] Using the same database and a similar specification, this study has found a 13-percent reduction in intimate partner female homicides among unmarried couples. These results are available upon request.

[6] Other potential problems of this identification strategy is the potential endogeneity in the timing of the adoption of reforms by different states, and the issue of "migratory divorce" - i.e. people choosing where to file a petition for divorce -(Allen and Gallagher, 2007).

tion period, while others do not. Also, those separation requirements may go from a few months to 2 years. In other states, changes in the grounds for divorce were accompanied by changes in property division, alimony, and custody rules. These differences matter. In fact, different coding of divorce regimes is one of the sources of the conflicting findings reached by previous empirical studies.

Stevenson and Wolfers (2006) also studied the impact on non-extreme domestic violence, and found that unilateral divorce law caused a reduction in around 30 percent in both female- and male-initiated conflict. The unfortunate timing of their data, however, prevented them from providing conclusive evidence. The first wave of their data was from 1976, when 31 states had already changed their divorce law, while the second wave was from 1985, when 6 more states had passed that reform. They considered these 37 states as treated and used two alternative control groups - the 9 states that already allowed unilateral divorce in their preexisting regime, and the 5 states that had not passed these reforms by 1985. Thus, their identification strategy relied on a differential evolution in domestic violence between the treatment and control groups, which was then attributed to the reforms. The main problem with this approach is that it may confound potentially different pre-existing trends in domestic violence between treatment and control states with the true effect of the policy change.[7]

Research in the sociology and criminology literature has also investigated the relationship between divorce and spousal violence. Some scholars support an "exposure reduction" approach, by which any mechanism that facilitates the dissolution of dysfunctional marriages should alleviate spousal violence by reducing the exposure of the victim to the offender (Dugan, Nagin, and Rosenfeld, 1999, 2003). Another line of research, however, states that a change towards less effective marital contracts may be ineffective to

[7]Other problems with these results are that 15 states were not sampled in the 1976 survey, and that the survey universe consisted only of intact marriages, which makes it impossible to disentangle the effect on domestic violence that occurs through a change in divorce propensity from the one related to changes in bargaining in intact relationships.

reduce domestic abuse if it continues between ex-spouses (Campbell, 1992), or even worse, it may intensify it if the abuser feels his or her dominant position is at stake (Wilson and Daly, 1992). Empirical research from these fields, in general, fails to prove causal relationships. For instance, Stolzenberg and D'Alessio (2007), by examining the cross-sectional relationship between divorce rates and domestic abuse in main U.S. cities, found that cities with higher divorce rates have higher levels of domestic crime between both spouses and ex-spouses. They argued that easier divorce does little to reduce the amount of domestic violence that occurs in a society, because after divorce, abuse continues between ex-spouses. However, they did not consider the potential reverse causality from domestic abuse to divorce rates. In a related study, Gillis (1996) used time-series data from 1852 to 1909 from France, and found a strong negative correlation between the rate of marital dissolution and female homicide. However, potential omitted variable bias prevented the author from claiming causation.

The present study's contribution to this literature is threefold. First, this study has employed a methodology that overcomes some of the shortcomings of previous research. This study has exploited an unexpected, large, and clearly defined change in divorce rules in Spain, where family law is mainly defined at the national level. Furthermore, other potentially relevant changes over the same time period have been accounted by using individuals not directly affected by the legal change (unmarried couples), to estimate the evolution in domestic violence in the absence of the reform. Second, the analysis of the impact on non-extreme violence is based on data from a large survey on violence against women conducted in Spain both before (1999 and 2002) and after (2006) the legal change. In addition, given that the survey universe consisted of all adult women living in Spain, independent of their marital status, this study could directly disentangle the two main channels through which easier divorce could affect domestic violence. Moreover, the richness of the individual-level data allowed us to go one step further than the previous research, by considering the potential heterogeneous impacts of the reform. The cost of divorce faced by an individual not only depends on the legal regime in place, but also on individual characteristics such as education and the presence of children, among the others. Third, in the analysis of the impact on extreme violence, this study has distinguished female homicide committed by spouses from those crimes involving ex-spouses. This distinction, so far neglected in the economics literature, is important because easier divorce could affect married and separated couples differently. The results obtained can be interpreted as supportive of an "exposure reduction" approach, because they point out the importance of the shortening of the length of the dissolution process as a key factor explaining the decline in lethal violence against ex-spouses. In this sense, this study also adds to the sociology and criminology literature.

The rest of the paper is structured as follows. Section 1.2 presents a simple theoretical framework for understanding the interaction between divorce law and spousal violence. Section 1.3 describes the main institutional context and the identification strategy. The data sources are described in Section 1.4. Section 1.5 presents the main empirical results and, finally, Section 1.6 provides the conclusion.

1.2 Theoretical Framework: Why easier divorce can affect domestic abuse.

This section presents a simple theoretical framework that attempts to shed light on the interaction between spousal violence and divorce costs. In this model, a marriage is seen as an institution that produces a valuable output which is distributed between spouses according to some predetermined shares.[8] After the marriage has taken place, spouses get to know the utility

[8]The marriage market is not explicitly modeled in this setup. Individuals are assumed to make their marriage decision on the base of the gains from the union and a certain distribution of those gains between them. They marry if their share of marriage gains is enough to compensate their utility from being single, otherwise they remain single. That distribution is based on some shares that may reflect their bargaining power in the marriage market. Factors such as sex ratios (Angrist, 2002) or legislation regarding

level they would obtain in case of divorce.[9] Utility upon divorce is considered as a threat point, since the continuation of the marriage will require that both spouses receive an utility level within marriage at least as high as what they would receive in case of divorce. A key assumption in this model is that those outside options remain private information for each spouse.[10] The model has two stages. In the first stage, each spouse observes the value of his or her own utility outside of marriage, and decides whether to continue married or to file for divorce. In absence of mutual consent for dissolution, the spouse seeking divorce has to pay a cost. In the second stage, conditional on the continuation of the marriage, they (re)negotiate about how to distribute the gains of the marriage. A bargaining process is explicitly modeled in this stage, which may involve the use of violence from the husband and may have divorce as a response from the wife.

a A simple model

Stage 1: Realization of payoffs and negotiation to continue the marriage

Individuals make their marriage decision on the base of the expected gains from the union, which are distributed between the partners according to some predetermined shares. Let us denote the utilities within the marriage as u_h and u_w for the husband and the wife, respectively. Once the marriage

property division after divorce (Chiappori, Fortin, and Lacroix, 2002) may influence individual's bargaining power in the marriage market. To make the marriage market endogenous to changes in divorce legislation is an important task left for future research.

[9] The assumption of ex-post information about the value of opportunities outside the relationship has been motivated in the literature by the arrival of new information during the marriage, such as the value of a potential new relationship or changes in the value of market opportunities (Becker, Landes, and Michael, 1977; Peters, 1986; Weiss and Willis, 1997).

[10] Zhylyevskyy (2008) and Friedberg and Stern (2010) provide empirical evidence coming from the National Survey of Family and Households supporting this assumption. They show that spouses have incorrect beliefs about the happiness or unhappiness of the other partner outside of marriage.

has taken place, each spouse observes his or her own outside opportunity, denoted by O_w and O_h, but they do not observe the outside opportunity of their partner. private information for each of them. Then, each spouse compares the utility levels he or she would receive in each of the states (marriage or divorce), and decides whether to propose the continuation of the marriage or to stand for divorce. When considering the possibility of divorce, each spouse takes into account that the lack of mutual consent for the termination of the marriage implies to pay a certain cost of divorce C.[11] On the contrary, if both spouses agree to a divorce, they do not have to pay any divorce cost.[12] As a result of individual assessments of utilities in each state, we would have three possible situations:

1. Both spouses prefer divorce: $u_w < O_w - C$ and $u_h < O_h - C$. Although they consider the divorce cost when making the decision, they get mutual consent for divorce and do not have to pay that cost.

2. Both want to continue married: $u_w > O_w$ and $u_h > O_h$. There is no conflict of interest and the marriage continues. Note that in this case the marriage yields a surplus $S = u_w + u_h - (O_w + O_h)$.[13]

3. Only one spouse wants to leave the marriage. Assume that the one

[11] There is no distinction in the model between mutual consent and unilateral divorce. The regime can be thought as unilateral since any spouse can make the decision of leaving the marriage without having the consent of the other, but incurring in a cost which is not present when there is mutual consent for termination. This setup is motivated in the actual divorce regime in Spain, which allows for unilateral separation based on certain grounds. These grounds include the usual considerations of fault or "de facto" separation, in which case, effective cessation of marital life for a period of 3 years is required. Having to prove fault in court or getting "de facto" separation is the cost that the spouse who wants to leave the marriage unilaterally has to pay.

[12] Notice that a condition for standing for divorce is to be able to afford it, that is: $u_i < O_i - C$. Otherwise, if $O_i - C < u_i < O_i$, the person would prefer to continue married. There will be of course cases in which both would like to get divorced but could not afford it if having to pay C. We can assume with no loss of generality that they would reach an agreement for mutual consent divorce.

[13] Cases in which only one spouse would like to leave the marriage but cannot pay the cost of divorce (i.e. $O_i - C < u_i < O_i$) are included here.

wanting to get divorce is the wife.[14] Two possibilities arise:

(a) She wants to leave and can afford it, but her husband can compensate her to stay together: $u_w < O_w - C$ and $u_h - O_h > O_w - C - u_w$. I assume that if compensation is possible, whatever the scheme of this compensation is, the compensation takes place and the marriage continues.[15]

(b) She wants to leave and can afford it, and her husband cannot compensate her to stay together: $u_w < O_w - C$ and $u_h - O_h < O_w - C - u_w$. Compensation is not feasible and there is divorce. Note that the husband would prefer to continue married but can not convince her to stay. The dissolution is unavoidable and conflict may arise with the decision of the wife of leaving the relationship.

Stage 2: (Re)negotiation of the distribution of gains of marriage

Conditional on the continuation of the marriage, spouses may (re)negotiate the distribution of the marital surplus. Marriages that survive the first stage are: (i) those in which both spouses want to stay married, (ii) those in which one spouse wanted to leave but was compensated by the other to stay together. To simplify, assume that renegotiation only takes place in case (i). Note that the surplus S is not known with certainty, since the outside option of the other partner is not observable.

Assume now that the husband can force a renegotiation of the distribution of the surplus in order to maximize his value of the marriage, and this renegotiation requires the use of violence. He can either choose violence (V) to claim a transfer ($T(V) = T$) from his wife, or choose no violence (NV)

[14] The case in which it is the husband instead, is completely symmetric.

[15] Note that the existence of divorce costs may imply that some inefficient marriages do not end up in divorce. This would be the case if $O_w + O_h - C < u_w + u_h < O_w + O_h$.

and remain with his original share of the surplus (that is, $T(NV) = 0$).[16] If he chooses violence and there is no divorce, his utility becomes $u_h + T$.[17]

The wife responds by deciding whether to stay in the marriage, accepting a lower share of the surplus (because of the transfer and the disutility from violence), or to file for divorce. If she stays, her utility is $u_w - T - V_w$, where V_w is the disutility from violence. If she divorces, her utility is given by $O_w - C$.

Wives differ in their outside option, such that $O_w \sim [O_w^{min}, O_w^{max}]$. We can interpret this as their labor market potential after divorce or their remarrying probabilities.[18] The husband does not know the true value of O_w, but only the distribution in the population.

The solution for the second stage of the game can be found by backward induction. The wife will choose between staying and leaving, given the decision of her husband. In the absence of violence her best strategy is to stay, given that $u_w > O_w - C$ for all women in this stage. If there is violence, she will divorce if and only if $O_w - C > u_w - V_w - T$. Otherwise, she will stay in the marriage and suffer violence from her husband.[19] The husband makes his decision about violence knowing only the probability that she will divorce if her utility inside marriage falls below the value of her outside option. To simplify the notation, let us call p the probability

[16] This transfer should be interpreted as any redistribution of the gains of the union in favor of the husband.

[17] This would imply that violence is "instrumental", in the sense that it is used as a means to get a higher share of S.

[18] In order to have that some wives would divorce in case of violence and others do not, we need to impose some restrictions to the distribution of O_w, such as: $O_w^{max} - C > u_w - V_w - T$ and $O_w^{min} - C < u_w - V_w - T$.

[19] Given the interpretation of the cost of divorce in the first stage, a natural question here would be why the wife has to pay a cost to get divorced, given that the husband has committed fault (violence). Nevertheless, we can think of C as the cost of having to prove violence in court, plus the period of mandatory separation that still should be incurred.

that she will divorce as a consequence of violence.[20] Then, he compares the extra utility he would receive if violence is accepted with the probability that she divorces and he is being left with his outside option. A condition for choosing violence, therefore, is that $(u_h + T)(1 - p) + O_h p > u_h$. If this inequality holds, the husband will choose violence; and the wife will stay in an abusive marriage with probability $1 - p = F_{Ow}[u_w - V_w - T + C]$, and divorce with probability $p = 1 - F_{Ow}[u_w - V_w - T + C]$.

b Comparative Statics and Implications

This simple model yields clear and intuitive predictions on the impact of a reduction in divorce costs on domestic abuse. The probability of domestic violence is: $F_{Ow}[u_w - V_w - T + C]$. This is increasing in the cost of divorce, C, which leads to the following proposition:

Proposition 1 *The prevalence of domestic abuse among married couples decreases after the reduction in the cost of divorce.*

It is important to notice that the reduction in domestic violence comes not only from the increase in dissolutions of abusive marriages, but also from a reduction in the incentives of husbands to choose violence. To see this more clearly, the condition for the husband to choose violence can be rewritten like this:

$$T(1 - p) > (u_h - O_h)p$$

If the reform changes the probability p, it also changes the incentives to choose violence in order to force a renegotiation of the surplus. The reform,

[20] A wife will leave an abusive marriage if she has a good enough outside option, that is: $p = Pr(O_w - C > u_w - V_w - T) = 1 - F_{Ow}[u_w - V_w - T + C]$, where F_{Ow} is the c.d.f. of Ow.

therefore, reduces the equilibrium level of domestic violence through an improvement of the bargaining position of the wife.

The model has also implications for the distribution of the effects in terms of individual characteristics. One of the main sources of heterogeneous responses to changes in divorce law is the presence and age of children. This argument can be rationalized at least in two ways. First, the reduction in the cost of divorce is larger for women without children under age 18. Having young children lengthen the divorce process since decisions about custody and maintenance payments have to be made. Second, the presence of young children has been found an important determinant of individual-specific cost of divorce (Del Boca and Flinn, 1995; Weiss and Willis, 1997). For instance, mothers of young children are likely to face higher emotional and economic costs of marital dissolution than non-mothers or mothers of older children. A simple extension of the model would be to assume that the cost of divorce for a certain woman, C_w, depends not only on the divorce cost determined by the current legal regime, say $\overline{C}$, but also on individual characteristics, c_w. While the reform in divorce law affects the general component of the cost of divorce, the presence of an specific component will lead to differences in the intensity of the treatment.

Corollary 1 *The reduction in the incidence of abuse is larger for women more affected by the reduction in divorce cost.*

A second source of heterogeneous responses to the law change are differences in women's outside option. The following corollary shows this:

Corollary 2 *The reduction in the incidence of abuse is larger for women with better outside options.*

Moreover, the assumption of a continuous distribution for women's outside opportunities leads to an interesting testable prediction: the reduction in domestic abuse should come not from women at the top end of that distribution, but from women with better outside of marriage prospects among those suffering abuse in the old regime. Figure 1.1 illustrates this point.

In this model, divorce happens in equilibrium not only as a response to spousal violence in the second stage but also as a consequence of realizations of outside options (net of divorce cost) in the first stage. The reduction in the cost of divorce, then, will affect the probability of divorce. In particular, it will increase the frequency of divorces in which one spouse prefers the continuation of the marriage.[21]

Proposition 2 *The rate of non-mutual consent divorce increases after the reform in divorce law.*

How is this related to spousal violence? As the sociology and criminology literature show, partner violence -and in particular, extreme violence- often occurs around important events in a relationship such as a unilateral breakup decision (Stolzenberg and D'Alessio, 2007; Wilson and Daly, 1992; Campbell, 1992). The reduction in the cost of divorce would lead to a higher demand for divorce and, in particular, a higher share of unilateral breakups, which could potentially lead to more conflict between separating spouses. At the same time, the legal change shortens the length of the whole dissolution process, from the decision to divorce until divorce is effectively obtained. Therefore, assuming that the highest risk of spousal violence occurs during the dissolution process, whether we should expect more or less violence between separating spouses would depend on which effect predominates: the increase in the number of spouses involved in conflicting dissolution or the reduction in the length of time required to dissolve the marriage. For instance, if N is the number of couples, d is the probability of divorce, h is the probability that a divorce ends in partner homicide during the divorce process (per unit of time at conflict), and t is the duration of

[21]The reduction in the cost of divorce makes it more difficult for the spouse who values the marriage more to compensate the other partner to stay.

that process, the number of partner homicides between separating spouses is: $N * d * t * h$. Suppose now that a reform of divorce law makes: $d_1 > d_0$ and $t_1 < t_0$, where 0 and 1 denote the pre- and post-reform period, respectively. Suppose also that N and h are unchanged by the reform.[22] Then, at a certain point in time during the new regime, the number of people at risk of spousal homicide will be lower if and only if $d_1/d_0 < t_0/t_1$.

1.3 Empirical Strategy

a The Reform of Divorce Legislation in Spain in 2005

In July 2005, the Spanish parliament approved a comprehensive reform of the rules governing marital dissolution in Spain.[23] This reform included two key modifications that substantially lowered the barriers to divorce. First, it eliminated the mandatory 1-year legal separation period before divorce.[24] Second, it allowed for unilateral and no-fault divorce.[25] As a consequence of these legal changes, the divorce regime suddenly went from one with fault and mandatory separation period to another with easy, unilateral, and no-fault divorce, dramatically reducing both the economic and emotional costs of marital dissolution.

The old regime, which was in place since 1981, was mainly characterized by a two-step process to deal with marital breakdown. The couple who wanted to dissolve the marriage generally had to resort to a period of separation

[22] Since h is the probability of spousal homicide during the divorce process per unit of time at conflict, it can be assumed as unchanged by the reform.

[23] Act 15/2005 of July 8th, modifying the Spanish Civil Code and the Civil Procedure Rules on matters of separation and divorce.

[24] Legal separation is left as an option for those not wanting to resort to divorce.

[25] Other modifications included the reduction of the waiting period after which it is possible to dissolve a union from 1 year to only 3 months since the celebration of the marriage, and the introduction of the notion of shared custody of children after divorce.

before being able to file for divorce.[26] Once the petition for legal separation had been filed, at least 1 year had to pass before filing for divorce. Separation, in turn, could be obtained by mutual consent or unilaterally, but based on a legal ground. The legal grounds for separation established in the Spanish Civil Code included the usual considerations of fault - unjustified abandonment of the family house, marital infidelity, abusive conduct, being sentenced, alcoholism, drugs addiction, etc. - or the effective cessation of marital life for a period of 3 years.[27]

The combination of unilateral and no-fault divorce with the possibility of filing for divorce directly, without legal separation as a necessary step, implied a substantial reduction in the length of time needed to obtain a divorce. Quantifying this time reduction is not an easy task, because it may depend on whether there was mutual consent for separation or not, and on the ground on which separation was based. A lower bound for this shortening of the process can be determined in 1 year, the period established in the old regime between the separation petition and the possibility of initiating the divorce process. Nevertheless, in some cases, this period can be much longer, particularly in those relationships in which there was no mutual consent for termination. The old regime made separation particularly difficult for a spouse who was unhappy in a relationship and wanted to leave without having the consent of the other partner. A person like this usually faced two alternatives. One was to go to court and claim separation on the base of fault, in case it existed, which may involve a lengthy and expensive legal battle with the other partner. A second alternative consisted of stopping marital life for a period of 3 years, and then claim legal separation on the base of *de facto* separation. In such a case, the change to unilat-

[26]There is one exception in which it is possible to directly file for divorce, which corresponds to the case in which there is risk of violence against the spouse or the children. For a more detailed description of the grounds for divorce in Spain before the reform of 2005 see Boele-Woelki, Braat, and Sumner (2003)

[27]This length corresponds to the case in which the cessation of marital life is not consented by both the spouses; otherwise, it would be reduced to only 6 months. However, this shorter period is somehow redundant, because mutual consent is a sufficient condition to file a petition for legal separation.

eral and direct divorce may imply a reduction to the dissolution process of about 4 years (3 years to file for legal separation on the ground of *de facto* separation plus 1 year before being able to file for divorce).[28]

As a consequence of the relaxation of the requirements to obtain a divorce, there was a huge increase in the number of divorce proceedings petitioned. Figure 1.2 shows the evolution of marital dissolution in Spain between 1975 and 2010. In the first year after the reform, the number of divorce petitions that entered into local courts increased by 170 percent. This was only partially compensated by a decline in separations, which can be explained by the fact that legal separations remain only as an option for those who do not want to opt for divorce directly.

Besides this increase in the number of marital dissolutions, the law change may have had a differential effect on women and men. For instance, in the old regime, women were more constrained than men to exit a relationship due to high costs of obtaining a divorce. The analysis of who is the spouse filing a petition for the dissolution of the marriage points to this direction. A separation or a divorce can be petitioned by one of the spouses or by both. Figure 1.3 shows the evolution of separations in which only one spouse has filed the petition, while Figure 1.4 shows the same for divorce proceedings.[29] In both cases, it is possible to observe an increase in the proportion of dissolutions initiated by wives after the reform, which provides evidence supporting the hypothesis that women are more benefited by the reduction in divorce costs.

[28] It is important to note that this is only an upper bound and, probably, in many cases, it would not be reached, even if it is not possible to prove that the other spouse has incurred any of the typified grounds for separation. This is because in those cases, courts usually refer to the so-called "lack of *affectio-maritalis*" as a valid ground for separation (Boele-Woelki, Braat, and Sumner, 2003).

[29] Both pictures are needed because during the old regime, marital ruptures were initiated with a demand for separation, while after the new regime, most of the dissolutions are obtained directly through divorce.

Other legal changes regarding domestic violence in Spain

During the period covered by this study, two integral plans and one main law aimed at preventing and combating domestic abuse were implemented. The First Action Plan against Gender Violence (1998-2001) and the Second Integral Plan against Domestic Violence (2002-2004) were elaborated and implemented by the Spanish Women's Institute. Those plans mainly included measures aimed at fostering awareness and prevention for potential victims, increasing the availability of resources for victims, and augmenting sanctions for aggressors. A major landmark in the fight against domestic violence, though, was the introduction in December 2004 of an integral law providing comprehensive protection measures against gender-based violence.[30] These measures can be grouped into three broad areas of intervention. The first consists of awareness-raising and prevention measures on the one hand, and education and training activities on the other hand. The main measures involve informational campaigns, raising-awareness advertising in the media, reinforcing the notion of equality of rights and opportunities between men and women in school curricula at all levels, training of healthcare professionals in detecting and preventing violence, and training of legal protection and support professionals. A second group includes penal and judicial measures such as increased penalties for gender-based offenses and the establishment of specialized courts to deal with this kind of crimes. Finally, the third group of measures aims at increasing protection for victims of gender violence.

b Identification Strategy

The identification strategy is essentially based on the reform in divorce legislation that took place in Spain in 2005, which can be considered as a source of exogenous variation in the rules of the game regarding marriage dissolution. As such, this reform constitutes a natural experiment and then provides a unique opportunity to identify the causal effect of easier divorce

[30] Organic Law 1/2004 of 28 December.

on domestic violence.

Two basic conditions should fulfill this legal change to constitute a valid natural experiment: being unanticipated and exogenous to the evolution of domestic violence. There are reasons to believe that these conditions are guaranteed. With respect to the first point, the reform in divorce legislation was part of a series of legislative measures concerning family law introduced by the Socialist Party right after winning the general elections in March 2004. The reason why these legal changes can be considered unexpected is that the election results themselves were totally unexpected. Until shortly before the national elections to the Spanish parliament were to take place, the incumbent party held a majority of public support according to available forecasts.[31] But a large-scale terrorist attack that hit the commuter train system in Madrid just 3 days before the date of the election suddenly changed the election outcome and resulted in a surprising victory of the opposition Socialist Party (Montalvo, 2010; Bali, 2007; Colomer, 2005; Chari, 2004).

With regard to the exogeneity of the legal change with respect to domestic violence, the stated purpose of the law was to give to the spouses the freedom to decide whether they want to continue married or not, and to eliminate the double procedure (fist separation and then divorce) usually needed to end a marriage, reducing both economic and emotional costs of marital disruption.

Then, the identification strategy used in this paper relies on a difference-in-differences approach (Angrist and Krueger, 1999; Heckman, Lalonde, and Smith, 1999), using married couples as the treatment group and cohabiting partners and individuals in a relationship but not legally married as a control group. That is, I compare the change in spousal violence for married women before and after the reform in divorce law, to the change in spousal violence for women not directly affected by the legal change (i.e. those in a relationship but not legally married). In this way, this empirical framework

[31] See for instance Center for Sociological Research (2004), Study 2559, April.

allows to control for systematic differences in the level of domestic violence both between married and unmarried women and between before and after the law change.

More formally, if Y_1 denotes the outcome of interest with treatment and Y_0 without, t' and t denote the pre- and post-treatment periods, and D is a binary indicator of program participation, the difference-in-differences estimator can be written as follows:[32]

$$\Delta^{DiD} = [Y_{1t} - Y_{0t'}|D = 1] - [Y_{0t} - Y_{0t'}|D = 0] \tag{1.1}$$

Since it is not possible to observe Y_1 and Y_0 for the same individual at the same time, this estimator relies on the following identifying assumption:

$$E[Y_{0t} - Y_{0t'}|D = 1] = E[Y_{0t} - Y_{0t'}|D = 0] \tag{1.2}$$

which is known as the common-trends assumption and requires that both the treatment and the control groups would have followed the same trend in the outcome variable, absent any reform. Under this assumption, it is possible to use the evolution of the population average difference over time in the control group as a benchmark to estimate the treatment effects. In terms of married and unmarried populations, this implies that the mean effect of the reform on spousal violence can be obtained as follows:

$$\begin{aligned}\Delta^{ATT} \quad = \quad & [E(Y_{it}|Married) - E(Y_{it'}|Married)] - \\ & [E(Y_{it}|Unmarried) - E(Y_{it'}|Unmarried)]\end{aligned} \tag{1.3}$$

where Y denotes some measure of spousal violence.

[32]Following the notation of Heckman, Lalonde, and Smith (1999)

Although there is no formal test to check the assumption of common trends between the treatment and control group, there are different ways to investigate its validity. The most straightforward is by graphically examining the data and comparing the trends of both groups in the pre-treatment period. An alternative test is to add controls for potentially different group-specific trends in the regressions and investigate whether there is enough evidence to reject the equal trends assumption. Both tests are carried out in the empirical analysis.

One potential threat to the validity of this assumption comes from aggregated shocks that have a differential impact across treatment and control groups. This may happen if the unobserved differences between both groups are correlated with those shocks. A potential candidate to constitute such a shock is the approval of the Law Against Gender Violence at the end of 2004. Nonetheless, most of the measures for protection against gender violence are aimed at all women, regardless of marital status. The only exception is given by measures aimed at facilitating separation and divorce procedures in cases in which domestic violence is alleged.[33] But even in the case that these measures have a differential impact between married and unmarried women, this effect would be intrinsically related to the main purpose of this paper, which is to assess the impact of easier divorce on the level of domestic violence.

Another key assumption of the difference-in-differences estimator is that there are no changes in the composition of the groups as a consequence of the reform. Otherwise, coefficients would be biased. To test the validity of this assumption, I use microdata from the census of marriages to evaluate two potential concerns in relation to the reform in divorce legislation. First,

[33] The Law Against Gender Violence of 2004 created specialized courtrooms to deal with gender violence crimes. When a criminal process is under the jurisdiction of these courts, they have also competence in civil law matters related to that process. This implies that separation or divorce procedures in which the women alleged spousal abuse are heard by these courts. Since 2005, when these specialized courts were created, around 4 percent of the total number of separations and divorces decreed in Spain fell under their jurisdiction.

I test whether there is evidence of a structural break in the time-series of marriages. Second, I check for potential changes in the composition of those who marry after the reform.

c Specifications

Non-extreme violence

The difference-in-differences approach translates into the following specification, in order to estimate the impact of easier divorce on non-extreme spousal violence:

$$DV_{igt} = \beta_0 + \beta_1 Married_g + \beta_2 (Married_g * Post_t) + \sum_t \lambda_t Year_t + X'_{igt}\gamma + \mu_{igt} \tag{1.4}$$

where DV_{igt} is a measure of domestic violence for individual i, marital group g, and year t, $Married_g$ is an indicator of the treatment group, $Post_t$ is a binary indicator for the post reform period and therefore β_2 is the difference-in-differences estimator.

Individuals affected by the legal change are those who were married or legally separated, but not yet divorced, when the law was passed. Given that the post-reform data were collected one year later, the definition of the treatment group should take into account potential transitions among marital states during this period, in order to avoid changes in the composition of groups. Available information about the duration of the relationship for intact marriages, and about elapsed time since the breakup for those who terminated, makes it possible to identify this group with precision. Then, the treatment group includes women who have been married for at least one year, or who are legally separated, or who have divorced during the previous year. Also, to ensure the comparability of the treatment group over time, the same definition is used for years 1999 and 2002.

There are two main measures of non-extreme domestic violence to be used as dependent variable. The first is a measure of self-reported abuse and is based on the interviewee's perception of having been victim of abuse from her intimate partner. The variable is defined as a binary indicator which takes value 1 if the woman reports abuse from intimate partner during the previous year. The second measure is called "technical abuse", since it is based on a series of 13 questions referred to behaviors or situations which are considered by experts as strong indicators of mistreatment. The survey contains information about the frequency with which these situations occur (i.e. frequently, sometimes, rarely, never) and about who is the offender. "Technical abuse" is a binary variable that takes value 1 if any of these 13 indicators occurs "frequently" or "sometimes" and the offender is the intimate partner of the victim. Also, this second measure can be disaggregated into four additional measures of abuse -physical, sexual, psychological in the form of control, and psychological in the form of emotional mistreatment, according to a classification elaborated by Alberdi and Matas (2002). In the tables below I consider these definitions of violence as alternative outcomes. The details of the construction of these measures as well as the description of the 13 indicators of abuse and the corresponding sampling frequencies are reported in Table A1.1.

These different measures of abuse lead to different sample definitions. On the one hand, when the dependent variable is self-reported abuse, since this information is available for all surveyed women, the sample includes all women who were in a relationship during the previous year. On the other hand, when the dependent variable is a measure of technical abuse, since that information is only available for women who are in a relationship at the moment of the survey, the sample is restricted to women who fulfil that condition.

Finally, the vector X_{igt} includes a rich set of control variables that can affect the level of domestic violence and also be correlated with marital status. It includes control variables for woman's age, education, labor market status, presence and number of children, religion beliefs, urban-rural residence, and region fixed effects. In some specifications, this vector also contains controls for education and labor market status of the partner.

Extreme violence

I estimate the following equation to capture the impact of the law change on female homicide:

$$\begin{aligned} FH_{gqt} &= \beta_0 + \beta_1 Married_{gqt} + \beta_2 (Married_{gqt} * Post_{qt}) + \\ &\quad \sum_q \gamma_q Quarter_q + \sum_t \lambda_t Year_t + \mu_{gqt} \end{aligned} \tag{1.5}$$

where FH_{gqt} refers to female homicides by intimate partner for group g, quarter q, and year t. In a first stage, the treatment group includes married and separated women, while the control is conformed of unmarried women. The reason to include both spouses and ex-spouses in the treatment group is that we are interested in the effect of easier divorce on spousal violence and we want to be sure that a potential effect on still married couples is not the consequence of the displacement of violence from married to separated couples.[34] In a second stage, I decompose the treatment group into two subgroups: Those victims who were still married and those who were already separated:

[34] An example may help to clarify this point. Suppose that the reduction in the length of time to obtain a divorce derived from the reform makes that an homicide, that otherwise would have occurred while the couple was still married, happens when they are already separated. In that case there is no reduction in spousal violence, but a displacement of violence from married to separated couples.

$$
\begin{aligned}
FH_{gqt} &= \beta_0 + \beta_1 Stillmarried_{gqt} + \beta_2 Separated_{gqt} + \\
&\beta_3(Stillmarried_{gqt} * Post_{qt}) + \beta_4(Separated_{gqt} * Post_{qt}) + \\
&\sum_q \gamma_q Quarter_q + \sum_t \lambda_t Year_t + \mu_{gqt} \qquad (1.6)
\end{aligned}
$$

The dependent variable is a measure of female homicides committed by intimate partner. It can be defined in at least three alternative ways, which lead to different econometric specifications. The first alternative, and probably the most natural, is to define it as a count. I use the aggregate number of intimate partner female homicides by marital status and quarter, for the period between 2000 and 2010. When the dependent variable is defined as a count, it is natural to assume it follows a Poisson process. Then, following the conventional parametrization of this kind of model, this implies that $ln(\lambda_{gqt}) = X'_{gqt}\beta$, where X'_{gqt} is a vector of explanatory variables and λ_{gqt} is the conditional mean of the number of homicides per group and period.[35]

An interesting property of Poisson regression models is that we can use individual or grouped data, with equivalent results. The only practical implication when using grouped data is that we need to include the logarithm of the population size for each group among the explanatory variables. On the other hand, one well-known limitation of Poisson models is the equidispersion property, by which the mean is equal to the variance (i.e. $E(FH_{gqt}) = var(FH_{gqt}) = \lambda_{gqt}$). This means that the usual assumption of homoscedasticity is not appropriate. The simplest way in which this concern can be addressed is by obtaining a robust estimate of the variance-covariance matrix of the estimator. Alternatively, the Negative Binomial regression model can be used, since it allows for overdispersed data (Cameron and Trivedi, 2005).

[35] This specification assumes that the number of homicides per group g and period of time given by q and t, FH_{gqt}, has a probability mass function equal to: $pr(FH_{gqt}) = \lambda_{gqt}^{FH_{gqt}} exp(\lambda_{gqt})/FH_{gqt}!$, for $\lambda > 0$.

A second alternative is to convert the count into a rate, by dividing the number of homicides by the corresponding group population size estimate, and estimate the model by OLS.[36] The choice of the functional form is not trivial. In fact, one of the reasons behind the conflicting results of past empirical studies is the use of different functional form. Then, investigating the stability of the results under different specifications is a way of assessing the robustness of those results.

The third alternative for the definition of the dependent variable consists of using the logarithm (instead of the level) of the homicide rate, in which case OLS is an appropriate model as well. The reason for this is that the homicide rate is always positive and therefore a linear model for the logarithm of the homicide rate is a more natural alternative (Lee and Solon, 2011).[37]

The main coefficient of interest is that of the interaction between the indicator of the treatment group and the dummy for the post reform period. This coefficient gives the average change over the post reform period in intimate partner homicide attributable to the law change.

In all cases I run the regressions with year and quarter fixed effects. In some cases, I also include linear group-specific time trends, in order to investigate the robustness of the results to the possibility that the common trends assumption fails.

Finally, there are two possibilities to define the beginning of the post reform period: to consider the date of announcement or the date of enactment. The law was approved by the Spanish parliament in July 2005 and was in force since that date, but was announced around 10 months earlier, when the first bill was approved by the Council of Ministers and submitted to the

[36] Population sizes for different marital groups can be obtained from the Spanish Labor Force Survey on a quarterly basis.

[37] Another possible alternative would be, given that the homicide rate is a fraction, to use linear models for the logit of the rate. See for instance Lee and Solon (2011) for a discussion on these issues applied to the impact of unilateral divorce on divorce rates.

Congress.[38] Since individuals may react to the introduction of new divorce regime right after its announcement, the post-reform dummy $Post_{qt}$ is set equal to 1 since the third quarter of 2004.[39] The empirical results shown in Section 1.5, however, are robust to using either date as the beginning of the post reform period.

Marital dissolution

Easier divorce can affect the incidence of domestic abuse by easing the dissolution of abusive relationships. Therefore, to complete the empirical analysis we need to assess the impact of the law change on marital dissolution. Evaluating this by looking at divorce rates directly is problematic, since the nature of the reform makes the before-after comparison meaningless.[40] To overcome this, I assess the impact of the reform on marital dissolution indirectly, by looking at the evolution of the stock of divorcees. The share of divorcees in the population at a point in time depends on both

[38]The whole process of approval of the legal change was actively followed by the media. To the best of my knowledge, the first newspaper article anticipating the reform to be introduced appeared on August 17th, 2004, in *El Mundo* newspaper (http://www.elmundo.es/elmundo/2004/08/17/espana/1092742690.html). After the Council of Ministers passed the first bill, it was first approved by the Congress of Deputies in April 2005, and later by the Senate in June 2005. The final enactment day was July 8th, 2005.

[39]The hypothesis that individuals became aware of the new policy around its announcement is supported by evidence provided by the search intensity on the internet for information about the legal change. This is shown by Figure A1.2, which depicts the evolution of the search intensity for the query *divorcio* -the Spanish word for divorce- in the search engine Google. There were two peaks in the search intensity for this query, coinciding with the announcement and enactment dates of the legal change. These data can be obtained at http://www.google.com/insights/search.

[40]Comparing divorce rates before and after would be misleading if we want to extract conclusions about the level of marital dissolution since some divorces after the reform simply substitute what otherwise would have been a separation. Comparing the total number of dissolutions (separation plus divorces) does not help either, since before the reform both were (in most of the cases) required to dissolve a unique marriage, while afterwards they could represent two different dissolution processes.

the propensity to divorce (the flow into divorce state) and the probability of remarrying (the flow out from divorce state). Then, abstracting from changes in remarriage rates, the evolution of the stock of divorced individuals can shed light on the impact of the reform on divorce probability.

To perform the analysis, I rely on data from the Spanish Labor Force Survey, which allows to construct fairly precise estimates of population size by marital status, on a quarterly basis. I use the stock of separated and divorced individuals -for simplicity I refer to this group as to divorcees- to estimate the following equation:[41]

$$\begin{aligned} divorcee_{it} &= \beta_0 + \beta_1 time_t + \beta_2 post_t^{2005} + \beta_3 timepost_t + \\ & X_{it}'\gamma + \sum_q \lambda_q Quarter_q + \mu_{it} \end{aligned} \tag{1.7}$$

where $divorcee_{it}$ is a dummy variable set equal to 1 if individual i is separated or divorced at time t, $time$ is a continuous variable indicating time in quarters from the start of the observation period, $post^{2005}$ is a dummy that equals 1 since the third quarter of 2005, when the reform in divorce legislation became effective, and $timepost$ is a continuous variable counting the number of periods after the law change. This flexible specification allows the stock of divorcees to trend linearly with potentially different slopes before and after the reform, and to have a change in level that can be attributed to the reform. That is, β_2 estimates the level change in the stock of divorcees immediately after the reform, while β_3 estimates the change in the trend in the mean number of divorcees after the reform. The vector of control variables, X_{it}', includes dummies for age and education groups, and also a dummy for gender when both men and women are included in the

[41] The survey does not distinguish between separated and divorced individuals, but this is not a problem, since both are a measure of marital dissolution. The main difference between the two cases is that divorce implies the termination of the marriage, while separation does not, since during this period reconciliation is still possible.

sample. Since I use quarterly data, quarter fixed effects are also added to control for seasonality.

Marital formation

The validity of the difference-in-differences approach proposed to estimate the impact of easier divorce on domestic violence requires that the reform neither affected the propensity to marry in the population nor the composition of those who marry. I test to what extent these two assumptions are supported by the data by using data on marriage records.

First, to investigate the possibility of a structural break in the series of marriages after the reform in divorce law, I estimate the following model using monthly data:

$$\begin{aligned} marriages_t &= \beta_0 + \beta_1 time_t + \beta_2 post_t^{2005} + \beta_3 timepost_t + \\ &\quad \beta_4 marriages_{t-1} + \beta_5 GDPgrowth_{t-12} + \\ &\quad \sum_t \lambda_t month_t + \mu_t \end{aligned} \tag{1.8}$$

where the dependent variable, $marriages_t$, is the number of new marriages in month t, $time$ is a continuous variable indexing the month; $post^{2005}$ is the usual indicator for the post-reform period, and $timepost$ is a continuous variable indicating time since the introduction of the reform. Variables $marriages_{t-1}$ and $GDPgrowth_{t-12}$ are included to control for autocorrelation and for the influence of economic conditions on the propensity to marry.[42]

Second, to investigate potential changes in the composition of new couples, I estimate the following equation:

[42] Including other lags of these two variables does not change the results significantly.

$$char_{it} = \beta_0 + \beta_1 time_t + \beta_2 post_t^{2005} + \beta_3 timepost_t + \sum_t \lambda_t month_t + \mu_{it} \quad (1.9)$$

where $char_{it}$ is a dummy variable that takes the value 1 if the individual i who gets married in month t has a particular observable characteristic and 0 otherwise, and $post2005_t$ is set equal to 1 since July 2005. The observable characteristics considered are spouses' main occupation, age at marriage, and previous legal civil status. As before, *time* and *timepost* are two continuous variables indicating time in months at time t, the first counting from the start of the observation period and the second from the enactment of the reform.

1.4 Data and Descriptives

a Databases

I employ two main databases to conduct the empirical analysis: a nationally representative survey on violence against women, and the official registry of female homicides by intimate partners.

Survey on Violence Against Women

To study the effects on non-extreme violence, I rely on microdata from the Survey on Violence Against Women conducted by the Spanish Women's Institute in 1999, 2002, and 2006. This survey is representative of all adult women (age 18 or older) living in Spain, irrespective of whether they are in a relationship or not.

The survey contains specific questions on abuse which make it possible to construct the measures of self-reported as well as technical abuse mentioned before. Respondents to the survey were queried about whether they think they have been victims of abuse from their intimate partner during the previous year and at any time in their adult life. They were also asked detailed questions about a series of situations considered indicators of violence, the frequency of this happening, and their relationship to the perpetrator.

The questionnaire also included detailed questions regarding the partnership status of the respondent, which allows to distinguish up to seven different marital groups: married, cohabiting, legally separated, divorced, widow, dating, and single. There is also information on the duration of the relationship. In addition, the survey also provides information -both for the woman and for her partner in case she has one- on demographic characteristics, labor market status, educational background, and household composition.

Data on female homicide by intimate partner

To study the impact of the reform on lethal spousal violence, I use data on female homicides by their intimate partner for the period between 2000 and 2010. Intimate partners include current and former husbands, opposite-sex cohabiting partners, boyfriends, and dates. There are two different sources for these data. The Spanish Women's Institute, an autonomous body attached to the Ministry of Health, Social Policy, and Equality; provides information on the annual number of fatal victims of intimate partner violence, disaggregated by victim-perpetrator relationship.[43] The Queen Sofia Center (QSC hereafter), a non-governmental institution devoted to the study of violence, provides similar information but on a monthly basis and with more details about the crime.[44] Besides knowing the victim-perpetrator relationship, QSC's data provide information about age for both of them,

[43]Their sources of data are the media and the Ministry of the Interior for 2000-2005, and the Government Office on Gender-based Violence for 2006-2010.

[44]Data come from the Ministry of the Interior, the media, and the courts responsible for handling cases.

place of residence of the victim, place where the crime was committed, and the motherhood status of the victim. For women who were legally married at the moment of the homicide, there is also information on whether the they had initiated the procedure to obtain legal separation. Because of the more detailed information and the possibility of defining the pre- and post-reform period with precision given the availability of data on a monthly basis, most of the empirical analysis below is based on QSC's information.

One limitation of both databases is that they do not distinguish between legally separated and already divorced victims in cases in which the perpetrator is the former spouse. In those cases, the victim-perpetrator relationship is coded as "ex-spouse". The importance of that differentiation is that while separated partners are affected by the legal reform (i.e. their dissolution process is subject to the new regime), those already divorced are not. This shortcoming of the data, however, appears to have little practical relevance. Both information contained in cases' description in QSC's data and anecdotal evidence seem to point to a majority of those cases corresponding to parters amid a process of separation and, therefore, not yet legally divorced.

Other sources of data

Besides these two main sources of data, I also employ other datasets to supplement the analysis. I use administrative data from Judiciary Statistics to study the evolution of the annual number of separations and divorces. Also, I employ microdata from marriage records provided by the Spanish Institute of Statistics to analyze the potential impact of the reform on both the quantity and the composition of new married couples. Finally, I use the Spanish Labor Force Survey, also conducted by the Spanish Institute of Statistics, to study the effect of the reform on the size and composition of the stock of divorcees. Population data employed to construct homicide rates is also obtained from this survey.

b Sample Definition and Descriptive Statistics

This section presents the basic features of the data used in the empirical analysis.

Non-Extreme Violence: Self-reported and Technical Abuse

The sample for the analysis of the impact of divorce law on non-extreme abuse consists of the waves of 1999, 2002, and 2006 of the Survey of Violence Against Women. Table 1.1 presents the main descriptives statistics of the data. The number of observations is 20.552 in 1999, 20.652 in 2002, and 32.426 in 2006. Important for the validity of the difference-in-differences approach with repeated cross-sectional data is that samples come from the same population. This seems to be the case when we observe the sample composition in terms of the main observed characteristics (Table 1.1).

It is interesting to see how the different measures of intimate partner abuse relate to each other. As expected, all correlation coefficients are positive and statistically significant. The coefficient for the correlation between self-reported and technical abuse is 0.326. Moreover, according to the correlation between self-reported abuse and the four types of violence in which technical abuse can be decomposed, it is possible to deduce that women who declare to be victims of abuse tend to associate this situation to physical abuse ($\rho = 0.464$), more than to psychological (emotional) abuse ($\rho = 0.374$), psychological abuse in the form of control ($\rho = 0.322$), or sexual abuse ($\rho = 0.153$).

The key assumption for the validity of the identification strategy (i.e. common trends) can be investigated by observation of the data. Figure 1.5 shows the proportion of married and unmarried women who reported to have been victims of abuse from intimate partner during the previous year in 1999, 2002, and 2006. Meanwhile, Figure 1.6 depicts the evolution of technical abuse by marital relationship during the same years. Although having only two data points during the pre-treatment period may be insufficient to convincingly prove the validity of the common trend assumption, the evidence available points in that direction.

Extreme Violence: Female Homicide

The sample for the analysis of extreme violence includes all 703 female homicides committed by intimate partners between 2000 and 2010 in Spain. During this period, the average number of female homicides per quarter is 16, with a minimum of 10 and a maximum of 24 (Table 1.2). In terms of the female population in Spain between 2000 and 2010, this translates into a quarterly prevalence of 0.88 female homicides per million women, or equivalently, 3.5 female homicides per million women and year.

According to the victim-offender relationship, in a typical quarter between 2000 and 2010, were killed in Spain 7.8 unmarried women, 6.1 married women, and 2.1 separated women.[45]

Figure 1.7 provides some evidence in favor of the common trends assumption for the difference-in-differences approach employed here. It shows the evolution of the number of intimate partner female homicides by marital group for the period between 2000 and 2010. Both the level and the year-to-year variation of the number of homicides are relatively similar for both treatment and control group, particularly in the years close to the legal change (2002-2004).

[45] As mentioned in Section 1.4, it is not possible to distinguish between separated and already divorced victims, since in both cases the victim-perpetrator relationship code is the same (i.e. "ex-spouse"). From now on, then, I refer to those cases as "separated" victims.

1.5 Empirical Results

a Non-Extreme Violence

Table 1.3 shows the results of equation 1.4 when the dependent variable is the dummy for self-reported abuse. Column 1 presents the results for a specification with no controls beyond the treatment indicator and year dummies. The difference-in-differences coefficient suggests a decline in self-reported abuse for the treatment group in comparison with the control group after the reform in divorce law by 0.75 percentage points. In column 2, I add individual-level controls -age, education, labor market status, legal civil status, presence and number of children, immigration status, and religion beliefs, while in column 3 I also include region fixed effects and a dummy for urban residence. After controlling for individual characteristics and aggregated variables, the estimated coefficient remains negative and statistically significant. In the preferred specification (column 3), easier divorce reduces self-reported abuse by 0.65 percentage points (29 percent of the sample mean). If we want to control for partner's education and labor market status, we need to restrict the sample to women with a partner at the moment of the interview.[46] This is reported in column 4, which shows that self-reported abuse decreases by 0.59 percentage points (27 percent of the sample average).

The estimate reported in column 3 reflects the impact of easier divorce on domestic violence through the two possible channels: the dissolution of abusive marriages and the decrease in violence among intact households. In order to capture the change in domestic violence explained by a change in wife's bargaining position within the household, column 6 reports the results when the treatment group is restricted to women who were already married when the law was passed and continued married at the moment of the survey. The coefficient not only remains negative and precisely es-

[46] While self-reported abuse refers to previous year, partner's information is only available for those women who declare to have a partner at the moment of the survey.

timated, but is also larger (equivalent to a reduction of 36 percent of the sample mean) than the estimate for the total effect of the legal change. This implies that the bulk of the decline in domestic abuse when the obstacles to divorce are lowered is explained by a decreasing propensity towards partner abuse within intact households. Lowering the barriers to divorce seems to act as a strong deterrent to spousal violence.

Finally, to test the robustness of these results, column 7 reports the results of a placebo test. In this case, the dependent variable is a dummy set equal to 1 if the person declares to have been victim of abuse at any point in life before -but not during- the last 12 months. This is a measure of self-reported abuse in a period that precedes the legal change and, consequently, should be unaffected by the reform. The result confirms this hypothesis. The coefficient is statistically insignificant and relatively low in magnitude, basically indicating no effect of the legal change on past abuse, as we would have expected.

The second measure of non-extreme violence is the indicator technical abuse as defined in Subsection 1.3 c. These results are shown in the first three columns of Table 1.4, which differ in terms of the control variables included in the regressions. The preferred specification, presented in column 3, controls for individual characteristics of the woman and her partner, year and region fixed effects, and urban-rural residence. The difference-in-differences coefficient indicates a reduction of 3.26 percentage points in the incidence of technical abuse (about 31 percent of the sample mean) since the introduction of easier divorce. The remaining columns show the results for the four different categories of abuse in which technical abuse can be disaggregated, according to Alberdi and Matas (2002). These results provide evidence confirming the main conclusion of a negative impact of easier divorce on domestic violence. In almost all cases the difference-in-differences coefficient is negative and precisely estimated.[47]

[47]The exception is the case of psychological abuse in the form of control, which is only statistically different from zero at a significance level of 12 percent or higher.

To test the robustness of these findings, Table 1.5 reports the results of using alternative definitions of technical abuse. So far, a person is considered technically abused if any of the 13 indicators of abuse available in the survey is present. Alternatively, all these indicators can be combined into one variable which reflects not only the existence of spousal violence but also its intensity. The first two columns of this table report the results of using this alternative measure as dependent variable. In column 2 the model is fitted by OLS, while in column 3 the count nature of the variable is taken into account and a poisson regression model is used to derive the results. In both cases the estimated coefficient is negative and statistically significant. In columns 3-7, the dependent variable becomes a binary indicator again, but now reflects different levels of spousal conflict. It is defined as a dummy that takes the value 1 if at least a certain number n of indicators of abuse are present, for $n = 2, ..., 6$. In all cases the estimated effect remains negative and strongly significant, confirming the decline in spousal violence after the introduction of easier divorce found before.

In sum, the evolution of the main measures of abuse over time and across groups points to both a statistically significant and economically relevant decline in domestic violence after the introduction of easier divorce.

b Heterogeneity of Impacts

The availability of individual-level data allows me to go one step further and test whether the effects of the reform vary across different types of women. I consider two sources of heterogeneous impacts of the legal change on non-extreme violence: the presence of young children, and education level of the woman.

The presence of young children and the intensity of the treatment

There are at least two reasons why we expect the effects of the reduction in the cost of divorce to vary across women depending on their motherhood status. First, the effective reduction in the length of time needed to dissolve a marriage is smaller when there are children below the age of majority (18 years), since decisions about child custody and maintenance payments slow the process. Second, the cost of divorce depends on individual-specific factors, besides the legal environment. The literature on family economics has identified the number and age of children as one of the main determinants of the cost of divorce among married couples (Becker, Landes, and Michael, 1977; Weiss and Willis, 1997). Parents of young children, for instance, may suffer more after divorce if it results in under-investment on their children (Del Boca and Flinn, 1995; Weiss and Willis, 1997). This implies that the intensity of the treatment varies across women and this can be used to test the consistency of the results obtained when looking at the average effect. If the reduction in the cost of divorce is less important for mothers of young children than for either non-mothers or mothers of older children, we would expect smaller reductions in domestic violence for the former than for the latter.

I consider the presence of children under 18 years of age who leave in the parental house as one of the main sources of differences in treatment intensity. The results of this exercise are reported in Table 1.6. Panel A of the table shows the results when the sample is restricted to women with young children, while Panel B does the same for women either without children or with older children not living with them. These results clearly show that the decline in domestic violence, measured both in terms of self-reported and technical abuse, is driven by the effects on women without young children at home. Difference-in-differences estimates for mothers of young children are not statistically different from zero in any of the measures of abuse considered in the analysis. On the contrary, those estimates are negative and precisely calculated in the case of women without young children at home. Not having young children, then, seems to be a necessary condition to take advantage of the reform in divorce legislation.

To test the robustness of these results, Table 1.7 presents an analysis based on a different identification strategy. Instead of using unmarried women as a benchmark for the no policy evolution of domestic abuse, this specification focuses on treated individuals and exploits differences in the intensity of the treatment. The sample consists of women who were married when the new divorce law became effective, and the differential effect of the reform on women without young children is captured by an interaction term between a dummy variable that takes the value 1 for those women without young children and the post-reform indicator. The parameter estimate for this interaction term is negative and statistically significant, independently of the measure of abuse considered, which suggests that the level of violence decreases more after the legal change among married women without young children. These results confirm the previous findings of a larger effect of the reduction in divorce cost on married women who did not have young children when the legal change was enacted.

Education as a measure of wives' outside option

A second reason why the effects of the reform may vary across women is that they differ in the value of their outside opportunities. In principle, married women with good prospects outside of marriage are less likely to remain in abusive relationships, even when the cost of divorce is high. Then, a reduction in the cost of divorce would lead to little change in the incidence of violence among those women. Women with poor alternatives outside of marriage, on the other hand, are less likely to benefit from a decrease in the cost of divorce, since they still would be better off in an abusive relationship than with divorce. Therefore, we would expect the effects of the law change to be larger, the closer is an abused wife to the margin of indifference between continuing in an abusive marriage or getting divorced.

One possible indicator of the value of the outside option for married women is their educational level. Table 1.8 presents the results when the total sample is disaggregated according to women's educational level. Panels A, B, and C, present the main coefficients for women with low (primary school or less), intermediate (high school), and high (university) education, while the dependent variables are self-reported abuse (column 1) and technical abuse (column 2). The parameter estimates indicate that the reduction in divorce cost is associated with a decrease in domestic violence among married women, with respect to unmarried women, although these coefficients are only estimated with precision in the cases of low and intermediate education groups. In other words, the level of domestic violence among married women, with respect to unmarried ones, only decreases toward the center and bottom part of the distribution of skills.

To test whether the effects are statistically different along those segments of the skill distribution, Table 1.9 reports the results for the full sample. The impact of the reform on low educated women, the omitted category, is captured by the interaction between variables *married* and *post*, while the differential effects on more educated women are captured by further interactions with binary indicators for intermediate and high education. For self-reported abuse (column 1), the parameter estimates suggest that the reform leads to a reduction in domestic violence that does not vary significantly across skills. For technical abuse (column 2), the effects of the reform are larger for low-skilled women than for intermediate- and high-skilled ones. Overall, these estimates suggest that the reduction in the cost of divorce results in a decline in domestic violence across all educational levels, and that this decline is larger for women at the bottom part of the distribution of skills, in particular when technical abuse is used as a measure of domestic violence.

Again, to test the robustness of these results, we can investigate how the impact of the reform varies among married women with different educational levels. Table 1.10 reports the results of a regression on the sub-sample of women who were married at the time of approval of the legal change. The post-reform variable captures the change in the level of violence for a married women with low educational level, while the interactions with the binary indicators for the other skill levels capture the differential effects for women with those skills. In columns 1 and 2, the sample includes married women of all ages, and the results point to a similar conclusion to the one obtained when unmarried women were used as a control group: the incidence on abuse decreases along the whole distribution of skills, and the reduction is larger among low-skilled women when the technical definition of abuse is used. In columns 3 and 4 of the same table, the sample is restricted to middle age women (i.e. between 30 and 50 years of age), to investigate the distribution of the impacts on a sub-group for which the education level may be a more appropriate measure of opportunities outside of marriage. Doing this exercise leads to a slightly different result. In the case of self-reported abuse (column 3), the only sub-group that benefits from the reduction in the cost of divorce is the one of women with intermediate education. Neither for low-skilled nor for high-skilled women there is a significant change in the incidence of spousal abuse. In the case of technical abuse (column 4), the estimated effect on domestic abuse is negative for women with low education, and although the impact seems to be larger around the center of the skill distribution, the difference is not statistically significant.

c Extreme Violence

This section presents the main findings for the impact of the reform in divorce legislation on female homicide. Column 1 of Table 1.11 shows the results for the estimation of equation 1.5 when the dependent variable is a count of all homicides committed by intimate partners per quarter. The coefficient of interest, the one of the interaction between the dummy variable for being married and the indicator for the post-reform period, is negative and statistically significant, indicating a negative effect on the probability of extreme violence. The magnitude of the coefficient reflects also a quantitatively relevant effect: a change of -0.326 in the log count translates into a reduction of 2.4 female homicides per quarter that can be attributed to the reform. With an average of 7.97 female homicides per quarter and group during the whole sample period, this is equivalent to a decline in spousal murder of about 30 percent.

Poisson models rely on the assumption of equidispersion (i.e. mean equal variance), which means that this model would not be appropriate had we found some signs of overdispersion. Nevertheless, several reasons justify the use of the poisson model. First, the distribution of the count of homicides does not show signs of overdispersion. The mean and the variance of the number of homicides per quarter and marital group are relatively similar: 7.97 and 6.36, respectively.[48] Second, the goodness-of-fit chi-squared test yields a statistics of 46.26 which leads to no rejection of the poisson model. Third, the likelihood ratio test of $\alpha = 0$ shows that α is not significantly different from zero, reinforcing the validity of the poisson model.

Column 3 presents the results when the dependent variable is specified as a rate and the model is fitted by OLS. The difference-in-differences estimator is -0.526 and it is estimated with precision. Considering a female homicide rate of 1.364 every million women per quarter during the pre-reform period, this estimate implies a reduction of about 46 percent after the law.

Finally, column 5 shows the results for the logarithm of the quarterly homicide rate as dependent variable. The coefficient is again negative and statistically significant, reinforcing the conclusion that the reform in divorce law had a negative effect on female homicides of married women by intimate partners.

One key assumption of the difference-in-differences approach is that the trend in the outcome variable for both treatment and control groups would have been the same, had the reform not been passed. To test the validity of this assumption, I include a group-specific linear trend in the regressions. This allows me not only to analyze which is the effect on the coefficients of assuming different trends for treatment and control groups, but also to test whether there is enough evidence against the assumption of common trends. The results are shown in even columns of Table 1.11. With respect

[48] The same conclusion is reached looking at the mean and variance of the total homicide count per quarter (i.e. without differentiation among marital groups), which is shown in Table 1.2.

to the change in the coefficients, as a result of including group-specific linear trends, they are estimated with less precision and are between one-third and one-half smaller in magnitude. But those results bring evidence supporting the common trend assumption. As can be seen from the statistics provided at the bottom of the table, in none of the 3 specifications we can reject the hypothesis of common trends between the treatment and control group.

One contribution of this study is to distinguish between female homicides committed by spouses from those crimes involving ex-spouses. This aspect of the relationship between divorce law and domestic violence has not yet been treated in the economic literature, even though many studies from criminology and sociology have pointed out the importance of marital disruption itself as a determinant of domestic violence between separating spouses (Campbell, 1992; Wilson and Daly, 1992; Stolzenberg and D'Alessio, 2007). Easier access to divorce would be ineffective to reduce the incidence of domestic violence in a society if, after marital dissolution, violence continues between ex-spouses (Campbell, 1992) or, even worse, if it escalates after the victim seeks a separation (Wilson and Daly, 1992).

This theoretical possibility is contemplated in the simple model developed in Section 1.2. It not only predicts an increase in marital dissolution after a reduction in the barriers to obtain a divorce, but also an increase in the share of dissolutions in which one spouse is unhappy with the termination of the marriage. This prediction is supported by empirical evidence on the evolution of the share of adversarial dissolutions (Figure 1.8).

Results presented in Table 1.12 test this possibility. The definition of the treatment group distinguishes between victims who were still married and those already separated or in the process of separation at the moment of the homicide. The difference-in-differences coefficient for separated victims is always negative and strongly significant, and larger in magnitude than the estimate for still married women. These findings not only reject the hypothesis of increased or continued violence during or after dissolution, but also suggest that an important portion of the reduction in female homicide as a consequence of the reform in divorce law comes from the reduction in violence against women who are amid a process of marital dissolution.

The strong negative effect of the law change on violence against ex-spouses also suggests an important role for the duration of the dissolution process as a key factor behind the reduction in violence. To see this, we need to rely on a series of assumptions regarding the link between divorce and spousal violence. First, that there is a positive probability that a conflicting dissolution ends up in extreme partner violence, such as homicide. Second, that this probability falls substantially, say to zero, once the dissolution process finishes and the divorce decree is issued. Third, that this probability -per unit of time at conflict- is unchanged by the reform. Then, a decreasing propensity toward spousal homicide among separating couples, in combination with a larger population dissolving their marriages with some degree of conflict, can be explained by a reduction in the length of time that those potential victims are at risk of extreme violence, that more than compensates the increase in the size of this population at risk.

d Effects on Marriage and Divorce

Divorce

Figure 1.9 shows the evolution over time of the stock of divorcees. Casual observation of the trends for both women and men indicates an acceleration of the growth rate of the stock after the reform in divorce law. Table 1.13 presents the results of fitting equation 1.7 by OLS, for the whole sample (column 1) and for women and men separately (column 2 and 3, respectively). In all cases the results of the modification in divorce law on the stock of divorcees follow a similar pattern. There is a statistically significant and positive impact on the population of divorcees immediately after the reform was in place, and a positive although statistically insignificant change in the trend. With regard to the magnitude of the impact, the rise in the number of divorcees was about 3.5 every 1000 people, when both women and men are included in the sample. With an average of 42 divorcees per thousand population before the approval of the reform, this impact translates to an immediate increase of about 7.6 percent that can be explained by the reform in divorce law. Similar conclusions can be derived for both women and men when they are considered separately.

Besides this very short run impact, it interesting to analyze which is the absolute effect of the law change on the size of divorced population, some time after the reform. The estimate of the absolute effect of the law change can be calculated as $\hat{\beta}_2 + \hat{\beta}_3 * timepost$, where $timepost$ refers to the period of time since the reform.This effect is shown in the bottom part of the table, for a period of 3 years after the reform in divorce law.[49] Three observations are worth noticing. First, the absolute impact on the stock of divorcees is positive and statistically significant both for women and men. Second, the magnitude of those impacts is relatively low if we consider the large increase in divorce rates shown in Figure 1.2. The size of the total effect is calculated with respect to the stock of divorcees after the law change, had the reform not been implemented.[50] This leads to an increase of about 2 percent in the stock of divorcees with respect to the counterfactual value had the reform not been implemented.[51] Third, the size of the increase is higher for men than for women (i.e. 3.3 percent versus 1 percent, respectively), which can be explained by the differential size of the stock of divorcees by gender (i.e. there are always fewer divorced men than women, as the propensity towards remarrying is higher for men).

I also test whether the legal change affected the composition of the stock

[49]The period of 3 years is chosen to approximately fit the mid-point of the post-reform period in the estimation of the impact on female homicide.

[50]Given that the outcome of interest has an increasing trend, comparing the total effect after the reform to the mean value before the reform would lead to misleading results (i.e. we would be overestimating the true impact).

[51]There are two possible explanations for this relatively low effect: the flow out of divorce because of remarrying, and the fact that the data do not distinguish between separated and divorced individuals, and therefore, do not capture transitions from separation to divorce.

of divorcees in terms of observable characteristics. Table A1.2 shows the results of running the specification in equation 1.7, using an indicator for a certain age or education group as dependent variable. Panels A and B of the table show the results of this exercise for women and men, respectively. The general conclusion we can extract from this table is that there was an increase in the share of relatively older and less educated individuals in the pool of divorcees after the reform in divorce law. For both women and men, the group of those aged between 50 and 60 years gained share at expenses of those between 30 and 50 years.[52] A similar pattern can be found between less and more educated women and men. While the share of divorced individuals with primary education or less increased, the share of those with secondary education or more fell.

To sum up, the analysis of the implications of divorce law liberalization reveals a statistically significant increase in the size of divorced population, though relatively low in magnitude, and a change towards an older and less educated composition of that population, a few years after the reform.

Marriage

The empirical literature on the effect of unilateral divorce in the U.S. has found a negative association between liberalization of divorce law and marriage rates (Wolfers, 2006; Rasul, 2006). Rasul (2006) develops a theoretical model of search and learning in the marriage market that shows how a less effective marriage contract would lead to a lower marriage rate in equilibrium and to better quality matches. In what follows I show that there is no evidence of a change in the propensity to marry that can be attributed to the law change, and that although there is little evidence of the law affecting the selection into marriage (i.e. changes in the composition of new couples), the magnitude of these effects is small enough to guarantee almost no impact on the composition of the stock of married couples after the reform.

[52]The total effect is shown in the bottom part of each panel of the table.

The evolution of the annual number of marriages for the period 1976-2009 is shown in Figure 1.10. To investigate whether there is a causal impact of the reform on the marriage rate, I estimate equation 1.8 using monthly data on marriages, and report the results in Table 1.14. Different columns correspond to different specifications for the trend -i.e. linear, quadratic, or cubic. The main conclusion is that there is no significant change in the propensity to marry that can be attributed to the legal change. The coefficient of the dummy for the post-reform period is always statistically indistinguishable from zero, and there is no significant change in the trend after the reform. The results of a Chow test, displayed at the bottom of the table, confirm that we can not reject the hypothesis of no structural break in the series of marriages.

To test whether there was an effect on the composition of those couples formed after the legal change, I use data form the Census of Marriages. This data contain information on main occupation, age at marriage, and previous legal civil status of spouses, for all marriages that take place in Spain. Figures A1.3 to A1.6 show the evolution of marriages according to these observable characteristics. Visual inspection of these figures points to little or no evidence of a sudden change in the observed characteristics of spouses after the reform in divorce law.

Tables A1.3 and A1.4 show the results of estimating equation 1.9 for the five main occupations of husbands and wives, respectively. The interpretation of the coefficients is the same as before. Since the total effect of the reform depends not only on the level change immediately after the reform but also on the change in the trend, its size depends on the period of time after the reform used in the calculation. The bottom part of the table reports the total effect on the outcome of interest measured 3 years after the introduction of the reform, with its corresponding standard error. The results show statistically significant impacts for 4 occupational categories of husbands and for 2 of wives. To analyze the magnitude of the effect, let us consider husbands who work in manufacturing, which represents around 40 percent of total employment of husbands in the pre-reform period. The share of husbands working in manufacturing fell by 2.5 percentage points after the legal change, which in terms of the counterfactual share (i.e. what manufacturing would have represented if the reform had not been introduced) is equivalent to a reduction of 6.30 percent. Is this effect relevant in terms of the stock of the population married, to an extent that would rise concern on the validity of the difference-in-differences estimates? The following exercise helps to illustrate this point. On average, between 2005 and 2010 the stock of married couples in Spain is 10.75 million, while the annual number of marriages (flow) is about 192 thousand. This means that each year approximately 1.8 percent (i.e. 192/10750=0.0178) of the stock of married couples is renewed.[53] Then, with a stock changing at a pace of 1.8 percent per year, the estimated effect (for the first 3 years) of the reform implies a change in the share of manufacturing workers within the stock of married couples of $1.8\% * 3years * (-6.30\%) = 0.34\%$. The same exercise can be performed for the rest of the major occupational groups and in no case is the final impact on the stock of married population above 1 percent. Then, we can conclude that there are no reasons to be concerned about the endogeneity of the treatment groups after the reform.

Table A1.5 reports the results for age at marriage (columns 1 to 3) and for the share of already divorced spouses. These results confirm the previous conclusion. There are statistically significant changes in age at marriage and in remarriage rates, but the magnitude of these changes in the flow of new couples has little relevance in terms of the composition of married population.

Summarizing, the analysis of the impact of the law change on the marriage market shows no evidence of a change in the propensity to marry and little evidence of an effect on the composition of new couples in terms of observable characteristics of spouses. Nevertheless, both the size of the impacts and the low ratio of new marriages to already married couples make

[53] To simplify I assume that the number of married couples dying or divorcing each year is the same as the number of new marriages in that year, so the size of the stock remains constant.

it safe to use a difference-in-differences approach to estimate the impact of easier divorce on domestic violence, using married and unmarried couples as treatment and comparison groups, respectively.

1.6 Conclusion

This study investigated whether easier access to divorce can reduce the incidence of spousal violence. To identify the causal effects, the study exploited exogenous variation in the cost of marital dissolution stemming from an unexpected reform of the divorce regime in Spain in 2005. This reform allowed for unilateral and no-fault divorce, and eliminated the 1-year mandatory separation period, reducing both economic and emotional costs of marital breakup. Furthermore, the study also took advantage of the fact that this change reduced the cost of terminating a relationship for couples who were legally married when the law became effective, but not for unmarried ones, and therefore, the empirical work follows a difference-in-differences methodology. The main findings point to a sizable decline in both non-extreme and extreme domestic abuse after the enactment of the new law.

The empirical analysis has revealed a decline in less extreme spousal conflict among married couples with respect to unmarried ones between 27 and 36 percent. Both self-reported spousal abuse and technical definitions of abuse based on recorded behavior confirm that the introduction of easier access to divorce has led to a decline in spousal conflict. These results are robust to the use of alternative definitions of domestic violence and are not driven by changes in the composition of the groups. Moreover, these findings are reinforced by the analysis of the heterogeneous responses to the legal change on the base of differences in the intensity of the treatment. Married women with young children are less affected by the reduction in the cost of divorce than childless women, because the presence of young children limits the reduction in the length of the divorce process. The results show a larger reduction in domestic violence against women without young children.

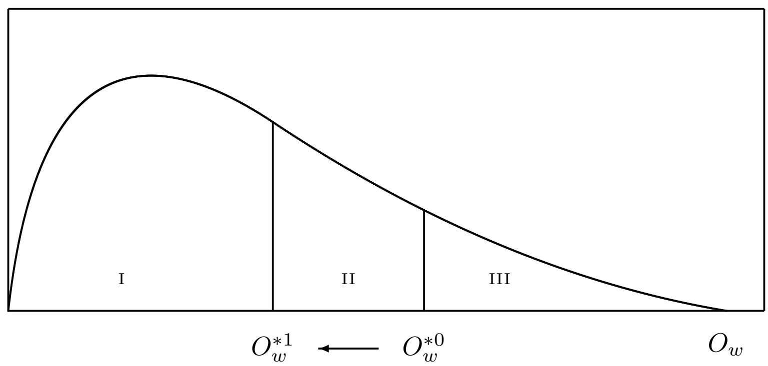
Easier divorce can reduce domestic violence either by increasing the propensity towards dissolution of abusive relationships, or by decreasing the propensity towards abuse in intact marriages. To disentangle these two channels, this study focused on the effects on couples who were married when the law was enacted and continue married 1 year later, when the data were observed. The results suggest that the bulk of the reduction in violence can be explained by the improvement in the bargaining position of wives in marriages that remain intact. The availability of easier access to divorce thus seems to make the threat of leaving the marriage more credible, which is shown to be a strong deterrent of spousal violence.

This study also measured the effect of the reduction in divorce costs on the most extreme measure of spousal conflict, partner homicide, and observed a decline of around 30 percent among married women, which can be attributed to the legal change. An important part of this decline is explained by a reduction in lethal violence against spouses who were amid the process of marital dissolution. These results, taken together with an increase in the share of conflicting dissolutions after the reform, suggest an important role for the duration of the divorce process as a key factor behind the reduction in lethal violence.

This study also examined whether the legal change affected marital formation and dissolution patterns. Investigation of the potential effect on the marriage market is crucial to assess the extent to which the homicide results may be influenced by selection into marriage. The study tested whether the reform had an impact on either the propensity to marry or the composition characteristics of new marriages, and found no significant evidence of either of these effects, suggesting that selection into marriage effects is not an important concern. Examining the potential effect on marital dissolution is important to assess the channels through which easier divorce affects domestic violence. This study found that the stock of separated and divorced individuals in the post-reform period is around 2 percent higher because of the reform, suggesting that the dissolution channel may explain only a small part of the reduction in violence.

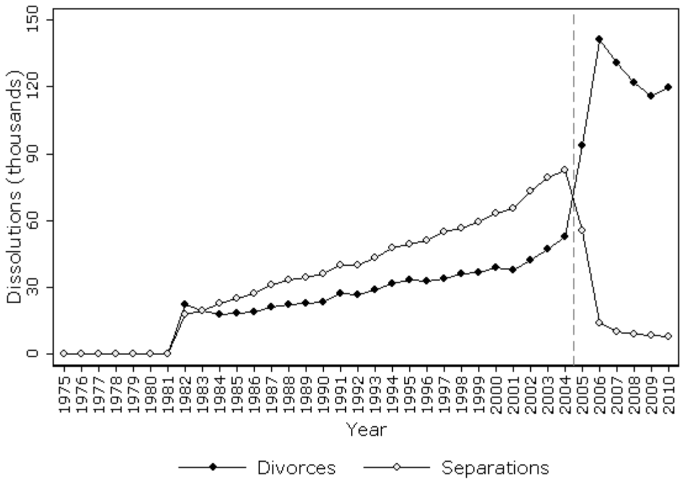
Figures

Figure 1.1: Distribution of wive's outside opportunities and reduction in the cost of divorce.



Notes: Before the reform, the marginal woman (i.e. the one that is indifferent between divorce and an abusive marriage) had an outside option given by O_w^{*0}. When the cost of divorce falls to $C_w^1 < C_w^0$, the new marginal woman places to the left, say at O_w^{*1}. For those husbands whose wife's outside option lies in between these two values, it was optimal to be violent under the old regime but it is not after the reduction in the cost of divorce. Women located in sector II, therefore, will benefit from the reduction in divorce costs. For women in sector I the reduction in the cost of divorce is not enough for them to credibly threaten with divorce, while for women in sector III the reform has no relevant effects since they were not affected by domestic violence.

Figure 1.2: Marital dissolution in Spain, 1975-2010

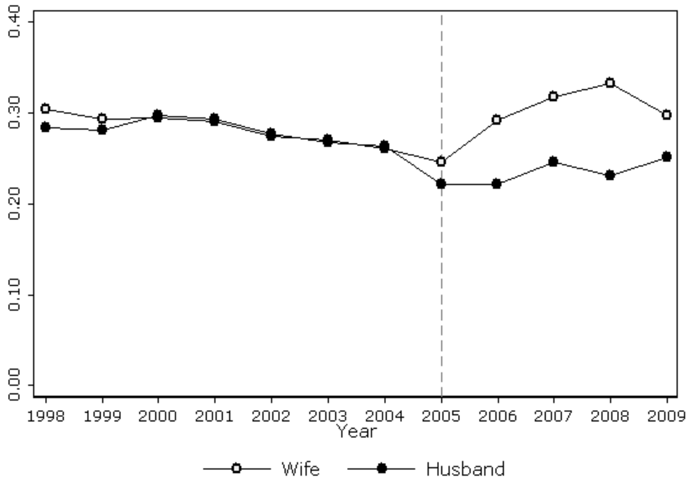


Source: Judiciary statistics provided by the General Council of the Judiciary.

Figure 1.3: Distribution of legal separations according to who is the petitioner

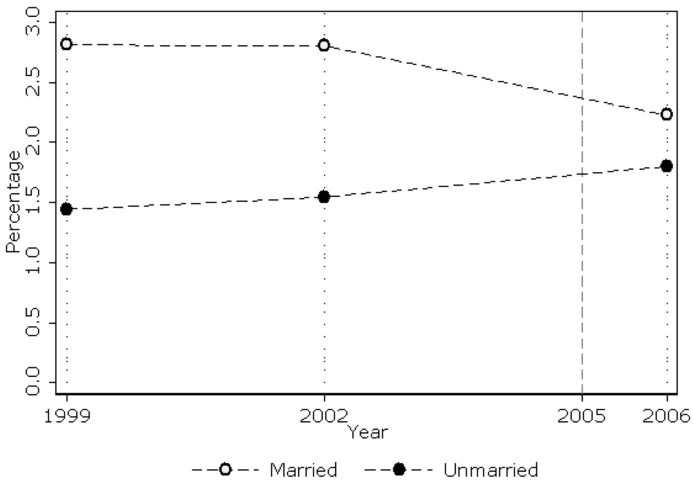
Source: National Institute of Statistics of Spain.

Figure 1.4: Distribution of divorces according to who is the petitioner



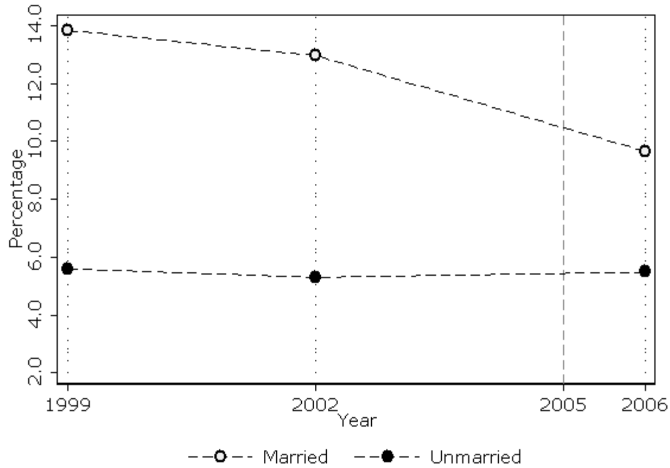
Source: National Institute of Statistics of Spain.

Figure 1.5: Self-Reported Abuse during previous year



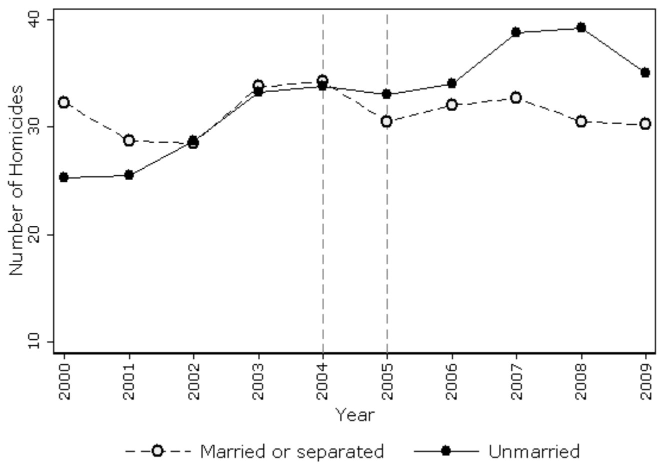
Source: Survey on Domestic Violence Against Women. Spanish Women's Institute

Figure 1.6: Technical Measure of Abuse.



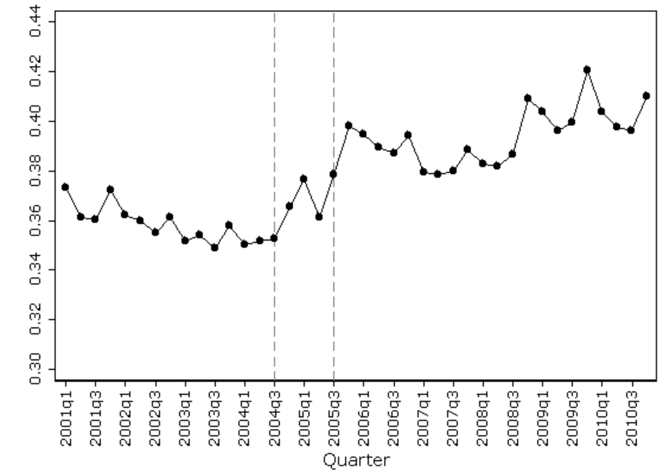
Source: Survey on Domestic Violence Against Women. Spanish Women's Institute

Figure 1.7: Evolution of the annual number of intimate partner female homicides, 2000-2010



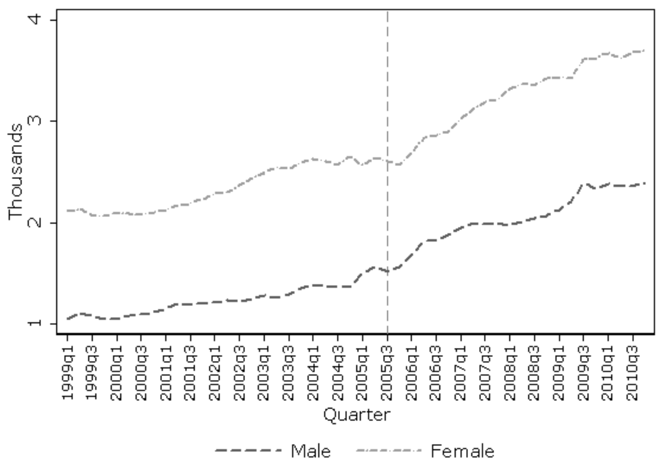
Source: Spanish Women's Institute. Moving average (centered - 2 years)

Figure 1.8: Share of Adversarial Dissolutions by Quarter.



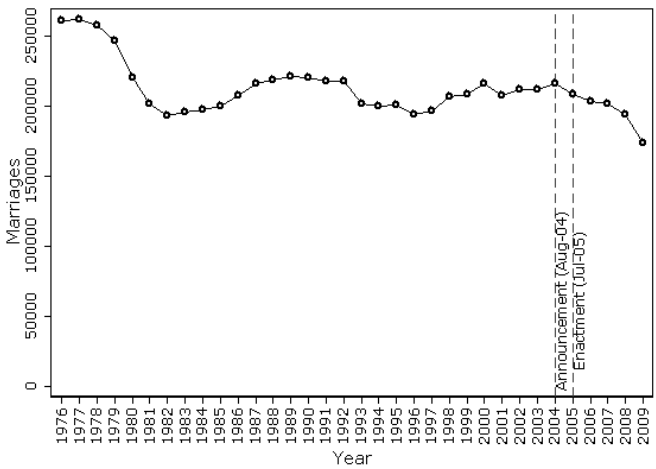
Note: Share of adversarial dissolutions (separation and divorces) over total dissolutions. Source: Judiciary statistics provided by the General Council of the Judiciary.

Figure 1.9: Evolution of the Stock of legally separated and divorced people.



Source: Spanish Labor Force Survey. National Institute of Statistics of Spain

Figure 1.10: Evolution of total annual number of marriages



Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain

Tables

Table 1.1: Descriptive Statistics. Survey on Violence Against Women.

	Total sample		1999		2002		2006	
	Mean	St dev	Mean	St dev	Mean	St dev	Mean	St dev
Woman's age								
18-29	0.212	0.409	0.238	0.426	0.209	0.407	0.197	0.398
30-39	0.193	0.394	0.179	0.384	0.184	0.388	0.206	0.405
40-49	0.169	0.375	0.153	0.360	0.163	0.370	0.183	0.387
50-59	0.148	0.356	0.143	0.350	0.152	0.359	0.149	0.356
60 or older	0.278	0.448	0.286	0.452	0.291	0.454	0.264	0.441
Woman's education								
Primary or less	0.384	0.486	0.442	0.497	0.413	0.492	0.330	0.470
Lower High School	0.238	0.426	0.212	0.409	0.240	0.427	0.252	0.434
Upper High School	0.198	0.399	0.192	0.394	0.193	0.395	0.205	0.404
University	0.180	0.384	0.154	0.361	0.154	0.361	0.213	0.409
Woman's marital status								
Single	0.112	0.316	0.122	0.328	0.109	0.312	0.108	0.311
Dating	0.100	0.299	0.108	0.310	0.100	0.300	0.094	0.292
Cohabiting	0.029	0.168	0.018	0.133	0.024	0.154	0.039	0.194
Married	0.617	0.486	0.604	0.489	0.625	0.484	0.619	0.486
Separated	0.021	0.142	0.020	0.139	0.018	0.133	0.023	0.149
Divorced	0.014	0.118	0.011	0.104	0.014	0.116	0.017	0.128
Woman's labor market status								
Employed	0.347	0.476	0.301	0.459	0.308	0.461	0.402	0.490
Unemployed	0.078	0.268	0.079	0.270	0.079	0.270	0.076	0.265
Out of labor force	0.573	0.495	0.619	0.486	0.612	0.487	0.520	0.500
Woman's partnerships								
In a relationship	0.756	0.429	0.739	0.439	0.759	0.428	0.765	0.424
Duration relationship	22.234	15.118	21.791	15.109	22.824	15.157	22.131	15.088
Children								
Children	0.709	0.454	0.694	0.461	0.714	0.452	0.714	0.452
N^{O} of children	1.659	1.442	1.685	1.512	1.689	1.457	1.623	1.386
Partner's age								
Age	51.526	14.635	51.477	14.840	51.958	14.583	51.282	14.539
Partner's education								
Primary or less	0.336	0.472	0.402	0.490	0.363	0.481	0.279	0.449
Lower High School	0.262	0.440	0.236	0.425	0.272	0.445	0.273	0.445
Upper High School	0.209	0.407	0.196	0.397	0.202	0.401	0.223	0.416
University	0.192	0.394	0.166	0.372	0.163	0.370	0.226	0.418
Partner's labor market status (during last year)								
Not employed	0.328	0.470	0.354	0.478	0.349	0.477	0.299	0.458
Part-time employment	0.028	0.165	0.030	0.171	0.034	0.181	0.023	0.151
Full-time employment	0.635	0.481	0.604	0.489	0.608	0.488	0.670	0.470
Religion								
Practicing Catholic	0.382	0.486	0.456	0.498	0.391	0.488	0.329	0.470
Not Practicing Catholic	0.481	0.500	0.443	0.497	0.487	0.500	0.501	0.500
Agnostic/Atheist	0.077	0.267	0.056	0.230	0.062	0.240	0.100	0.301
Other religion	0.024	0.153	0.019	0.135	0.022	0.146	0.029	0.167
None religion	0.036	0.187	0.027	0.163	0.039	0.194	0.040	0.197
Urban/Rural								
Urban (more than 10k pop)	0.772	0.420	0.755	0.430	0.749	0.433	0.797	0.402
Sample size	73630		20552		20652		32426	

Table 1.2: Descriptive Statistics. Female Homicides by Intimate Partner.

	Homicides per quarter				Homicides per quarter and marital status		
Year	mean	st dev	min	max	Unmarried	Married	Separated or divorced
2000	12.5	1.3	11	14	4.3	7.0	1.3
2001	11.5	2.4	10	15	4.8	4.0	2.8
2002	13.0	2.4	10	15	6.8	4.8	1.5
2003	17.5	2.4	14	19	8.8	6.0	2.8
2004	17.3	3.6	12	20	6.8	8.0	2.5
2005	15.8	1.0	15	17	9.5	5.3	1.0
2006	17.5	4.8	13	23	8.0	7.0	2.5
2007	18.0	2.9	14	21	9.0	6.3	2.8
2008	18.8	4.8	12	23	10.5	6.0	2.3
2009	15.3	3.3	12	19	8.0	5.3	2.0
2010	18.8	4.1	14	24	9.3	7.3	2.3
Total	16.0	3.8	10	24	7.8	6.1	2.1

Note: Marital status defined in terms of victim-offender relationship. Cases in which the victim was not legally married to the aggressor at the moment of the homicide nor before are classified as Unmarried. Married victims are those who were legally married to the aggressor. Separated victims include those victims who were previously married to the aggressor or who were still legally married but in process of separation or divorce. Source: Queen Sofia Center.

Table 1.3: Impact on Non-Extreme Violence: Self-Reported Abuse

	Dependent variable: Self-Reported Abuse (dummy)					
	Abuse during previous year					Abuse before previous year (placebo)
	(1)	(2)	(3)	(4)	(5)	(6)
$Married * Post$	-0.746***	-0.651***	-0.647***	-0.599**	-0.686***	0.078
	(0.212)	(0.212)	(0.212)	(0.295)	(0.205)	(0.211)
$Married$	1.720***	3.457***	3.450***	0.726	1.633***	-0.819
	(0.140)	(0.763)	(0.764)	(0.755)	(0.587)	(0.595)
Individual controls	No	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Region controls	No	No	Yes	Yes	Yes	Yes
Partner controls	No	No	No	Yes	No	No
Adj. R^2	0.002	0.024	0.025	0.008	0.007	0.030
N	69895	69838	69838	54757	67895	69838
RMSE	14.682	14.515	14.512	14.473	13.507	12.910
Mean (depvar)	2.209	2.208	2.208	2.159	1.872	1.748

Notes: The sample includes adult females in 1999, 2002, and 2006, who had a partner during the year before the interview. The dependent variable is a binary indicator for self-reported abuse during the previous year (columns 1-5) or any time in life before the previous year (column 6). The treatment group includes women who were married at the moment of the reform in divorce legislation, independently of their current marital status, with the exception of column 5, which restricts the treatment group to women who were married when the reform was passed and continue married when the survey was conducted. The control group includes women with partner during the previous year but who are not legally married. Individual control variables include age group dummies, education dummies, labor market status, dummies for legal civil status, a dummy for the presence of children, the number of children, immigration status, and dummies for religion beliefs. Region controls include region fixed effects and a dummy for urban residence. Partner controls include dummies for education and labor market status of the partner. Since partner variables refer to the current partner, including these controls (column 4) implies restricting the sample to women with partner at the moment of the interview. All regressions include year dummies. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.4: Impact on Non-Extreme Violence: Technical Measures of Abuse

	Dependent variable: Measures of technical abuse (dummies)						
	Technical abuse			Physical abuse	Sexual abuse	Psych. abuse (control)	Psych. abuse (emo- tional)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$Married * Post$	-4.206***	-3.278***	-3.258***	-0.592**	-1.472***	-0.487	-2.230***
	(0.543)	(0.541)	(0.546)	(0.233)	(0.311)	(0.306)	(0.444)
$Married$	8.126***	0.655	0.709	0.577	1.241*	-1.193	0.154
	(0.367)	(1.415)	(1.421)	(0.360)	(0.723)	(0.989)	(1.159)
Individual controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Partner controls	No	No	Yes	Yes	Yes	Yes	Yes
Adj. R^2	0.009	0.020	0.022	0.009	0.012	0.006	0.017
N	55535	55495	54757	54757	54757	54757	54757
RMSE	30.662	30.472	30.449	13.342	18.410	16.107	25.845
Mean (depvar)	10.613	10.601	10.605	1.830	3.556	2.681	7.334

Notes: The sample includes adult females in 1999, 2002, and 2006 who have a partner at the moment of the interview. The dependent variable are binary variables for technical abuse (columns 1-3), physical abuse (column 4), sexual abuse (column 5), psychological abuse in the form of control (column 6), and psychological abuse in the form of emotional mistreatment (column 7). The treatment group includes currently married women who were married at the moment of the reform in divorce legislation. The control group includes women with partner but who are not legally married. Individual control variables include age group dummies, education dummies, labor market status, dummies for legal civil status, a dummy for the presence of children, the number of children, dummies for religion beliefs. Region controls include region fixed effects and a dummy for urban residence. Partner controls include dummies for education and labor market status of the partner. All regressions include year dummies. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.5: Impact on Non-Extreme Violence: Alternative Definitions of Technical of Abuse

	Dependent variable: Measures of technical abuse						
	N of Indicators of abuse		Dummies for at least n indicators of abuse (n=2...6)				
	OLS	Poisson	+2 indicators	+3 indicators	+4 indicators	+5 indicators	+6 indicators
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$Married * Post$	-0.074***	-0.323***	-1.342***	-0.771***	-0.711***	-0.438***	-0.386***
	(0.015)	(0.119)	(0.341)	(0.260)	(0.209)	(0.165)	(0.132)
$Married$	0.010	0.068	0.419	0.020	-0.183	0.118	0.084
	(0.036)	(0.266)	(0.832)	(0.628)	(0.538)	(0.331)	(0.323)
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Partner controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj. R^2	0.018		0.015	0.010	0.008	0.006	0.005
N	54757	54757	54757	54757	54757	54757	54757
RMSE	0.882		20.273	15.638	12.517	10.214	8.486
Mean (depvar)	0.221	0.221	4.361	2.535	1.605	1.061	0.729

Notes: The sample includes adult females in 1999, 2002, and 2006 who have a partner at the moment of the interview. In columns 1-2, the dependent variable is a continuous variable for the number of indicators of abuse present for each individual. In columns 3-7, the dependent variable is a dummy that takes the value 1 if at least n indicators of abuse are present (for n=2...6). The treatment group includes currently married women who were married at the moment of the reform in divorce legislation. The control group includes women with partner but who are not legally married. Individual control variables include age group dummies, education dummies, labor market status, dummies for legal civil status, a dummy for the presence of children, the number of children, dummies for religion beliefs. Region controls include region fixed effects and a dummy for urban residence. Partner controls include dummies for education and labor market status of the partner. All regressions include year dummies. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.6: Heterogeneous impact by presence of young children using unmarried women as control group.

	Dependent variable	
	Self-Reported Abuse	Technical Abuse
	(1)	(2)
Panel A: Women with young children		
Married * Post	1.475	0.103
	(1.221)	(2.215)
Married	6.598**	-3.154*
	(3.203)	(1.755)
Post	-1.913	-2.688
	(1.215)	(2.211)
Panel B: Women without young children		
Married * Post	-0.854***	-4.123***
	(0.246)	(0.617)
Married	3.012***	2.959***
	(0.727)	(0.690)
Post	0.182	-0.173
	(0.166)	(0.525)

Notes: The sample is split between mothers of children under 18 years of age and women without young children, independently of whether they are mothers or not. Each sub-sample includes adult females in 1999, 2002, and 2006. Dependent variables are dummy variables for different measures of abuse. Self-reported abuse refers to the last 12 months, while all technical measures of abuse refers to current situation. The treatment group includes women who were married at the moment of the reform in divorce legislation, independently of their current marital status. The control group includes women with partner but who are not legally married. The control variables included in the regressions are: age group dummies, education dummies, age and education of the husband, number of children, region fixed effects and year fixed effects. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.7: Heterogeneous impact by presence of young children (Only married women).

	Dependent variable	
	Self-Reported Abuse	Technical Abuse
	(1)	(2)
Without young children * Post	-0.621*	-1.298*
	(0.355)	(0.749)
Without young children	0.639*	-0.267
	(0.331)	(0.688)
Post	-0.495*	-2.706***
	(0.257)	(0.553)
Individual controls	Yes	Yes
Year dummies	Yes	Yes
Region controls	Yes	Yes
Partner controls	Yes	Yes
Adj. R^2	0.004	0.016
N	29812	29812

Notes: The sample includes married women between 30 and 60 years of age. The treatment group includes mothers of young children (under 18 years of age), while women either without children or with children older than 18 years of age are left in the control group. Dependent variables are dummy variables for different measures of abuse. Self-reported abuse refers to the last 12 months, while technical abuse refers to current situation. The control variables included in the regressions are: age group dummies; education dummies; age, education, and labor market status of the husband; number of children; region fixed effects; and year fixed effects. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.8: Heterogeneous impacts by education level using unmarried women as a control group.

	Dependent variable	
	Self-Reported Abuse	Technical Abuse
	(1)	(2)
Panel A: Women with low education level		
Married * Post	-0.753**	-8.842***
	(0.316)	(3.050)
Married	2.837***	3.662*
	(1.027)	(1.976)
Post	0.351	3.794
	(0.248)	(3.027)
Panel B: Women with intermediate education level		
Married * Post	-0.836**	-2.742***
	(0.347)	(0.748)
Married	3.826***	1.157
	(1.418)	(0.802)
Post	-0.108	-0.376
	(0.284)	(0.665)
Panel C: Women with high education level		
Married * Post	-0.324	-1.350
	(0.495)	(0.956)
Married	4.013**	1.164
	(1.628)	(0.987)
Post	-0.252	-0.966
	(0.351)	(0.825)

Notes: The sample is split by education level of women. Low education includes women with primary school or less, intermediate education accounts for women with high school, while high education accounts for women with a university degree. Each sub-sample includes adult females in 1999, 2002, and 2006. Dependent variables are dummy variables for different measures of abuse. Self-reported abuse refers to the last 12 months, while all technical measures of abuse refers to current situation. The treatment group includes women who were married at the moment of the reform in divorce legislation, independently of their current marital status. The control variables included in the regressions are: age group dummies, age and education of the husband, presence of young children at home, number of children, a dummy for urban-rural residence, region fixed effects, year fixed effects, immigration status, and religion beliefs. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.9: Heterogeneous impacts by education level using unmarried women as a control group. Full sample.

	Dependent variable	
	Self-Reported Abuse	Technical Abuse
	(1)	(2)
Married * Post	-0.484*	-4.381***
	(0.269)	(0.678)
Married * Post * Intermediate Education	-0.332	1.641**
	(0.294)	(0.650)
Married * Post * High Education	-0.210	1.993***
	(0.368)	(0.757)
Married	3.445***	0.414
	(0.768)	(1.435)
Post	0.008	-0.429
	(0.169)	(0.500)
Intermediate Education	0.122	-0.459
	(0.193)	(0.579)
High Education	-0.168	-1.476**
	(0.231)	(0.660)
Adj. R^2	0.024	0.022
N	69886	54779

Notes: The sample includes all adult women in 1999, 2002, and 2006. Low education (omitted category) includes women with primary school or less, intermediate education accounts for women with high school, while high education accounts for women with a university degree. Dependent variables are dummy variables for different measures of abuse. Self-reported abuse refers to the last 12 months, while all technical measures of abuse refers to current situation. The treatment group includes women who were married at the moment of the reform in divorce legislation, independently of their current marital status. The control group includes women with partner but who are not legally married. The control variables included in the regressions are: age group dummies, age and education of the husband, presence of young children at home, number of children, a dummy for urban-rural residence, region fixed effects, year fixed effects, immigration status, and religion beliefs. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.10: Heterogeneous impacts by education level including only married women in the sample.

	All married women		Married women aged 30-50	
	Dependent variable		Dependent variable	
	Self-Reported Abuse (1)	Technical Abuse (2)	Self-Reported Abuse (3)	Technical Abuse (4)
Intermediate Education * Post	-0.450	1.134	-1.088**	-0.592
	(0.329)	(0.711)	(0.541)	(1.209)
High Education * Post	-0.205	2.296***	-0.621	0.833
	(0.433)	(0.878)	(0.619)	(1.331)
Intermediate Education	0.738**	-0.379	0.953**	0.479
	(0.310)	(0.646)	(0.417)	(0.925)
High Education	0.786*	-2.034**	0.936*	-1.282
	(0.430)	(0.832)	(0.532)	(1.098)
Post	-0.469*	-4.525***	0.015	-2.890**
	(0.264)	(0.597)	(0.495)	(1.132)
Adj. R^2	0.004	0.015	0.004	0.014
N	40535	40535	20741	20741

Notes: The sample includes all married women in columns 1-2, and middle aged women (30-50 years) in columns 3-4. Low education (omitted category) includes women with primary school or less, intermediate education accounts for women with high school, while high education accounts for women with a university degree. Dependent variables are dummy variables for different measures of abuse. Self-reported abuse refers to the last 12 months, while technical abuse refers to current situation. The control variables included in the regressions are: age group dummies; education dummies; age, education, and labor market status of the husband; number of children; region fixed effects; and year fixed effects. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Survey of Violence Against Women 1999, 2002, and 2006.

Table 1.11: Effect of the Divorce Law Reform on the Female Homicide by Intimate Partner.

	Poisson		Ordinary Least Squares			
	Homicide counts		Homicide rate		Homicide rate in logs	
	(1)	(2)	(3)	(4)	(5)	(6)
Post	0.356*	0.272	0.568***	0.456**	0.421**	0.335*
	(0.187)	(0.190)	(0.193)	(0.219)	(0.187)	(0.196)
Married	-0.813***	0.471	-0.758***	0.950	-0.775***	0.530
	(0.091)	(1.272)	(0.119)	(1.809)	(0.111)	(1.609)
Post*Married	-0.326***	-0.164	-0.526***	-0.303	-0.376***	-0.205
	(0.114)	(0.201)	(0.155)	(0.270)	(0.137)	(0.238)
Quarter dummies	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	$\chi^2 = 21.99$	$\chi^2 = 26.74$	$F = 1.09$	$F = 1.24$	$F = 1.80$	$F = 2.12$
Group linear trend	No	$\chi^2 = 1.03$	No	$F = 0.90$	No	$F = 0.68$
Adj. R^2			0.719	0.718	0.741	0.740
Goodness-of-fit chi2	46.259					
Prob > chi2(71)	0.99					
Mean dependent variable	7.977	7.977	1.140	1.140	-0.046	-0.046

Notes: The sample includes the number of adult female homicides by quarter, 2000-2010. The dependent variable is constructed by aggregating the number of homicides per group and quarter, and is defined as a count variable (columns 1-2), as a rate in terms of the size of corresponding group population (columns 3-4), and as the logarithm of the rate (columns 5-6). Columns 2, 4, and 6 are similar to columns 1, 3, and 5, respectively, except for the inclusion of group-specific linear trends. The treatment group includes homicides of women who were either married to or separated at the moment of the homicide. All other victim-perpetrator relationships (cohabiting couples, romantic partners) are included in the control group. All regressions include 88 observations (11 years x 4 quarters x 2 groups). Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Mortality Statistics collected by Queen Sofia Center.

Table 1.12: Effect of the Divorce Law Reform on the Female Homicide by Intimate Partner: Married versus Separated Women.

	Poisson	Ordinary Least Squares	
	Homicide counts (1)	Homicide rate (2)	Homicide rate in logs (3)
Post	0.402*	2.064**	0.673**
	(0.235)	(0.969)	(0.280)
Married	-0.556***	-0.391*	-0.579***
	(0.123)	(0.201)	(0.150)
Separated	1.509***	3.249***	1.417***
	(0.166)	(0.600)	(0.170)
Post*Married	-0.315**	-0.354	-0.328*
	(0.148)	(0.261)	(0.179)
Post*Separated	-0.657***	-1.437**	-0.643***
	(0.211)	(0.723)	(0.212)
Quarter dummies	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes
Adj. R^2		0.494	0.723

Notes: The sample includes the number of adult female homicides by quarter, 2000-2010. The dependent variable is constructed by aggregating the number of homicides per group and quarter, and is defined as a count variable (column 1), as a rate in terms of the size of corresponding group population (column 2), and as the logarithm of the rate (column 3). There are two treatment groups, depending on the legal status of the victim at the moment of the homicide: (i) Women who were legally married to the perpetrator, and (ii) Women who were already separated or in the process of separation from the perpetrator. All other victim-perpetrator relationships (cohabiting couples, romantic partners) are included in the control group. All regressions include 132 observations (11 years x 4 quarters x 3 groups). Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Mortality Statistics collected by Queen Sofia Center.

Table 1.13: Impact on Marital Dissolution.

	Dependent variable: 1 if divorced		
	Total (1)	Women (2)	Men (3)
time	0.518***	0.623***	0.407***
	(0.027)	(0.043)	(0.032)
$post^{2005}$	3.252***	1.659**	4.866***
	(0.447)	(0.701)	(0.550)
timepost	0.056	0.091	0.020
	(0.042)	(0.066)	(0.052)
Divorcees (per 1000)	42.398	53.947	30.542
Effect	3.926***	2.746***	5.105***
	(0.615)	(0.975)	(0.741)
Change (%)	2.02	1.04	3.38
Adj. R^2	0.018	0.017	0.012
N	3404397	1724559	1679838

Notes: The sample includes individuals between 20 and 60 years of age in all quarters between 2001 and 2009. The dependent variable is a dummy variable set equal to 1 if the person declare to be separated or divorced at the moment of the interview. The control variables are dummies for age groups and education levels, plus quarter fixed effects to control for seasonality. When both men and women are included in the sample (column 1), a dummy for sex is also included. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics, Spain.

Table 1.14: Structural break tests for time series of marriages.

	Dependent variable: number of marriages per month		
	Linear trend (1)	Quadratic trend (2)	Cubic trend (3)
$Time$	-6.67***	-35.11***	-79.49***
	(1.79)	(7.52)	(22.00)
$Timesq$		0.07***	0.34***
		(0.02)	(0.13)
$Timecu$			-0.00**
			(0.00)
$Post^{2005}$	259.42	-1966.27	-1423.39
	(908.72)	(1427.96)	(2057.07)
$Timepost$	-17.14	45.90	135.51
	(27.39)	(111.69)	(294.66)
$Timepostsq$		-1.70	-3.95
		(1.96)	(12.06)
$Timepostcu$			0.03
			(0.14)
$Marriages(L)$	0.18***	0.13***	0.12**
	(0.05)	(0.05)	(0.05)
$GDPgrowth(L12)$	347.71***	332.48***	314.67***
	(104.07)	(105.12)	(105.92)
Constant	6051.71***	8823.96***	10965.51***
	(968.23)	(1183.33)	(1546.12)
Chow test of structural break			
Ho:	b[$Timepost$]=0	b[$Timepost$]=0 b[$Timepostsq$]=0	b[$Timepost$]=0 b[$Timepostsq$]=0 b[$Timepostcu$]=0
F test	0.392	1.811	0.335
p-value	0.532	0.165	0.715
Adj. R^2	0.831	0.837	0.839
N	395	395	395
Durbin-Watson	1.968	1.961	1.955

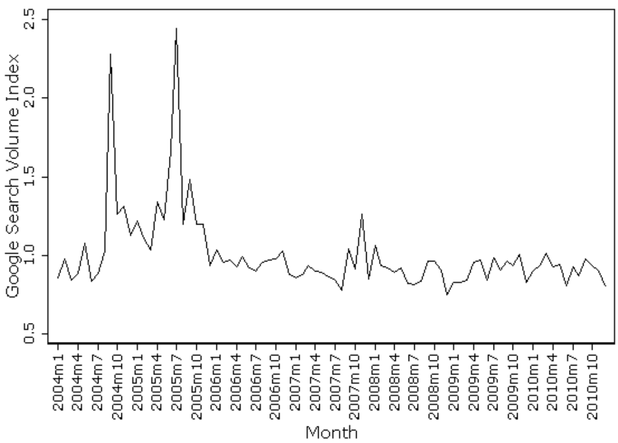
Notes: The sample includes all marriages occurred between 1976 and 2009 on a monthly basis. $Post^{2005}$ is a dummy variable set equal to 1 since July 2005. All regressions have month fixed effects. Robust standard errors are reported in parentheses. *, **, and*** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Appendix A1. Additional Figures and Tables

Figure A1.1: Divorces according to who is the petitioner

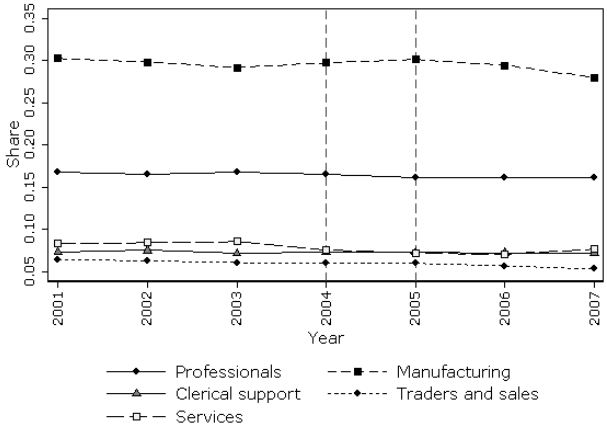
Source: National Institute of Statistics, Spain.

Figure A1.2: Google Search Volume Index for the query "divorcio"



Source: Google Trends.

Figure A1.3: Evolution of marriages by husband's occupation

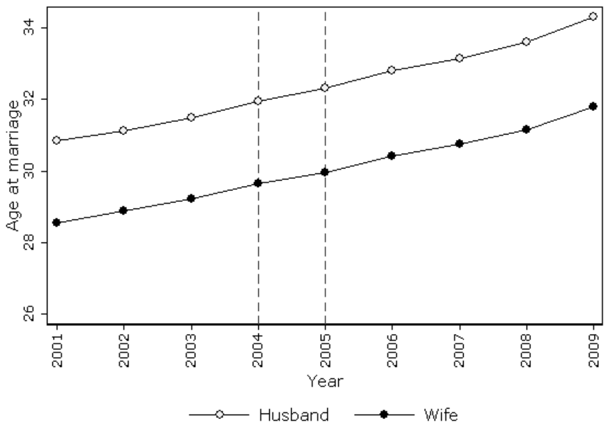


Notes: These five occupations of husbands account for almost 90 percent of all husbands. A change in the coding of occupation after 2007 makes it not possible to continue the series for a longer period at the same level of disaggregation. Vertical lines in 2004 and 2005 indicate the years of announcement and enactment of the legal change, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Figure A1.4: Evolution of marriages by wife's occupation

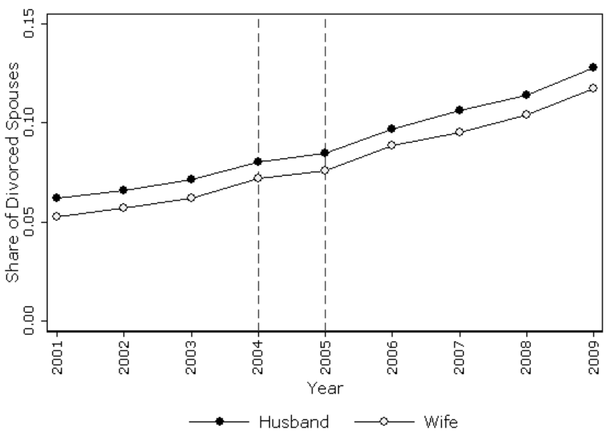
Notes: These five occupations of wives account for 87 percent of all wives. A change in the coding of occupation after 2007 makes it not possible to continue the series for a longer period at the same level of disaggregation. Vertical lines in 2004 and 2005 indicate the years of announcement and enactment of the legal change, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Figure A1.5: Average age at marriage



Notes: Vertical lines in 2004 and 2005 indicate the years of announcement and enactment of the legal change, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Figure A1.6: Evolution of marriages by spouses' civil status: Divorced



Notes: Vertical lines in 2004 and 2005 indicate the years of announcement and enactment of the legal change, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Table A1.1: Measures of Technical Abuse: Definitions and Frequencies.

Definition of Technical Abuse			Classification according to Alberdi and Matas (2002)		
Indicator of abuse	mean	st. dev.	Type of abuse	mean	st. dev.
Insults or threatens you	0.0129	(0.1126)	Physical abuse	0.0190	(0.1367)
Makes you to become afraid	0.0096	(0.0974)			
Pushes or hits you when he becomes angry	0.0063	(0.0793)			
Insists on having sexual intercourse even if he knows you do not want to	0.0370	(0.1888)	Sexual abuse	0.0370	(0.1888)
Prevents you from visiting your family or relate to your friends, neighbors	0.0127	(0.1122)	Psychological abuse (control)	0.0282	(0.1654)
Takes the money you earn or does not give what you need	0.0037	(0.0606)			
Decides what yo can do or not do	0.0179	(0.1326)			
Does not care about your needs	0.0181	(0.1333)	Psychological abuse (emotional mistreatment)	0.0752	(0.2638)
Says where would you go without him	0.0142	(0.1185)			
Says that everything you do is always wrong, that you are clumsy	0.0176	(0.1313)			
Ridicules you or does not value your beliefs (religious, political, etc)	0.0136	(0.1157)			
Does not value the job or tasks you do	0.0445	(0.2063)			
Blames you in from of your children	0.0213	(0.1443)			
Technical abuse	0.1095	(0.3123)			

Notes: The measure of technical abuse is based on a series of 13 questions included in the survey as indicators of abuse according to the opinion of experts. This part of the questionnaire is answered only by women who declare to be in a relationship at the moment of the survey, independently of their legal civil status. For each indicator of abuse, there is information on the frequency of occurrence (i.e. frequently, sometimes, rarely, never) and on who is the offender. In this paper, I follow the same criterion the Spanish Women's Institute established when published the data, that is, to consider a situation of intimate partner abuse exists when there is a positive response to the correspondent question, the situation occurs "frequently" or "sometimes", and the offender is the intimate partner. I also follow Alberdi and Matas (2002) classification of these 13 indicators of abuse into 4 categories: physical, sexual, and two forms of psychological abuse.

Table A1.2: Impact on the Composition of the Stock of Divorcees

	Age Group				Education Level			
	16-29	30-39	40-49	50-60	Less than primary	Primary	High school	University
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Women								
time	-0.089***	-0.131***	0.136***	0.065	-0.160***	-0.457***	0.452***	0.165***
	(0.021)	(0.044)	(0.048)	(0.043)	(0.024)	(0.039)	(0.048)	(0.039)
$post^{2005}$	-0.212	2.482***	-1.520**	-1.682***	0.199	-0.705	0.017	0.489
	(0.263)	(0.593)	(0.658)	(0.596)	(0.294)	(0.504)	(0.665)	(0.554)
timepost	0.015	-0.247***	0.005	0.356***	0.127***	0.251***	-0.421***	0.043
	(0.026)	(0.057)	(0.063)	(0.058)	(0.030)	(0.050)	(0.064)	(0.053)
Share before	3.952	25.905	40.568	28.081	5.081	17.461	55.276	22.182
Effect	-0.029	-0.478	-1.466	2.586***	1.727***	2.308***	-5.034***	0.999
	(0.439)	(0.922)	(1.014)	(0.915)	(0.483)	(0.815)	(1.031)	(0.839)
Change (%)	-0.96	-1.93	-3.43	9.36	64.08	19.33	-8.09	4.32
Panel B: Men								
time	-0.074***	-0.018	0.267***	-0.166***	-0.241***	-0.386***	0.462***	0.165***
	(0.024)	(0.059)	(0.066)	(0.061)	(0.032)	(0.056)	(0.067)	(0.054)
$post^{2005}$	-0.214	-1.063	-1.740**	1.456*	1.326***	-1.593**	-0.723	0.990
	(0.282)	(0.768)	(0.886)	(0.825)	(0.380)	(0.699)	(0.895)	(0.749)
timepost	0.031	-0.219***	-0.245***	0.559***	0.162***	0.237***	-0.327***	-0.073
	(0.028)	(0.074)	(0.086)	(0.081)	(0.039)	(0.069)	(0.087)	(0.072)
Share before	2.477	23.143	40.621	32.016	4.916	18.594	54.140	22.350
Effect	0.154	-3.696***	-4.683***	8.163***	3.267***	1.256	-4.643***	0.120
	(0.490)	(1.237)	(1.399)	(1.299)	(0.636)	(1.162)	(1.421)	(1.162)
Change (%)	8.62	-14.61	-10.27	31.40	394.39	8.46	-7.65	0.51

Notes: The sample includes individuals between 20 and 60 years of age in all quarters between 2001 and 2009. In each column, the dependent variables is a dummy variable set equal to 1 if the person corresponds to that particular age or education group. All regressions include quarter fixed effects to control for seasonality. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics, Spain.

Table A1.3: Impact on the composition of new marriages according to main occupation of the husband.

	Five main occupations				
	Manufacturing	Professionals	Services	Clerical support	Sales and trading
	(1)	(2)	(3)	(4)	(5)
Time	0.024***	0.002	-0.022***	0.007***	-0.006***
	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)
$Post^{2005}$	1.203***	-0.141	-0.810***	-0.094	0.101
	(0.200)	(0.169)	(0.123)	(0.122)	(0.110)
Timepost	-0.103***	0.024***	0.070***	0.003	-0.023***
	(0.010)	(0.008)	(0.006)	(0.006)	(0.005)
Share before (%)[a]	40.124	22.301	10.620	9.868	8.071
Effect[b]	-2.493***	0.705***	1.716***	0.019	-0.717***
	(0.320)	(0.272)	(0.200)	(0.196)	(0.177)
Change (%)[c]	-6.30	3.41	13.55	0.24	-9.56
Adj. R^2	0.001	0.001	0.001	0.001	0.000
N	1075700	1075700	1075700	1075700	1075700

Notes: The sample includes all marriages occurred between 2001 and 2007. $Post^{2005}$ is a dummy variable set equal to 1 since July 2005. The dependent variable is a dummy for each of the five main professions of husbands, which has been multiplied by 100 to ease the readability of the results. Then, the share of each occupation should be interpreted as a percentage. All regressions have month fixed effects and a linear time trend.

a. Mean of the dependent variable during the pre-reform period.

b. Absolute effect of the reform on the dependent variable, measured 3 years after its introduction (half of the post-reform period). The corresponding standard error is reported in parentheses below.

c. Relative effect of the reform, calculated as the ratio of the absolute total effect to the predicted value of the dependent variable, had the reform not been implemented.

Robust standard errors are reported in parentheses. *, **, and*** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Table A1.4: Impact on the composition of new marriages according to main occupation of the wife.

	Five main occupations				
	Professionals	Clerical support	Services	Housekeeping	Sales and trading
	(1)	(2)	(3)	(4)	(5)
Time	0.019***	0.017***	0.016***	-0.047***	0.017***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)
$Post^{2005}$	-0.329*	0.091	-0.284**	0.088	0.104
	(0.176)	(0.163)	(0.143)	(0.127)	(0.129)
Timepost	0.020**	-0.006	0.028***	-0.004	-0.037***
	(0.009)	(0.008)	(0.007)	(0.006)	(0.006)
Share before (%)[a]	24.655	19.969	14.243	11.619	10.969
Effect[b]	0.385	-0.140	0.713***	-0.043	-1.213***
	(0.282)	(0.262)	(0.230)	(0.205)	(0.204)
Change (%)[c]	1.78	-0.99	4.00	-0.25	-13.12
Adj. R^2	0.001	0.003	0.001	0.006	0.001
N	1075700	1075700	1075700	1075700	1075700

Notes: The sample includes all marriages occurred between 2001 and 2007. $Post^{2005}$ is a dummy variable set equal to 1 since July 2005. The dependent variable is a dummy for each of the five main professions of wives, which has been multiplied by 100 to ease the readability of the results. Then, the share of each occupation should be interpreted as a percentage. All regressions have month fixed effects and a linear time trend.

a. Mean of the dependent variable during the pre-reform period.

b. Absolute effect of the reform on the dependent variable, measured 3 years after its introduction (half of the post-reform period). The corresponding standard error is reported in parentheses below.

c. Relative effect of the reform, calculated as the ratio of the absolute total effect to the predicted value of the dependent variable, had the reform not been implemented.

Robust standard errors are reported in parentheses. *, **, and*** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

Table A1.5: Impact on age and civil status of new couples.

	Age at marriage			Civil status at marriage		
	Husband's age	Wife's age	Age gap	Divorced husband	Divorced wife	Both divorced
	(1)	(2)	(3)	(4)	(5)	(6)
Time	0.031***	0.029***	0.001***	0.049***	0.051***	0.020***
	(0.001)	(0.000)	(0.000)	(0.002)	(0.002)	(0.001)
$Post^{2005}$	-0.030	-0.021	-0.008	-0.028	-0.021	0.030
	(0.023)	(0.020)	(0.015)	(0.083)	(0.079)	(0.052)
Timepost	0.009***	0.007***	0.002***	0.037***	0.029***	0.023***
	(0.001)	(0.001)	(0.001)	(0.003)	(0.003)	(0.002)
Average before[a]	32.336	29.989	2.347	8.880	7.959	3.332
Effect[b]	0.300***	0.232***	0.068***	1.292***	1.034***	0.869***
	(0.035)	(0.030)	(0.023)	(0.123)	(0.117)	(0.076)
Change (%)[c]	0.82	0.70	2.12	6.59	5.65	11.48
Adj. R^2	0.034	0.035	0.002	0.015	0.016	0.008
N	1829289	1829289	1829289	1829289	1829289	1829289

Notes: The sample includes all marriages occurred between 2001 and 2007. $Post^{2005}$ is a dummy variable set equal to 1 since July 2005. In columns 1-3, the dependent variable is either the age of spouses or the age gap between them, and is expressed in years. In columns 4-6, the dependent variable is dummy variable for the correspondent characteristic, which has been multiplied by 100 to ease the readability of the results. All regressions have month fixed effects and a linear time trend.

a. Mean of the dependent variable during the pre-reform period.

b. Absolute effect of the reform on the dependent variable, measured 3 years after its introduction (half of the post-reform period). The corresponding standard error is reported in parentheses below.

c. Relative effect of the reform, calculated as the ratio of the absolute total effect to the predicted value of the dependent variable, had the reform not been implemented.

Robust standard errors are reported in parentheses. *, **, and*** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Census of Marriages, National Institute of Statistics of Spain.

2 The Effect of Property Division Laws on Divorce and Labor Supply: Evidence from Spain

2.1 Introduction

Does the legal criterion for the division of matrimonial property in case of divorce influence the behavior of spouses within the marriage? And does it have an effect on the incidence of divorce? In this paper I address these questions by exploiting evidence from a natural experiment in Spain, where different regions have different marital property regimes. I argue that rules regulating the division of joint property in case of marital dissolution are relevant to determine household outcomes. Not only the decision about whether to dissolve the marriage can be influenced by the distribution of rights over family assets in the event of a separation, but also the incentives, and then the behavior, of spouses within the marriage may be affected. Specifically, I study how changes to laws governing the division of family assets at divorce affect the probability of divorce, and for those couples that stay together, their incentives to supply labor in the market.

The intuition behind this relation is quite straightforward. The rule for division of joint assets in case of divorce determines the outside option of spouses, which in turn may affect their bargaining position within the marriage. In the traditional model of the household (Becker, 1981), the distribution of property rights over family assets is irrelevant to determine household outcomes, since the family would re-allocate optimally. However, the literature on family economics seems to have arrived to a consensus about the necessity of treating the household as composed by different members with heterogeneous preferences, resulting in the so-called non-unitary models of household behavior.[1] These models include a wide range of theoretical constructions, but the key point in all of them is that the

[1] See Chiappori and Donni (2009) for a recent review of this literature.

intra-household balance of power matters.

One of the main difficulties of the empirical counterpart of this literature has been to find exogenous sources of variation in bargaining position within the household. I overcome this problem by exploiting a natural experiment given by differences in Family Law across regions in Spain, and two institutional changes that took place during the nineties. The Spanish Civil Code provides a regime of community of property, which is the default regime in all the regions except two (Catalonia and the Balearic Islands). In these two regions, the default regime for all married couples is separation of property. The source of variation in bargaining power comes from two changes to the law in Catalonia. First, in 1993, an economic compensation for the financially weaker spouse in case of marital dissolution was introduced. I argue that this change exogenously and unexpectedly improved the position of the wife within the household. Second, in 1998, the scope of marital contracts was extended, allowing them to include provisions referring to the dissolution of marriage, which was not possible before. In particular, this legal change opened the possibility that a couple could write a contract limiting or even canceling out the economic compensation introduced five years earlier.

I find that the introduction of the economic compensation for the financially weaker spouse in case of divorce led to a reduction in married women labor supply of between 0.6 and 2.5 hours per week. Part of this effect is explained by changes on the extensive margin. The probability of employment for married women fell by about 2 percent when their were favored by the redistribution of rights over marital assets. These effects were partially reversed when marital contracts were allowed to include provisions referring to divorce, since this implied the possibility of limiting or even eliminating the economic compensation introduced before. Indeed, this latter change led to an increase in married women labor supply of around of 1.2 hours per week, and of 2.6 percent in the probability of employment. I also find an increase in marital dissolution after the introduction of the economic compensation. This positive effect was larger in the first years after the reform and then decreased, but remained still positive one decade later.

This research relates to two strands of literature. On the one hand, there are several papers that show that the distribution of power within the family is relevant to determine the final allocation of resources of the household (McElroy and Horney, 1981; Lundberg and Pollak, 1993; Lundberg, Pollak, and Wales, 1997; Chiappori, 1988, 1992; Chiappori, Fortin, and Lacroix, 2002). According to these papers, the household cannot be considered as a unique decision-making unit subject to a unique budgetary constraint in which only total family income matters.

On the other hand, by exploiting variation in divorce law across regions this paper is close to the vast literature on the impact of divorce legislation on several economic outcomes. This literature has mainly focused on the reforms in divorce laws across U.S. states during the 1970's, when many states removed fault as a ground for divorce and almost all of them allowed one of the spouses to file a petition for divorce without the consent of the other. One of the outcomes most often considered is the incidence of divorce.[2] Peters (1986) found that allowing for unilateral divorce does not have any significant impact on divorce rates, a result that was criticized by Allen (1992). Later, Friedberg (1998) found that the divorce rate in states that allowed for unilateral divorce was significantly higher than in other states, and that this legal change could account up to one sixth of the increase in divorce rates in the U.S. during the 1970s and 1980s. However, recently Wolfers (2006) has shown that the increase in divorce rates due to the adoption of unilateral divorce policy was small and faded out within a decade. In an analysis of the impact of different divorce law reforms on the divorce rate in several European countries, González and Viitanen (2009) find that reforms that made divorce easier were followed by significant increases in divorce rates.

[2]Other outcomes are labor supply (Gray, 1998; Stevenson, 2008), fertility (Drewianka, 2008; Alesina and Giuliano, 2006), marriage-specific investments (Stevenson, 2007), implications for children (Gruber, 2004), domestic violence (Stevenson and Wolfers, 2006), and marital formation (Mechoulan, 2006; Rasul, 2006).

Within this literature there are also some papers studying the relation between the rule to divide family assets and labor supply. Gray (1998) evaluates whether the adoption of unilateral divorce law by some states in the U.S. had an impact on married women's labor supply, and finds that this reform had no significant impact unless the underlying marital-property laws in each state are considered. Controlling for these property laws, he finds that the labor supply of wives does appear to respond to their states adopting unilateral divorce statutes and, in particular, that a wife's labor supply is an increasing function of her bargaining position within the marriage. Stevenson (2008) criticizes these results and argues that they are biased due to sample size problems and potentially endogenous controls. Once she accounts for those problems, her results indicate that the incentives provided by unilateral divorce are independent of how matrimonial property is divided. Finally, in a similar setting to the one analyzed here, Kapan (2008) focuses on a House of Lords decision which led to a more equitable distribution of assets between divorcing spouses in England and Wales and finds a negative and significant relationship between married women's bargaining position and female labor supply.

The contribution of this paper is two-fold. First, it contributes to the literature that studies how the spouses' bargaining position within marriage affects their labor supply decisions using a natural experiment in Spain. Moreover, the legal change analyzed here unambiguously improves the position of the wife within the marriage, making it easier to interpret the results from a bargaining perspective. This is important since part of the previous literature cannot disentangle the effect of changes in property division laws from that of unilateral divorce reform, and consequently it is unclear which partner's position is improved after the reform (Gray, 1998; Stevenson, 2007, 2008). One paper that does not suffer from this problem is Kapan (2008). One advantage of the setup studied here over the last paper is the use of within country variation in property division laws instead of cross-country variation, plus a richer legal reform given by two legal changes: the introduction of an economic compensation first and the possibility of eliminating it by means of a contract later. Second, this paper brings new evidence to the debate on how divorce legislation affects (if it does) the incidence of divorce. Moreover, the impact on divorce rates of a legal change like the one analyzed here has never been studied before: a change in the rule for the distribution of assets at divorce without any other change in the grounds for divorce.

The rest of the paper is structured as follows. Section 2.2 presents the institutional background and describes the main reforms to the marital property regime in Catalonia. Section 2.3 presents the theoretical framework in which this analysis is embedded. The data and methodology are described in Section 2.4. Section 2.5 presents and discusses the results and, finally, Section 2.6 concludes.

2.2 Institutional Background

a Divorce Law and Marital Property Regime in Spain

In Spain, the general regulation provided in the national Civil Code may coexist with territorial legislation regarding some specific civil law matters. The area of Family Law is an example of this plurality of norms. The general rules regulating the formation and dissolution of marriage are established at the federal level by the Civil Code; however, the Regional States are left with the right to define their own regulations governing some specific family law aspects, as for instance the marital property regime. This particular set-up configures an interesting case to study how different marital property regimes affect several economic outcomes.

The marital property regime is the set of rules governing the ownership of property during the marriage and the division of it in case the marriage dissolves. In Spain, spouses have the right to choose the property regime by writing a marital contract. If nothing is agreed, the default regime defined in the territorial legislation applies. Two regions have established that, in the absence of marital contracts opting for one particular marital regime, the property of the spouses will be subject to a Separate Property regime (i.e. in case of divorce, property is divided according to who has the legal title). In the rest of regions in Spain, the default rule is the Community Property (i.e. all assets and wealth accumulated since marriage are equally divided between spouses at divorce).

The legal dissolution of a marriage is possible in Spain since 1981, when divorce was reintroduced after four decades of being banned. The divorce law passed in 1981 established a two-step process to deal with marital breakdown. The couple that want to dissolve the marriage should generally resort to a period of separation before being able to file for divorce. Then, the grounds for divorce are closely related to the grounds for legal separation.[3] There are two types of separation: by mutual agreement and based on a legal ground. In the first case, either both spouses or one with the consent of the other can file a petition for legal separation. In the second case, adversary separation occurs when one of the spouses files a petition for separation given that the other has incurred in fault.[4] However, in practice the divorce regime can be considered as close to an no-fault regime, since the Courts have given a loose interpretation of the grounds for separation.[5]

[3]There is one exception in which is possible to directly file for divorce, that corresponds to the case in which there is risk of violence against the spouse or the children. For a more detailed description of the grounds for divorce in Spain during the years under analysis see Boele-Woelki, Braat, and Sumner (2003)

[4]The legal grounds for separation established in the Spanish Civil Code include situations such as unjustified abandonment of the family home, marital infidelity, and abusive or offensive conduct, among others.

[5]According to Boele-Woelki, Braat, and Sumner (2003), the Courts have referred quite often to the so-called "lack of *affectio maritalis*" as a ground for separation, which can be interpreted as the loss of affection between spouses, continuous arguments and reproaches or the existence of a cold and distant relationship between them.

b The reforms to the Regime in Catalonia

Catalonia is the second most populated of the seventeen autonomous communities in Spain, with more than 15 percent of the total Spanish population according to the 2001 Census. It is as well one of the richest regions, occupying the fourth position in per capita GDP as of 2009.[6] During the nineties there were two important modifications to the Catalan marital property regime: the introduction of an economic compensation for the financially weaker spouse and the extension of the scope of marital contracts.

Economic compensation in case of divorce

In 1993, an economic compensation for the financially weaker spouse in case of divorce was introduced in the separation of property regime in Catalonia.[7] The norm established that if one spouse was working during the marriage either for the house or for the other partner with an insufficient economic remuneration or without it, then he or she has the right to perceive an economic compensation from the other spouse in the event of divorce.[8]

In a separation of property regime, this compensation for the financially weaker spouse can be interpreted as a step towards a more equitable distribution of the family property when the marriage breaks up (Lamarca i Marqués, Farnós Amorós, Azagra Malo, and Artigot i Golobardes, 2003).

The amount of the compensation and whether it should be awarded or not

[6] Data taken from the National Institute of Statistics, http://www.ine.es/.

[7] Art 23 of Act 8/1993. The spouse who has been working for the household or for the other spouse, without compensation or with inadequate remuneration, is entitled to receive, when the marriage ends by legal separation, divorce or annulment, an economic compensation if for that reason a disequilibrium has been generated between his or her assets and those of the other spouse.

[8] It is worth mentioning that this compensation is compatible with any other economic rights to which the favored spouse may be entitled to at divorce, such us alimony payments for instance.

is decided by the judge intervening in the dissolution of the marriage. According to Lamarca i Marqués, Farnós Amorós, Azagra Malo, and Artigot i Golobardes (2003), between 1993 and 1998 many claims for the compensation were either denied or received relatively little amount of money. This institution gained importance in the Catalan marital property regime after some landmark decisions by the Catalan Supreme Court of Justice regarding the criteria to apply the norm, the first in October 1998.[9] In fact, this strengthening of the economic compensation within the marital property regime in Catalonia is largely related to its introduction into the Family Code of Catalonia in 1998 (Act 9/1998).[10] In Catalonia, the Family Code is a norm of considerable practical relevance to deal with family law matters.

Scope of marital contracts

A second reform to the marital property regime in Catalonia occurred in 1998, when the scope of marital contracts was extended to allow their use, not only to organize the economy of the family, but also to liquidate it.[11] That is, since 1998 marital contracts can contemplate the possibility and the consequences of a potential crisis in the marriage.

[9] In those interventions the Supreme Court stated clearly that "always when one spouse works for the house or for the other without a retribution, it generates an (unfair) enrichment in favor of the other spouse", and "to award the economic compensation to one spouse their assets should be compared", being the difference between them the basis to calculate the amount of compensation (Lamarca i Marqués, 2003).

[10] Art 41 of Act 9/1998. Economic compensation on the grounds of work: In cases of judicial separation, divorce or marriage annulment, the spouse who has worked for the household or for the other spouse without receiving any payment in exchange or who has received insufficient payment, shall be entitled to receive economic compensation from the other spouse, in the event that this fact has produced a situation of inequality between the two patrimonies, which implies an unfair enrichment.

[11] Art 15 of Act 9/1998. In marital contracts, it is possible to determine the matrimonial economic system, agreements on inheritances, make donations and establish licit stipulations and pacts that are deemed convenient, even in anticipation of a marriage break- up.

In general, marital agreements are legal instruments that allow the spouses to make contracts about issues regarding the matrimonial property regime. They are different from the more usual pre-nuptial agreements in the sense that it is possible to write them not only before the marriage but also during it, and even after a possible separation.

Before 1998, marital contracts were a valid contracting instrument for the period during the marriage, but once the marriage was dissolved, this contract lost its legal validity. After 1998, marital contracts can be enforceable even after the couple divorces. This implies that it is now possible for the spouses to contract about economic transfers between them after a potential divorce. Specifically, this opens the possibility that spouses write a marital contract establishing conditions related to the economic compensation for the financially weaker partner (i.e. limiting or even eliminating it).

Figure 2.1 shows the number of marital contracts signed in Catalonia and in the rest of Spain between 1988 and 2002. As can be seen, while in the rest of Spain the annual number of contracts grew steadily during the whole period, in Catalonia there was a huge increase after 1998, when those agreements were allowed to contemplate the dissolution of the marriage. This seems to support the hypothesis that the new contracting behavior among Catalan couples is directly associated with the two reforms to the marital property regime: the introduction of the economic compensation in 1993 and the possibility to make legal agreements related to it since 1998. That is, although the content of contracts is private and consequently, unobservable, the fact that the only legal change in 1998 is that they can include provisions regarding divorce seems to be an important factor explaining the rise in the number of agreements. For instance, since 1998 a couple can write a contract agreeing that in the case of dissolution of the marriage, none of them has the right to claim the economic compensation established in the marital property regime since 1993, independently of the financial situation of each of them.

The number of marital contracts relative to the annual number of marriages can give us an idea of their quantitative importance. In the rest of Spanish regions, the ratio of contracts to hundred marriages goes gradually from 11.8 in 1988 to 39 in 2002. In Catalonia, this ratio remains quite constant around an average of 1.7 annual contracts per hundred marriages until 1998, and increases sharply to reach the figure of 12.5 in 2002.[12] Thus, although the increase in the number of contracts in Catalonia is important, the fact that they represent a small proportion of new marriages seems to indicate that their impact on the stock of married couples should be observed only gradually.

2.3 Theoretical Framework and Expected Outcomes

The introduction of an economic compensation in case of divorce redistributes the rights over total marital assets between spouses. Thus, it modifies the nature of the marriage contract by changing the value of the option outside of marriage for both of them. Assuming the wife is the financially weaker spouse, this legal change implies a redistribution of wealth towards her and against the husband. This is a valid assumption insofar as household assets are disproportionately held in the husband's name (Chiappori, Fortin, and Lacroix, 2002; Gray, 1998). Also, it is supported by the evidence provided by court cases regarding the economic compensation, which shows that in almost all cases this compensation is claimed by the wife.[13]

On the other hand, the possibility of writing marital contracts including provisions referring to a potential end of the marriage, makes it possible

[12] Ideally, it would be better to calculate the ratio of annual contracts to the number of marriages that involve certain level of wealth (i.e. poorer couples have less incentive to enter into a marital agreement), but this information is not available.

[13] Information available from the legal service *Westlaw* for Spain, which provides case law information about all decisions from the Superior Court of Justice, Provincial and National Hearings and the most interesting decisions from lower courts.

to use them to restrict or even eliminate the compensation. By the same logic used before, this is expected to have an opposite effect on the relative position of married women within the household. However, we should distinguish between the effect on existing couples at the moment of the reform and the effect on couples formed afterwards. In the case of existing couples, wives whose bargaining position have been enhanced by the compensation would have no incentive to enter into a contract that restricts this benefit. But for new couples, both the compensation and the possibility to modify it by means of a contract are in force at the moment they make their marriage decision. Then, we expect that the effect of the modification of the scope of marital contacts on the intra-household balance of power occurs through changes in the marriage market. Although I will address this issue in the following paragraphs in more detail, the key point here is that a reform that only affects the flow of new couples will have effects on the stock of married people that are noticeable only gradually.

The next two sections benefit from existing theoretical models in the literature of household economics to derive predictions for the effects of the reforms to the marital property regime on the probability of divorce first, and on the labor supply of intact couples, later.

a Marital Dissolution and Formation

The key point is the distinction between the effects of the reforms on the existing stock of married couples (what in the literature is called a "pipeline effect") and the effects on couples formed under the new regime (a "selection effect")(Rasul, 2006; Mechoulan, 2006; Matouschek and Rasul, 2008). When the economic compensation was introduced into the Catalan marital property regime in 1993, there was an unexpected redistribution of wealth and assets within the household. For those marriages that are close to the brink of divorce, the favored spouse, whose utility outside the marriage has increased given the higher share of the assets she would be entitled to in case of separation, may want to end the marriage.[14] Then, this incentive effect will affect existing couples by increasing the likelihood of marital dissolution in the population.[15] Moreover, since this is an effect on the stock of existing couples, it could be quantitatively important in the short run.

On the other hand, for those couples that get married under the new regime there is a selection effect, that could affect the composition and the quality of new matches. With regard to the composition, under the new regime we expect fewer matches between heterogeneous partners in terms of wealth, in particular between 1993 and 1998, when the economic compensation was in force and contracts could not contain provisions regarding divorce. With regard to the quality of new matches, the economic compensation is expected to foster cooperation between spouses and investments in marriage-specific capital, leading to a reduction in the probability of marital breakdown (Stevenson, 2008).

The second reform, the extension of the scope of marital contracts, should not have, in principle, an effect on the probability of divorce. Since I argued before that this reform is less relevant for existing couples, the main effect should come through impacts on the marriage market. To think of the potential effect of this reform on divorce intensity through changes in the marriage market, we need to ask how the selection into marriage would change as a consequence of the reform. In other words, which couples that would not have married under the contracting rules before 1998, are willing to do it after the legal change? These could be couples characterized

[14] An implicit assumption here is that there is not perfect Coasian bargaining. The strict application of the Coase theorem would lead to the prediction of no changes in the incidence of divorce, since spouses would bargain to reach the efficient outcome. There is, however, enough empirical evidence suggesting that the assumptions in which this theorem is based are not realistic (Peters, 1986; Stevenson and Wolfers, 2006).

[15] This reasoning is based on the assumption that the divorce regime in Spain during the period of analysis can be considered in practice as a no-fault regime. The lack of "affectio maritalis" as an accepted and widely used ground for separation makes this a reasonable assumption. See for instance Boele-Woelki, Braat, and Sumner (2003) and the legal literature cited there.

by more wealth heterogeneity, that is, couples in which the richer partner would not want to risk his or her assets in the case of separation after marriage. These couples can now marry and write a contract agreeing upon the distribution of assets in case of separation. However, there is no reason to expect a different probability of divorce for those couples. Therefore, we expect no effect of the extension of marital contracts on the probability of divorce.

To sum up, the introduction of the economic compensation into the marital property regime in Catalonia 1993 is expected to have a positive impact on the probability of marital dissolution, as a consequence of the change in incentives for existing couples. This effect would tend to fade out as the composition of new matches changes due to a selection effect through the marriage market. Also, the modification of the scope of the contracts in 1998 is expected not to have an (independent) effect on divorce rates.

b Intra-household Allocation and Labor Supply

We should distinguish again between the effects on existing couples from the effects on individuals not yet married. As mentioned, the introduction of the economic compensation into the Catalan marital regime redistributed family wealth in favor of the wife. While this could have led to more divorces, it could have affected intact marriages as well. According to the collective model of the household (Chiappori, 1988, 1992), a reform like this improves the bargaining position of the wife, and then shifts each spouse's commodity and time use to more strongly reflect her preferences.[16]

The theoretical link between the intra-household bargaining position and labor supply is provided by Chiappori, Fortin, and Lacroix (2002). They

[16] It should be noticed that the prediction would be entirely different under the so-called unitary model to household modeling Becker (1981). That model is based on the assumption that household members act as if they maximize a unique utility function under a common budget constraint, which implies that the distribution of property rights within household is irrelevant to determine household outcomes.

show that a redistribution of family wealth in favor of the wife would be equivalent to a higher share of non-labor income allocated to her. Then, to the extent that spousal labor supply is responsive to income, standard income effects should, all else equal, lead to a reduction in female labor supply and an increase in male labor supply. The total effect on husband's labor supply is less clear, since a substitution effect operates in the opposite direction.

The modification of the scope of the contracts in 1998 could have affected married people labor supply mainly through changes in the marriage market.[17] We expect that this new reform deteriorated the position of wives in those marriages formed after 1998, which according to the collective model of the household would lead to an increase in their labor supply.

2.4 Data and Identification Strategy

a Data

The data for the estimation of the impact on marriage dissolution rates come from the administrative registries of judicial statistics. These data are comprised of the total number of marital dissolutions (divorces, separations, and marital annulments) at the region level, from 1990 to 2004.

In the estimation of the impact on labor supply I will use data coming from the Spanish Labor Force Survey (*Encuesta de Población Activa*), covering all quarters since 1990 and until 2002. This survey is carried out every quarter by the Spanish National Institute of Statistics on a sample of some 60,000 households, and it is designed to be representative of the Spanish

[17] As mentioned before, although contracts can be written at any moment, wives whose balance of power within the household has been improved due to the economic compensation would not have incentive to restrict that benefit by means of a contract. And since the two spouses have to agree to write a contract, we reasonably can expect that this reform did not affect existing couples in an important way.

population. The survey has a rotating scheme by which in each quarter one sixth of the sample is renewed, so households are expected to be in the survey for six quarters.[18]

b Econometric Specification

To study how the rules governing the division of property at divorce influence household outcomes this paper benefits from a natural experiment in Spain, given by the regional variation in marital property regime across Spanish regions. Unexpected and exogenous law changes in some regions but not in others are an ideal source of variation for the estimation of causal effects. In this paper, I take advantage of the main modifications to the Catalan Family Law during the nineties to identify variation in the bargaining position of spouses within the household.

As noted earlier, the economic compensation for the financially weaker spouse in case of divorce was introduced in the marital property regime of Catalonia in 1993. Later, in 1998, a new law change extended the scope of marital contacts, allowing them to refer to the consequences of marital dissolution, which I argued made it possible to use these legal instruments to limit or even eliminate the economic compensation for the financially weaker spouse.[19]

[18]The data are available in two formats: (i) the cross-sectional dataset, and (ii) the longitudinal dataset. The latter has the advantage of including a unique identification code for each individual that allows to match observations from quarter to quarter. However, the former dataset is richer in information since some key variables are dropped from the panel dataset (e.g. the household identifier and the region of residence, two key variables for this study, are some of the variables missing). To overcome these difficulties, I match both datasets in a way that allows me to have all the information included in the cross-sectional dataset plus the individual code to identify individuals over time. This is done using only information contained in both datasets, and as a result of the procedure employed to perform the matching 100 percent of the observations are matched correctly.

[19]It should be remembered that the legal modification to the Catalan Family Code in 1998 encompassed also the introduction of the economic compensation into this legal norm. Given the importance of the Family Code in Catalonia as a systematization in

Then, the natural experiment to exploit here is to analyze whether the introduction of the economic compensation first, and the modification of the contents of marital agreements later, had any impact on household outcomes in Catalonia, using as a control group individuals from the rest of Spanish regions where the community of property regime is the norm and where there were no relevant legislative changes during the period.[20]

In cases like this in which there is only one region treated and several untreated regions that potentially could be part of the control group, there is always the question of which regions conform an adequate control group. I will follow the criterion of including in the control group only those regions with similar trends in the outcome to Catalonia during the pre-treatment period. This is because difference in differences is a valid identification strategy only if the treatment and control groups have similar trend in the outcome of interest in the pre-treatment period (Galiani, Gertler, and Schargrodsky, 2005; Heckman and Hotz, 1988).[21]

Therefore, in the regression results below the control group is selected according to the following criterion. The outcome of interest, y_{rt}, is regressed during the pre-treatment period ($t < 1993$), on a linear time trend, a full set of dummies for all seventeen Spanish regions (μ_r), and the interactions of those dichotomic variables with the linear trend. That is:

$$y_{rt} = t + \sum_r \mu_r + \sum_r \mu_r * t + u_{rt} \tag{2.1}$$

one legal body of all norms regarding family law, this introduction could have had an additional effect on household outcomes of an opposed sign to the one predicted for the extension of the scope of marital contracts.

[20]There is one exception to this given by the fact that in Balearic Islands the default system is the separate property regime. However, since there was not any relevant change in this regime during the period of analysis, it will form part of the control group.

[21]Alternatively, I run all the regressions including all the remaining regions as the control group, and the main results are mostly the same.

Setting Catalonia as the omitted category, only regions with coefficients in the interaction term not significantly different from zero are selected and included in the control group.

Divorce Probabilities

The data to estimate how changes to the marital property regime affect the incidence of divorce come from judicial statistics. I use administrative information on the number of marital dissolutions aggregated at the region level for the period since 1990 to 2004. The sample period starts in 1990 to have four years of data before the treatment (the first law modifying the property division regime was applicable since the end of 1993, so the treatment indicator equals 1 since 1994 onwards) and it is truncated in 2004 to avoid obtaining results that may be confounded with the effects of another important law passed in 2005, which modified the grounds for divorce.

The dependent variables in the analysis are the divorce rate and the separation rate, defined as annual divorces or separations per thousand people, respectively. To account for pre-existing differences across regions in the level of marital dissolution I include region fixed effects in the regressions. Also, given that the control group is conformed by regions with the same linear trend in the rate of dissolution, it is not necessary to control for unobservable factors that may induce region-specific linear trends. Then, the two equations to estimate are the following:

$$d_{rt} = \beta_1 cat * post93 + \beta_2 cat * post98 + \sum_r \mu_r + \sum_t \lambda_t + u_{rt} \tag{2.2}$$

$$d_{rt} = \sum_d \beta_d cat * year_{td} + \sum_r \mu_r + \sum_t \lambda_t + u_{rt} \tag{2.3}$$

where d_{rt} refers to either the divorce rate or the separation rate, μ_r and λ_t represents region and year fixed effects, respectively. Variables *post*93 and *post*98 are binary variables set equal to one for the period after the property division regime was modified in Catalonia, while *cat* is another dummy variable set equal to one for Catalonia. Hence, the coefficients of the interaction terms of those variables measure the impact of the reforms on marital dissolution rates. That is, β_1 should be interpreted as the average change in the dependent variable due to the legal change in 1993, while β_2 is the average change in the dependent variable attributable to the legal change in 1998. Notice that, given the definition of the variables *post*93 and *post*98, the net impact of the reforms after 1998 is given by the summation of the two coefficients β_1 and β_2.

Equation 2.3 differs from equation 2.2 in that it allows for dynamic effects of the reforms. Wolfers (2006) states that this type of specification is preferable when the reform is expected to have initially a large effect (i.e. due to a "pent-up" demand for divorce in this case), but the long run effect may be negligible. So $year_{td}$ is a vector of dummy variables that equal 1 if the (first) reform has been effective for d years at time t.[22]

Labor Supply

The main empirical strategy is again to compare changes between Catalonia and the regions included in the control group in the labor supply of wives and husbands before and after the reforms.

The main dependent variable is the number of usual hours worked per week reported by married individuals, including the zeros. The data for the estimation of the impact of the reforms on labor supply come from the Spanish Labor Force Survey. First I use the pooled cross sections of all quarters from 1990 to 2002 to estimate both an OLS regression and a Tobit

[22]Same as in Wolfers (2006), I combine years into two-year groups: one dummy for the first two years after the reform, another for the next two years, and so on.

specification. The reason why we need a tobit specification is the inclusion of the zeros in the hours regression. Therefore, the main specification for the pooled cross-sections is the following:

$$h_{irt} = \beta_1 cat * post93 + \beta_2 cat * post98 + x_{it}\delta + \sum_r \mu_r + \sum_t \lambda_t + u_{it} \quad (2.4)$$

where $post93$ and $post98$ are two binary variables defined the same as before, cat is a dummy variable for Catalonia, and x is a set of control variables.

Since hours of work is a non-negative random variable that equals zero for some fraction of the sample, the difference in hours across treatment groups can be decomposed in two parts: the difference in the probability of being employed (participation effect), and the difference in hours conditional on employment.[23] We may be interested in analyzing how these reforms change the probability of employment.[24] Then, I use the linear probability model to estimate the same equation but with a binary indicator for employment status in the left-hand side:

$$e_{irt} = \beta_1 cat * post93 + \beta_2 cat * post98 + x_{it}\delta + \sum_r \mu_r + \sum_t \lambda_t + u_{it} \quad (2.5)$$

The coefficient of interest are the interaction terms β_1 and β_2, which are interpreted as the average change in the usual number of hours worked per week attributable to the reforms in 1993 and 1998, respectively. Again, the

[23] See Angrist and Pischke (2008) for the details.

[24] The second part, the difference in hours conditional on participation, has no special interest, since it does not have a causal interpretation. Angrist and Pischke (2008) show that the treatment changes the composition of the group with positive working hours resulting in a kind of selection bias.

variable *post*93 is set equal to 1 over the whole period after 1993. This means that β_1 should be interpreted as the average impact of the reform in 1993 for the rest of the estimation period, while β_2 should be interpreted as the additional impact of the law change in 1998. Then, adding these two coefficients we would obtain the net impact after 1998. The coefficient β_1 is expected to be negative for married women (both the income and the substitution effects go in the same direction) and could be either positive or negative for married man (income and substitution effects have opposite signs). For the same reasons, the coefficient β_2 is expected to be positive for married women and negative for married men.

The control variables included in the regressions are a second order polynomial in age, a set of educational attainment dummies, a dummy for being in school, the regional unemployment rate and per capita GDP to control for business cycles, age and education of the spouse, and dummies for different quarters to control for seasonality. In addition, in all specifications region and time fixed effects and included.

I select a sample of married individuals aged between 30 and 50 years old in order to better capture labor supply decisions as a consequences of intra-household bargaining and avoid the confounding effects of both education related decisions of younger individuals and also earlier retirement decisions of older people. I also restrict the sample to those individuals that are observed along the six interviews and whose marital status is unchanged over that period. This allows me to focus the attention on the impact of the reform on the spouses labor supply as a consequence only of changes in their bargaining position within the household (i.e. all individuals whose marriages break down within the six quarters span are dropped from the sample).

To fully take advantage of the available data, I also run the same equations including fixed effects at the individual level. The gain of including individual fixed effects is to control for differences in unobservable characteristics between individuals in the treatment and in the control group. Then, the equations to be estimated are similar to equations 2.4 and 2.5 but including individual fixed effects(δ_i):

$$h_{irt} = \beta_1 cat * post93 + \beta_2 cat * post98 + x_{it}\gamma + \sum_r \mu_r + \sum_t \lambda_t + \delta_i + u_{it} \quad (2.6)$$

$$e_{irt} = \beta_1 cat * post93 + \beta_2 cat * post98 + x_{it}\gamma + \sum_r \mu_r + \sum_t \lambda_t + \delta_i + u_{it} \quad (2.7)$$

Finally, an important concern regarding the correlation between the reforms and spousal labor supply is that it could be driven by some other (unobserved) socioeconomic factors, different from changes in the bargaining position of the spouses. Then, as a robustness check, I perform a Difference-in-Difference-in-Differences analysis, using single individuals as an additional control group. That is, controlling not only for changes in labor supply in the rest of Spanish regions, but also for changes in the labor supply of unmarried individuals (i.e. people similar to the treatment group who should not be affected by the policy), it is possible to rule out the effect of factors that could be potentially correlated with the two variables of interest.[25]

[25] Some papers in the literature try to solve this by doing placebo tests. That is, they look at the effects of the reform on groups that should not be affected. For instance, Chiappori, Fortin, and Lacroix (2002) propose to test wether the change in the divorce rule had an impact on the labor supply of single individuals, who are not supposed to be affected if the collective model is the true explanation.Stevenson (2008) performs a similar placebo test and finds a significant impact of the changes to unilateral divorce on the labor supply of single women, which she attributes to some anticipation effect.

2.5 Empirical Results

a Impact on the Divorce Rate

The potential impact of rules governing the division of joint property upon divorce on marital dissolution has special interest for two reasons. First, if the modification of the rules affecting the division of matrimonial property has an impact on aggregate dissolution rates, it could also have an impact on the labor supply of married individuals, to the extent that labor supply decisions of married people are sensitive to the probability of divorce. Moreover, insofar as marriages that break down because of the law change are not randomly selected from the pool of marriages, there could be a compositional effect that will alter married individuals' labor supply as well. Second, whether a change in divorce legislation has a causal effect on divorce rates is an interesting question in itself that has generated a long and still open debate. Thus, bringing new evidence to this literature is an important contribution, particularly since the impact of a change only in the rule governing the division of joint assets without any other modification in the grounds for divorce has never been studied before. This literature has focussed on the effects of unilateral revolution in the U.S. (Peters, 1986; Allen, 1992; Friedberg, 1998; Wolfers, 2006) and of norms that made divorce easier in Europe (González and Viitanen, 2009). To the best of my knowledge, this is the first paper asking whether a change in property division laws at divorce can have an impact of the probability of marital dissolution.

Figure 2.2 shows the evolution of the divorce rate by region, measured as annual divorces per thousand people. We can notice an increase (both over the trend and in comparison to the rest of Spanish regions) in the annual number of divorces per 1000 people between 1993 and 1998 in Catalonia, when the financially weaker spouse can claim an economic compensation which can not be restricted or eliminated by a marital contract. After 1998 the evolution of the annual number of divorces in Catalonia seems to be similar to that of the rest of Spain, although there seems to be a broader gap in levels between the two groups. Although this graphical evidence seems to point to an increase in the number of marriages that break up between 1993 and 1998 in Catalonia, there is still room for an explanation related to the two-step process that the dissolution of a marriage requires in Spain. That is, given that when the economic compensation was introduced in 1993 there was a stock of separated but not divorced people, the apparent increase in the number of divorces in Catalonia could be just an advance of divorce proceedings of couples already separated, without any change in the number of marriages breaking up. To test this hypothesis, Figure A2.1 in the appendix shows the evolution of the annual number of legal separations in the same two groups and during the same period of time. We can observe an increase in the separation rate in Catalonia with respect to the rest of Spain between 1993 and 1998, a behavior compatible with the hypothesis that changes in marital property regime have an effect on the incidence of marital dissolution. I explore more formally this conjecture with the regression analysis that follows.

Table 2.1 reports the estimates for equations 2.2 and 2.3, and for the two dependent variables under analysis, divorce and separation rates. The specification in column 1 shows the average impact of the two reforms to the marital property regime in Catalonia during the period under analysis. We can see that the coefficient of the interaction term $cat * post93$ is positive and statistically significant, while the coefficient of the interaction term $cat * post98$ is also statistically significant but negative and lower in magnitude. This suggests that the introduction of the economic compensation had a positive impact on the divorce rate in Catalonia, an effect that is only partially reversed after 1998, when marital contracts can refer to the consequences of a crisis in the marriage. It should be noticed that the net effect of the two reforms on the number of divorces is still positive after 1998. This implies that the possibility of imposing limits to the economic compensation by means of a contract have mitigated, but not eliminated, the positive impact of that institution on the incidence of divorce. Regarding the magnitudes of the estimates, given an average of 0.933 divorces per thousand people in Catalonia, the increase in divorce rates by about 0.123 between 1993 and 1998 translates to a increase of about 13 percent in annual divorces than can be explained by the economic compensation. The effect after 1998 is sill positive and equal to $0.123 - 0.05 = 0.073$, and it implies that the average divorce rate in Catalonia remains about 8 percent higher due to the combined effect of the two reforms.

Column 2 shows the results obtained with the more flexible specification given in equation 2.3. We can see that all the coefficients for the two-year periods after 1993 are positive and statistically significant. They show, however, an interesting dynamic pattern for the divorce rate in Catalonia as a consequence of the reforms. The impact of the economic compensation reached its maximum three or four years after its introduction into the Catalan marital property regime, and then started to decrease. It remained positive, however, during the whole period of analysis.

There are at least two possible explanations for this behavior. One is the existence of a "repressed" demand for divorce. When the compensation improved the outside option for some spouses whose marriages were on the brink of divorce, they decided to dissolve their relationships. The maximum effect is obtained three to four years after the legal change, somehow expected given the normal delay of divorce proceedings (in particular given the two-step process required). A second explanation is that the selection into marriages is playing a role. The law change may have induced more homogeneous marriages in Catalonia, reducing the incidence of divorce some years later. Given the short period of time we are referring to, I do believe that the first explanation is more likely to be driving this behavior in the divorce rate.

Columns 3 and 4 present the same two specifications but using annual number of separations per thousand people as the dependent variable. The results somehow confirm the conclusions obtained by analyzing the response of number of divorces. There is an increase in the number of separations in Catalonia after 1993, which can be attributable to the inclusion of the economic compensation into the Catalan marital property regime. A subtle difference when looking at separations instead of divorces, is that the partial reversion after 1998 of the initial jump in 1993 in the number of couples filing for separation is not statistically significant. The coefficient of the interaction term $cat * post98$ is negative but insignificant. The magnitude of the effect is similar to the one obtained when using divorce rate as the dependent variable. An increase of 0.168 in the average annual number of separations after 1993, in terms of an average of 1.2 separations per thousand people in Catalonia before the law change, is equivalent to a 14 percent increase. Finally, the last column presents the result when the dynamic response of the separation rate is explicitly taken into account. We can derive the same conclusions than in the case of the divorce rate. There is an increase in the number of annual separations in Catalonia after the introduction of the economic compensation that reaches its maximum about three or four years after the reform. After that, this impact seems to fade out but remains positive during the whole period under analysis.

Overall, the main conclusion seems to be that the introduction of the economic compensation generated a significant (both statistically and economically) increase in marital dissolution rates, that can be partially explained by couples already separated advancing their divorce proceedings, but also by some new marital breakdowns as a consequence of the new rules of the game.

b Impact on Labor Supply

Figure 2.3 shows the evolution of weekly hours worked by married women in Catalonia and the rest of Spanish regions.[26] We can notice an increasing trend in working hours for both groups during the whole period, with a more or less constant difference in levels of about two hours in favor of Catalonia.

[26] Non-employed married women are included with number of working hours set to zero, in order to capture adjustments in both the extensive and the intensive margins.

Main Results

The main results of the labor supply reduced-form regressions for married women are summarized in Table 2.2.[27] The period of analysis goes from 1990 to 2002, including data from all quarters. I choose this period of time in order to include four years before the first reform and four after the second reform. The estimation sample is restricted to married women between 30 and 50 years of age who have been interviewed during six consecutive quarters.

Columns 1, 3, and 4 report the results when the dependent variable is the number of hours per week a married woman works in the market, while columns 2 and 5 report the results when the binary variable for employment status is on the left-hand side. The two coefficients of interest are those of the interaction terms $post93 * cat$ and $post98 * cat$, which give the average change in female hours of market work attributable to the reforms to the marital property regime in Catalonia. All the specifications shown in the table control for differences in levels of the dependent variable that are constant across regions during the sample period by including region fixed effects, and also for differences across time that are common for all regions, by including year fixed effects. Also, reported standard errors are clustered at individual level to account for the presence of correlation within individuals over time (Bertrand, Duflo, and Mullainathan, 2004).

In all specifications the coefficients have the expected sign. The introduction of the economic compensation for the financially weaker spouse, usually the wife, is expected to cause a reduction in married women's labor supply, while the possibility of diminishing or eliminating this compensation by mean of a contract is expected to have the opposite effect. The first column shows the results when the labor supply equation is fitted with an OLS criterion. The coefficient of $post93 * cat$ is negative and statistically significant (at 10 percent level), and equal to -0.688, while the variable $post98 * cat$ has a positive and significant coefficient of 1.288. This means that (married)

[27]Full regression results are available from the author upon request.

women reduced their labor supply when they were entitled to a higher share of marital assets in case of divorce by less than one hour per week, but this effect reversed when contracts were allowed to have provisions about the situation after divorce. The net effect after the two reforms according to the OLS specification is an increase in female labor supply by less than one hour (-0.688+1.288=0.6) per week. Column 3 reports the coefficients when the equation is fitted with a Tobit model. This model is preferred over OLS given the large numbers of zeros in hours worked. In this case, the introduction of an economic compensation reduced married women labor supply by about 2.5 hours per week, and this effect was not reverted by the reform to the scope of the contracts (the coefficient is plus 1.127 but insignificant).

We want to know as well to what extent the average change in hours worked comes from changes in the extensive margin (i.e. changes in participation into employment). The estimates in column 2 indicate that indeed part of the response comes through changes in the extensive margin. The probability of being employed for married women fell by 1.8 percent as a consequence of the economic compensation, but increased by 2.6 percent following the modification of the scope of marital contracts.

So far, the benefit of having more than one observation per individual over time is that we can focus on married people who continue married during the six-quarter windows of the survey. In this manner, we can concentrate our attention on the relationship between rules for division of marital property and the labor supply decision of intact marriages, interpreting the results as arising from changes in the intra-household balance of power.

Nevertheless, the main advantage of having more than one observation per individual is that it allows us to control for unobserved (fixed) effects. Even though, it should be kept in mind that this is a short panel (only six observations per individual covering a period of one year and a half), and the result would refer to the very short term impact of the reform. That is, given that the identifying variation for each coefficient comes from a discrete policy shock, results are determined by the variation in the dependent variable around the policy change (to be more precise, from changes in the number of hours that are not further away from each reform than 5 quarters). Then, columns 4 and 5 report the results of the panel estimation with individual fixed effects for working hours and employment status, respectively. Looking at hours, we can see again a negative and statistically significant impact of the economic compensation on married women's labor supply of about half an hour per week, and no impact of the extension of the scope of the contracts. And similar results can be derived from changes in probability of employment: the economic compensation had a significant negative impact in female employment, while the extension of the scope of marital contracts in 1998 had no significant effect.

Finally, there is one additional result that is worth mentioning. According to the theoretical framework, the introduction of the economic compensation should affect the stock of existing couples at the moment or the reform, while the change in the scope of the contracts is expected to have an impact more on the flows than on the stock, through changes in the marriage market. That is, while for the first reform we expect to find an almost immediate effect, for the latter, if there is any effect, it should appear more gradually. The estimation that exploits the longitudinal aspect of the data is an interesting manner of testing this hypothesis, given the short time dimension of the panel. In fact, the results reported in the last two columns are compatible with the expected timing of the reforms: the reform that theoretically should have affected the stock of married women (economic compensation) had a statistically significant impact on labor supply in the very short run, while the change whose effects are expected to be noticeable only gradually (scope of marital contracts) had no significant effect within the first six quarters after the reform.

The results of the regressions for married men are reported in the Appendix (Table A2.3). The specifications that exploit the repeated cross-section version of the sample yield coefficients for the variables of interest that are of small magnitude and not statistically different from zero. These results are consistent with the abundant empirical and theoretical evidence that shows a low labor supply elasticity for males of prime age.The panel estimation shows a negative and significant coefficient for married men after the introduction of the economic compensation in 1993, and again a negative and significant coefficient after the reform to the scope of the contracts in 1998.

Groups more affected by the reform

The main results showed above support the main prediction regarding the relationship between rules for division of property at divorce and female labor supply: the higher the share of marital property that goes to the wife, the lower her labor supply. This relationship is expected to be stronger for couples with higher level of assets and wealth, although unfortunately, there is no direct information in the data about how wealthy a family is that can be used to test this hypothesis. The data contain, however, information on variables that could help identifying those couples with higher levels of assets. One of these variables is the type of job the husband is performing. If we assume that there is a correlation between being the owner or manager of a firm and the level of marital assets, we can use this information to proxy the level of wealth of the family. Table 2.3 show the results of labor supply regressions for wives whose husbands declare in the survey to be the owner or manager of a firm. As expected, we can see larger impacts of the reforms to the Catalan marital property regime on married women in wealthier couples.

c Robustness Checks

Table 2.4 reports the results of the triple difference estimation for married women, where the impacts of the reforms to the marital property regime on labor supply are estimated controlling not only for changes in the labor supply of married women in the rest of Spain, but also for changes in the labor supply of unmarried women in all regions. Then, the coefficients of interest are those of the interaction terms between the dummies for the periods after each reform and being a married women residing in Catalonia. As we can see, the sign for the coefficient of $post93 * cat * marriedw$ remains negative for the introduction of the economic compensation in all specifications, but it is only statistically different from zero (at 10 percent level) in the regression for hours worked when individual fixed effects are controlled for (column 4). The coefficient of $post98 * cat * marriedw$, on the other hand, has no stable sign and is always insignificant.

2.6 Conclusion

In this paper I analyze empirically the relationship between the rules governing the division of marital property in case of divorce and two economic outcomes of the household: the incidence of marital dissolution and the labor supply decision of intact couples. The rule for the division of marital property at divorce is important because it determines the outside option of spouses, and consequently, their balance of power within the household. According to non-unitary models of the household, the relative bargaining position of spouses matters for the household decision-making process. I argue that unanticipated changes in this rule may alter the incentive of certain couples to dissolve their marriage as well as the labor supply decision of couples that stay together.

The variation in family law across regions in Spain, where different regions have different marital property regimes, offers an ideal setting to study this. While in the majority of regions the default marital property regime is the community of property, in two regions, one of them is Catalonia, the default rule is separation of property.[28] Moreover, the Catalan regime was modified twice during the nineties. First, in 1993 an economic compensation for the financially weaker spouse in case of divorce was introduced, which can be

[28]This refers to the period analyzed in this paper. The Valencian Community modified its marital property regime in 2008, adopting a separation of property rule.

interpreted as a step towards a more egalitarian distribution of marital assets in case of divorce. Second, 1998 the scope of marital contract was extended allowing them to contemplate the dissolution of the marriage, which gave them the freedom of agreeing about how to divide the assets if the marriage breaks up. These two legal modifications can be seen as sources of exogenous variation in spouses' relative bargaining power within the household that can be used address the questions stated before.

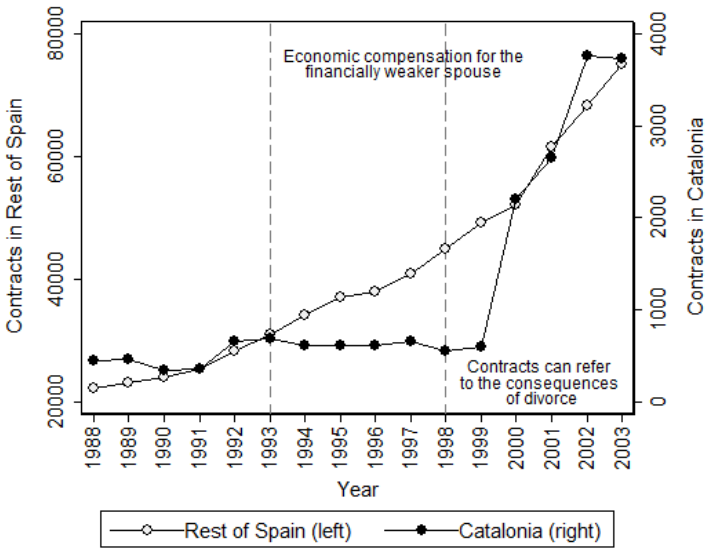
I find that the introduction of the economic compensation increased divorce rate in Catalonia by about 13 percent, and part of this effect is reversed after 1998, when contracts can contemplate the possibility of divorce. The net effect remained positive and close to 8 percent until one decade after the first reform. Looking at the dynamics of the response in divorce rates, I find that the impact of the economic compensation reached it maximum between three and four years after its introduction, and then started to decrease. Similar results are obtained when the analysis is performed with separation instead of divorce rate. The results suggest that, although part of the increase in the incidence of divorce can be explained by couples already separated when the reform took place that advanced their divorce proceedings, there was an increase (both statistically and economically) in marital dissolution rates.

I also find that intact couples were affected by the reforms. Wives entitled to a higher share of family assets in case of divorce reduced their labor supply between 0.6 and 2.5 hours per week, depending on the specification. Looking at the effect on the extensive margin, I find a reduction in the probability of employment for married women of about 2 percent that can be attributable to this redistribution of rights over marital assets. These effects are reversed (partially or totally, depending on the specification) when marital contracts are allowed to include provisions referring to divorce. Given that since that moment marital contracts can be used to limit or eliminate the compensation, this was interpreted as lowering wives' bargaining power. Consistent with this interpretation, married women labor supply increased by around 1.2 hours per week, while the probability of employment did it by 2.6 percent.

Overall, these results are compatible with the predictions of the non-unitary approach to household modeling. The relative position of spouses within the household matters to determine household outcomes. Family law has an important role in contributing to determine the bargaining position within a family and then in shaping economic outcomes.

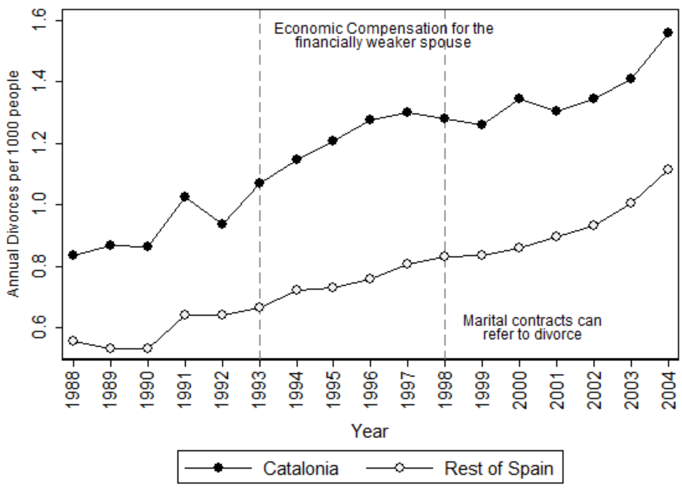
Figures

Figure 2.1: Marital Contracts



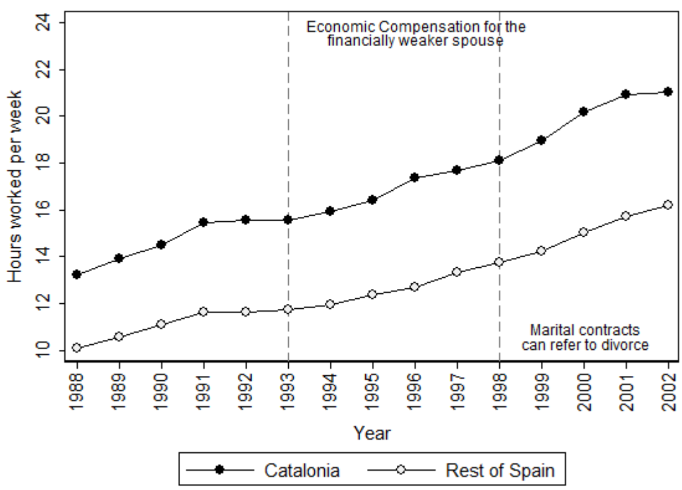
Source: Administrative data taken from Registries and Notaries Yearbook (*Anuario de la Direccion Nacional de los Registros y del Notariado*).

Figure 2.2: Divorce Rate



Source: Administrative data taken Judicial Statistics (*Consejo General del Poder Judicial*).

Figure 2.3: Usual weekly hours. Married Women 30-50 years old



Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics, Spain.

Tables

Table 2.1: Impacts on Marital Dissolution

Dependent variable	Divorce rate		Separation rate	
	(1)	(2)	(3)	(4)
cat*post93	0.123***		0.168***	
	(0.018)		(0.030)	
cat*post98	-0.050***		-0.043	
	(0.017)		(0.030)	
Years 1-2		0.099***		0.120***
		(0.023)		(0.040)
Years 3-4		0.155***		0.209***
		(0.023)		(0.041)
Years 5-6		0.093***		0.171***
		(0.023)		(0.040)
Years 7-8		0.095***		0.139***
		(0.023)		(0.040)
Years 9-11		0.056***		0.106***
		(0.020)		(0.034)
Year effects	$F = 125.3$	$F = 123.1$	$F = 220.1$	$F = 215.4$
Region effects	$F = 293.4$	$F = 293.3$	$F = 147.8$	$F = 148.9$
Adj. R^2	0.969	0.969	0.960	0.959
N° of obs	286	286	286	286

Notes: Divorce (Separation) rate is the annual number of divorces (separations) per 1000 people. Estimated using region's population weights. Sample period 1990-2004. The control group includes regions 1-4, 6, 10-12, 14, 16-17. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 2.2: Impacts on Married Women Labor Supply

	Ordinary Least Squares		Tobit	Panel-FE	
Dependent variable	Hours (1)	Employment (2)	Hours (3)	Hours (4)	Employment (5)
post93*cat	-0.688*	-0.018*	-2.508***	-0.463**	-0.011**
	(0.402)	(0.010)	(0.971)	(0.210)	(0.005)
post98*cat	1.288***	0.026***	1.127	0.237	0.004
	(0.374)	(0.009)	(0.812)	(0.242)	(0.006)
Region FE	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes
Individual FE	no	no	no	yes	yes
Adj. R^2	0.108	0.124		0.001	0.002
N^o of Obs.	493277	493277	493277	493277	493277

Notes: The sample includes women aged 30-50 years, who appear appear in 6 interviews with the same marital status. Sample period 1990-2002. The control group includes regions 1-4, 7-8, 11, 13, 16-18. The vector of control variables contains age, age squared, educational dummies, regional employment rate, per capita GDP at the regional level, and spouse-level controls such us age and education. Cluster-robust (at individual level) standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 2.3: Impacts on particular subgroups: Wives of firm's owners

	Ordinary Least Squares		Tobit	Panel-FE	
Dependent variable	Hours (1)	Employment (2)	Hours (3)	Hours (4)	Employment (5)
post93*cat	-1.477***	-0.034***	-3.612***	-0.748*	-0.010
	(0.423)	(0.010)	(0.918)	(0.438)	(0.011)
post98*cat	1.586***	0.031***	2.416***	-0.170	-0.004
	(0.377)	(0.009)	(0.777)	(0.477)	(0.014)
Region FE	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes
Individual FE	no	no	no	yes	yes
Adj. R^2	0.068	0.075		0.001	0.002
N^o of Obs.	117036	117036	117036	117036	117036

Notes: The sample includes wives aged 30-55 years whose husband declares to be the owner/maganer of a firm with or without employees, who appear appear in 6 interviews with the same marital status. Sample period 1990-2002. The control group includes regions 1-5, 7-8, 11, 13, 16-18. The vector of control variables contains age, age squared, educational dummies, regional employment rate, and spouse-level controls such us age and education. Cluster-robust (at individual level) standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

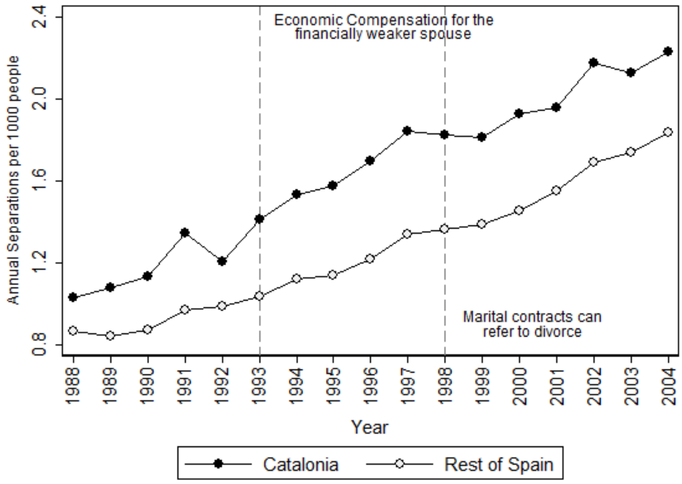
Table 2.4: Impacts on Married Women Labor Supply. Triple Difference

	Ordinary Least Squares		Tobit	Panel-FE	
Dependent variable	Hours (1)	Employment (2)	Hours (3)	Hours (4)	Employment (5)
post93*cat*marriedw	-0.692	-0.029	-2.971	-4.365*	-0.051
	(3.102)	(.074)	(5.402)	(2.497)	(0.043)
post98*cat*marriedw	0.088	-0.025	-1.652	1.205	0.043
	(2.179)	(.050)	(3.637)	(1.489)	(0.028)
Region FE	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes
Individual FE	no	no	no	yes	yes
Adj. R^2	0.117	0.133		0.002	0.002
N	471570	471570	471570	471570	471570

Notes: The sample includes women aged 30-50 years, who appear appear in 6 interviews with the same marital status. Sample period 1990-2002. The control group includes regions 1-4, 7-8, 11, 14, 16-18. The vector of control variables contains age, age squared, educational dummies, regional employment rate and GDP per capita. Cluster-robust (at individual level) standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

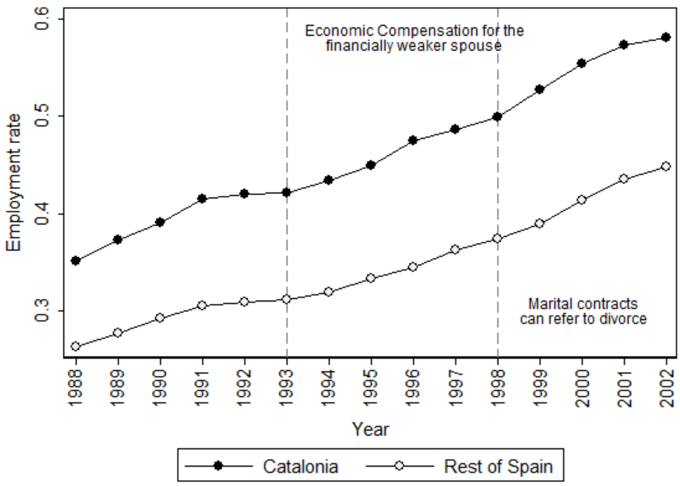
Appendix A2. Additional Figures and Tables

Figure A2.1: Separation Rate



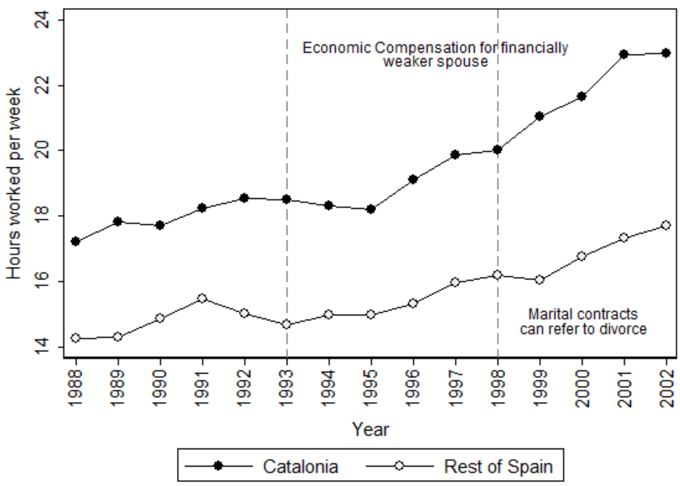
Source: Administrative data taken Judicial Statistics (*Consejo General del Poder Judicial*).

Figure A2.2: Employment rate. Married Women 30-50 years old



Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics, Spain.

Figure A2.3: Usual weekly hours. Wives of firm's owners



Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics, Spain.

Table A2.1: Summary Statistics. Married Women 30-55 years old

	Catalonia		Rest of Spain	
	Mean	**St. Dev.**	**Mean**	**St. Dev.**
1-1990 to 3-1993				
Weekly labor hours	38.985	12.929	37.550	15.889
Age	42.479	7.324	42.457	7.357
No education	0.362	0.481	0.395	0.489
Primary school	0.286	0.452	0.288	0.453
Secondary school	0.182	0.386	0.160	0.367
Tertiary/Univ school	0.170	0.376	0.157	0.364
Studying	0.007	0.085	0.012	0.108
Employment rate	0.376	0.010	0.314	0.034
Wife age	39.770	7.660	39.782	7.791
No education (wife)	0.375	0.484	0.420	0.493
Primary school (wife)	0.315	0.465	0.311	0.463
Secondary school (wife)	0.181	0.385	0.158	0.364
Tertiary/Univ school (wife)	0.128	0.335	0.112	0.315
In Labor Force	0.967	0.178	0.953	0.213
Employed	0.922	0.267	0.878	0.327
Obs	39,455		316,327	
4-1993 to 3-1998				
Weekly labor hours	37.904	14.381	36.529	16.881
Age	42.795	7.209	42.823	7.239
No education	0.054	0.226	0.076	0.265
Primary school	0.363	0.481	0.405	0.491
Secondary school	0.341	0.474	0.299	0.458
Tertiary/Univ school	0.241	0.428	0.220	0.414
Studying	0.018	0.131	0.017	0.130
Employment rate	0.372	0.015	0.313	0.038
Wife age	40.190	7.566	40.263	7.600
No education (wife)	0.069	0.254	0.089	0.285
Primary school (wife)	0.344	0.475	0.410	0.492
Secondary school (wife)	0.354	0.478	0.314	0.464
Tertiary/Univ school (wife)	0.233	0.423	0.187	0.390
In Labor Force	0.963	0.189	0.948	0.221
Employed	0.897	0.304	0.856	0.351
Obs	49,621		418,189	
4-1998 to 4-2002				
Weekly labor hours	39.864	12.484	38.462	15.230
Age	43.344	7.191	43.279	7.063
No education	0.034	0.182	0.058	0.233
Primary school	0.244	0.429	0.300	0.458
Secondary school	0.470	0.499	0.412	0.492
Tertiary/Univ school	0.252	0.434	0.231	0.422
Studying	0.023	0.151	0.020	0.139
Employment rate	0.421	0.012	0.360	0.040
Wife age	40.925	7.392	40.885	7.322
No education (wife)	0.040	0.196	0.064	0.244
Primary school (wife)	0.232	0.422	0.296	0.457
Secondary school (wife)	0.484	0.500	0.431	0.495
Tertiary/Univ school (wife)	0.244	0.430	0.209	0.407
In Labor Force	0.964	0.187	0.945	0.228
Employed	0.936	0.245	0.897	0.304
Obs	39,753		354,857	

Table A2.2: Summary Statistics. Married Men 30-55 years old

	Catalonia		**Rest of Spain**	
	Mean	**St. Dev.**	**Mean**	**St. Dev.**
1-1990 to 3-1993				
Weekly labor hours	15.237	19.367	11.467	18.501
Age	42.185	7.321	42.178	7.433
No education	0.410	0.492	0.458	0.498
Primary school	0.317	0.465	0.309	0.462
Secondary school	0.158	0.365	0.134	0.340
Tertiary/Univ school	0.114	0.318	0.100	0.300
Studying	0.009	0.095	0.014	0.118
Employment rate	0.259	0.007	0.196	0.035
Husband age	45.258	8.424	45.218	8.477
No education (husband)	0.389	0.488	0.425	0.494
Primary school (husband)	0.286	0.452	0.285	0.452
Secondary school (husband)	0.163	0.369	0.141	0.348
Tertiary/Univ school (husband)	0.162	0.368	0.149	0.356
In Labor Force	0.483	0.500	0.381	0.486
Employed	0.411	0.492	0.303	0.460
Obs	41,184		328,949	
4-1993 to 3-1998				
Weekly labor hours	16.932	19.604	12.715	18.801
Age	42.293	7.287	42.378	7.339
No education	0.090	0.286	0.110	0.313
Primary school	0.378	0.485	0.442	0.497
Secondary school	0.322	0.467	0.278	0.448
Tertiary/Univ school	0.210	0.407	0.170	0.376
Studying	0.019	0.136	0.022	0.148
Employment rate	0.275	0.014	0.208	0.038
Husband age	45.305	8.415	45.264	8.390
No education (husband)	0.073	0.260	0.096	0.294
Primary school (husband)	0.391	0.488	0.428	0.495
Secondary school (husband)	0.307	0.461	0.266	0.442
Tertiary/Univ school (husband)	0.229	0.420	0.210	0.407
In Labor Force	0.568	0.495	0.452	0.498
Employed	0.464	0.499	0.344	0.475
Obs	52,210		439,215	
4-1998 to 4-2002				
Weekly labor hours	20.066	19.617	15.104	19.259
Age	42.740	7.238	42.640	7.207
No education	0.055	0.228	0.080	0.272
Primary school	0.264	0.441	0.325	0.468
Secondary school	0.456	0.498	0.400	0.490
Tertiary/Univ school	0.225	0.418	0.195	0.396
Studying	0.027	0.164	0.026	0.161
Employment rate	0.330	0.014	0.254	0.045
Husband age	45.586	8.320	45.393	8.142
No education (husband)	0.048	0.213	0.072	0.259
Primary school (husband)	0.273	0.446	0.322	0.467
Secondary school (husband)	0.439	0.496	0.381	0.486
Tertiary/Univ school (husband)	0.240	0.427	0.225	0.417
In Labor Force	0.620	0.485	0.503	0.500
Employed	0.553	0.497	0.416	0.493
Obs	42,477		377,769	

Table A2.3: Impacts on Married Men Labor Supply

	Ordinary Least Squares		Tobit	Panel-FE	
Dependent variable	Hours (1)	Employment (2)	Hours (3)	Hours (4)	Employment (5)
post93*cat	0.250	-0.000	0.264	-0.800**	-0.013*
	(0.298)	(0.006)	(0.326)	(0.319)	(0.007)
post98*cat	-0.174	-0.002	-0.199	-0.396*	-0.007
	(0.274)	(0.005)	(0.296)	(0.241)	(0.005)
Region FE	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes
Individual FE	no	no	no	yes	yes
Adj. R^2	0.022	0.035		0.002	0.002
N	255170	255170	255170	255170	255170

Notes: The sample includes married men aged 30-50 years, who appear appear in 6 interviews with the same marital status. Sample period 1990-2002. The control group includes regions 2, 4, 5, 7, 10-12, 14-15. The vector of control variables contains age, age squared, educational dummies, regional employment rate, and wife-level controls such us age and education. Cluster-robust (at individual level) standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

3 Sex Ratios in College and Family Formation

3.1 Introduction

The question of who marries whom and when has received considerable attention among economists as well as among scholars from other social sciences. Understanding the determinants of the matching process is important given its influence on many relevant demographic and economic dimensions. Marriage market outcomes have been shown to substantially affect the level of income inequality in a society, marital and non-marital fertility rates, labor market outcomes, and children related outcomes (see, for instance, Fernández and Rogerson, 2001; Fernández, Guner, and Knowles, 2005; Chiappori, Fortin, and Lacroix, 2002).

One key factor for understanding marriage market outcomes is the sex ratio (i.e. the relative number of men to women) in the relevant population, as already discussed by Becker (1981) in his Treatise on the Family. Since then, a number of studies have investigated empirically the effects of imbalanced sex ratios on partnership formation and dissolution patterns.[1] One dimension in which these studies differ is in the definition of what should be considered the relevant marriage market for an individual. Factors such as geographical area, race, ethnicity, religion, social class, education, age, and workplace, among others, have been proposed to define the limits within which the sex composition of potential partners matters.

In this paper, we argue that a relevant marriage market is school, in particular in countries where university education is strongly segregated by field, so that students are exposed to a fairly stable school class during their entire college education. Average age at first marriage suggests that

[1] Wilson (1987); South and Lloyd (1992); Angrist (2002); Chiappori, Fortin, and Lacroix (2002); McKinnish (2007); Svarer (2007); Nielsen and Svarer (2009); Charles and Luoh (2010); Abramitzky, Delavande, and Vasconcelos (2011).

the initial search for a spouse often takes place before entering the labor market, especially amongst more educated individuals. Indeed, educational institutions being important marriage markets is quite well established in the sociology literature (see, for instance, Mare, 1991; Blossfeld and Timm, 2003; Schwartz and Mare, 2005). Moreover, the available evidence about where people meet potential partners seems to suggest that school is possibly the most important marriage market. Using data from the National Health and Social Life Survey for 1992, Laumann, Gagnon, Michael, and Michaels (1994) found that school is the top category where married individuals report having met their spouse, with 23 percent of spouses in their sample having met their partners during college.

The purpose of this paper is to address whether the probability of finding a suitable match while at school depends on the sex-mix of school mates that young adults are exposed to during their university education. We test this hypothesis with Spanish data, Spain being one of the countries with strong field segregation in college and where high quality data are available on sex ratios by year, field and university. To conduct the empirical analysis, we combine administrative data on the sex composition at time of graduation by university and field of study with survey data on partnership status some years after graduation. To address the potential endogeneity of the choice of field of study, we include university and field-specific fixed-effects in the estimations. This implies that the effect we are interested in is identified from variation in the fraction of women within university and field and over time.

The main findings of the paper are that the probability of starting a family shortly after school increases with the balance in the sex-composition of one's school class, and that this probability is increasing in the number of members of the opposite sex, although this effect is only statistically significant for females.

The literature relating sex ratios to partnership formation patterns is quite large. Among the earliest studies, Easterlin (1961) suggested that the decline in marriage could be linked to the decline in sex ratios among the foreign born in the 1920s. Freiden (1974) presented a cross-sectional analysis of sex ratios in states and counties with the 1960 Census. Grossbard-Shechtman (1984, 1993) studied links between sex ratios and marriage rates in cities, as well as female labor supply. South and Lloyd (1992) found that greater marriage opportunities for women increase marriage and divorce rates, and decrease illegitimacy ratios. Wilson (1987) argued that marriage rates for black women are low relative to white women because of the limited supply of employed black men available as potential spouses. Jemmott, Ashby, and Lindenfeld (1989) and Uecker and Regnerus (2010) studied sex ratios and romantic commitment on college campuses. Brien (1991) found evidence that local sex ratios (at the county level) influence the decisions to enter marriage and to have children out of wedlock. Chiappori, Fortin, and Lacroix (2002) estimated the relationship between sex ratios and labor supply across states for couples in the PSID.[2]

An important concern with these early studies is omitted variables bias and reverse causality in the relationship between sex ratios and measures of economic and social conditions. More recent studies take advantage of exogenous shocks to sex ratios to overcome these concerns. Angrist (2002) uses "exogenous" variation in immigration flows to study the effects of sex ratios within immigrant groups on marriage outcomes of first and second-generation immigrants. Abramitzky, Delavande, and Vasconcelos (2011) exploit exogenous variation in sex ratios coming from WWI's casualties to study the impact of male scarcity on marital assortative matching and other marriage market outcomes. Charles and Luoh (2010) also study the effects of male scarcity on marriage market outcomes, although in this case the variation in sex ratios comes from male incarceration.

Our paper relates to a couple of studies that analyze the impact of changing sex ratios in specific marriage markets on partnership formation and disso-

[2]Other related papers from Sociology and Demography are Pitt-Rivers (1929); Keeley (1977); Secord (1983); Lichter, LeClere, and McLaughlin (1991); Stier and Shavit (1994); Lichter, Anderson, and Hayward (1995); South and Trent (2010)

lution patterns. Svarer (2007) examines whether the sex composition that individuals are exposed to in the workplace affects their partnership formation and dissolution probabilities. He finds that the fraction of coworkers of the opposite sex positively affects the risk of dissolution both for females and males, but it does not explain partnership formation for single individuals. In a related paper, McKinnish (2007) investigates the extent to which the sex-mix an individual encounters on the job affects his or her marital status, and finds that individuals who work with a larger fraction of coworkers of the opposite sex are more likely to be divorced.

A general problem in the literature is that in defining a "marriage market" people usually combine geography with personal characteristics. Svarer (2007) and McKinnish (2007) do differ by focusing on the workplace, and on the same line we focus on a specific marriage market: college, which is more important for initial partnership formation.

The rest of the paper is structured as follows. In Section 3.2 we provide evidence of the importance of college as a marriage market and describe the expected effects of changing sex ratios on family formation patterns. In Section 3.3 we outline the empirical methodology. In Section 3.4 we describe the main sources of data used in the empirical analysis. Section 3.5 presents the main empirical results, and finally, Section 3.6 concludes.

3.2 College as a marriage market and expected effects of changing sex ratios

In the past 50 years, there has been a remarkable increase in the fraction of women who have access to college education, a phenomenon common to many different countries (Goldin, Katz, and Kuziemko, 2006; Becker, Hubbard, and Murphy, 2010). In the United States, the gender gap in college attendance and graduation started to narrow for cohorts born in the 1950s, and had already reversed for cohorts born in the 1960s and afterwards (Goldin, Katz, and Kuziemko, 2006).

A similar trend can be found for Spain. Figure 3.1 shows, separately by sex, the fraction of each birth cohort who completed college education. College graduation rates are similar for males and females in the cohort born in the late 1950s, and since then have been higher for females. The sex ratio among college graduates went from 1,620 men per 1,000 women in the cohort born in the 1940s, to 723 men per 1,000 women in the cohort born in the 1970s.

This dramatic change in the sex-composition of the school class may have had profound implications for family formation patterns, if individuals tend to form partnerships while in school. A number of papers in the sociology literature suggest that educational institutions are well established marriage markets (see, for instance, Mare, 1991; Blau, 1994; Blossfeld and Timm, 2003; Schwartz and Mare, 2005). University institutions provide contact opportunities with other individuals presumably with similar preferences, and this is the first step for friendship to develop (Blau, 1994). Moreover, opportunities for establishing social relationships are not limited to contacts made directly within the classroom or the educational institution itself, but also to those within the extended social network (i.e. friends of friends) that this implies (Blossfeld and Timm, 2003).[3]

There is also evidence in the economic literature pointing in this direction. For instance, Nielsen and Svarer (2009) use Danish register-based data to investigate how much of the systematic relationship between the educational level of the partners is explained by opportunities and how much by preferences, and find that half of the systematic sorting is explained by lower search frictions within educational institutions. In their data, 20 per-

[3]Blossfeld and Timm (2003) propose another reason why the educational system has an important role as a marriage market. Since those who pursue an university degree usually postpone the starting of a family longer, "the probability will grow that they will then quickly "catch up" with their age cohort after leaving school and eventually marry the partner who became a boy or girl friend during the period of education".

cent of couples are such that both partners attended the same educational institution.

Available data on new marriages in Spain also suggest a high degree of educational assortative mating. For instance, 71 percent of men with a college degree who got married between 2008 and 2010 did so to a woman with a college degree. Similarly, 51 percent of women with a college degree married a man with a college degree (See Table 3.1). Out of the total number of few marriages celebrated between 2008 and 2010, in 19 percent of cases both spouses had a college degree.[4] And finally, 36 percent of all women who got married between 2008 and 2010, and 26 percent of men, were college-educated.

a School class sex-composition and partnership formation

The potential impact of changing sex ratios on marriage market outcomes has been extensively studied in the economic literature (e.g. Becker, 1981; Chiappori, Fortin, and Lacroix, 2002). We focus the analysis on a specific outcome: stable partnership formation.

Search and matching models applied to partnership formation decisions are an useful tool for understanding the link between sex ratios and marriage opportunities (Oppenheimer, 1988; Mortensen, 1988; Weiss, 1993). Let's assume that all individuals are single and search for a partner while in school. In each period, individuals meet potential partners with a certain probability. The decision whether to form a partnership or not will depend on the expected gain from the current union compared to the expected gain from continued search. In standard search models, if the gain from the current partnership is higher than a reservation level, individuals decide to form the marriage or stable relationship. Otherwise, they continue

[4] In this calculation, the denominator is the pool of new marriages in which both spouses reported their education level. This is the case in 75 percent of all new marriages celebrated between 2008 and 2010.

searching.

For a given reservation level, an increase in the number of potential partners will increase the number of encounters an individual faces per unit of time and, consequently, the probability of forming a stable partnership. There are two straightforward predictions from this simple framework: (i) the larger the number of classmates of the opposite sex, the higher the probability of finding a stable partner per unit of time, and (ii) the number of matches will be larger the more balanced is the sex-composition of the school class. Clearly, however, the number of opposite-sex classmates will affect the reservation level as well, since individuals can become more or less selective depending on whether they know they will have more or fewer chances to meet new potential partners in each period. Although theoretically the aforementioned effects of the sex ratio on matching probabilities may be offset by its effect through changes in the reservation level, in the job-search literature the first effect has been shown to dominate under a broad set of conditions (van den Berg, 1994; Svarer, 2007).

3.3 Empirical Methodology

We estimate empirically the impact of the sex-composition an individual faces while in college on family formation patterns. We test our two main predictions. First, we expect to find a negative association between the degree of balance in the sex-composition of one's school class and the probability that an individual from that class (male or female) starts a family shortly after finishing college. Second, we expect that a higher proportion of opposite sex members in the individual's school class will increase his or her chances of finding a partner (and this effect would thus have opposite signs for men and women).

To assess the first prediction, we estimate the following linear probability model using a sample of recent college graduates:

$$Y_{ifugt} = \alpha + \beta SI_{fug} + \gamma X_{ifugt} + \mu_f + \rho_u + \delta_g + \lambda_t + \epsilon_{ifugt} \tag{3.1}$$

where Y_{ifugt} is an indicator of partnership status for individual i, who studied field f at university u, graduated in year g, and was observed in year t. SI_{fug} is a measure of "sex imbalance", that is, the extent to which the actual sex-mix in the school class differs from the value that would correspond to a totally sex-balanced class (i.e. one in which the fraction of females is 50 percent of the total number of students). We define this variable as the absolute value of the difference between the actual fraction of women in the individual's class and 0.5. Therefore, β is the coefficient of interest and is expected to be negative: the more unbalanced is the sex-composition of the school class, the less likely the formation of partnerships. X_{ifugt} is a vector of control variables, including a binary indicator for females, a third-degree polynomial in age, a binary indicator for whether the individual is still at school, a binary indicator for short (3-year) degrees, and the region- and field-specific unemployment rate.

The main concern with this empirical strategy is the potential endogeneity of the choice of field of study. This would bias our results if preferences for family formation are correlated with the choice of field. To address this concern, we include university and field fixed-effects to sweep out any unobserved field and university characteristics. In addition, we include a set of fixed-effects for region characteristics (at the province level), year of graduation, and year in which the individual is surveyed. Thus, our model is identified by the within-field and university (plausibly random) variation over time in the fraction of women.[5]. In further specifications, we test the robustness of the results obtained with the baseline approach by adding all pairwise interactions of the 4 sets of fixed-effects.

To assess our second prediction, we estimate the following equation sepa-

[5]Two papers that follows a similar identification strategy are Hoxby (2000) and Gould, Lavy, and Paserman (2009)

rately for males and females:

$$Y_{ifugt} = \alpha + \beta FW_{ugf} + \gamma Z_{ifugt} + \mu_f + \rho_u + \delta_g + \lambda_t + \epsilon_{ifugt} \tag{3.2}$$

where, again, Y_{ifugt} is an indicator of partnership status for individual i, who studied field f at university u, graduated in year g, and was observed in year t. Now, FW_{fug} is the fraction of women that this individual encountered during her period of education. In this case, the coefficient of interest, β, is expected to be negative for females and positive for males. Finally, the vector Z_{ifugt} is similar to X_{ifugt} in equation 3.1, but without including the dummy for females.

We consider four indicators of partnership status: (i) being currently in residence with an opposite-sex partner, (ii) being legally married, (iii) having ever been married, and (iv) currently residing with a partner who studied the same education field. The first three indicators capture the effect of imbalanced sex ratios on family formation decisions through a potentially broader network, while the fourth one aims at capturing more specifically the effect explained by contact opportunities within the school class.

3.4 Data and Descriptive Statistics

a Data

We combine two sources of data to conduct our empirical analysis: administrative University Education Statistics, and the Spanish Labor Force Survey, both provided by the National Institute of Statistics of Spain.

Sex Ratios

The University Education Statistics offer administrative data on the most relevant characteristics of the student body on an annual basis, which can be broken down by sex, university, and field of study, among other dimensions. This information is publicly available from the Spanish National Institute of Statistics in electronic format since 1998.

For every year between 1998 and 2008, we collect the number of men and women who graduated from university, by college and exact field of study. These data cover 69 universities, 113 fields of study (77 corresponding to long degrees -usually lasting 5 years- and 36 to short degrees -3 years), and 11 graduating years. We then aggregate these fields in order to match the classification of fields provided in the Labor Force Survey. Finally, we compute the fraction of students who are female in each field, university center and year, as the relevant measure of sex-composition.

Figure 3.2 depicts the evolution of the fraction of women among college graduates by year of graduation, over the entire period for which this information is available.[6] On average, in each year between 1998 and 2008 604 out of 1000 individuals graduating from college were females. This proportion is slightly increasing since 2002, and reaches a maximum for this period in 2007 with 613 women out of 1000 college graduates.

Figure 3.3 plots the fraction of women in each graduating cohort for selected fields, and shows that there is considerable variation across fields in the sex-composition of graduates. For example, Computer Science and Engineering, two relevant fields in terms of total number of students, are among those with a low fraction of women graduating from college, with averages of 238 and 253 out of 1000 graduates, respectively, during the entire period. On the other hand, Teaching and Medicine are among those with a high

[6] In the empirical analysis below we only use the data for 1998-2004, since the data from which we obtain the family status indicators only contain information on field of study until 2004.

fraction of women among college graduates, with 783 and 777 out of 1000, respectively.

There is also substantial variation within fields and over time in the fraction of women among college graduates. Figure 3.4 shows this for the same four selected fields: Computer Science, Engineering, Teaching, and Medicine.[7]

Finally, Figure 3.5 shows how the fraction of females among those graduating from college varies by region. The region with the lowest fraction Cantabria (553 out of 1000), while the Balearic Islands is the one with the highest (654 out of 1000).

Family Formation

Individual-level information on family status and educational background comes from the Spanish Labor Force Survey, conducted by the National Institute of Statistics on a quarterly basis. From this source of data, we have information on family status, education level and field of study, year of graduation, province of birth and residence, for a large sample of college graduates.

There are some data limitations that are worth mentioning. First, information on field of study is only available in the surveys conducted between 2000 and 2004, therefore we have to restrict our analysis to individuals who were interviewed in those years. This implies that all college educated individuals in our sample finished their studies in 2004 or before. Since information about sex ratios by university center and field of study is only available since 1998 onwards, we also have to restrict the analysis to those who finished college in 1998 or afterwards. Second, the labor force survey does not provide the exact university in which the individual graduated, so we proxy it with province of residence and field.[8] Also, we observe field

[7]Figure A3.1 in the Appendix plots the evolution of the fraction of women among college graduates for a broader selection of fields.

[8]Our data cover 69 universities, and some of them have schools in more than one

of study slightly aggregated. The field of study the respondent reports in the survey is coded into 21 long degrees and 14 short degrees. These two approximations lead to some measurement error in our main explanatory variable.

b Sample Definition and Descriptive Statistics

The main sample for the analysis of the impact of the sex ratio in college on family formation patterns consists of native college graduates between 22 and 40 years of age, who graduated from college between 1998 and 2004, and were surveyed between 2000 and 2004. After excluding those who do not report either the year of graduation or the sector of study, we are left with a sample of 19,609 individuals (11,413 females and 8,196 males).

In the final sample, 9.4 percent of individuals report to be living with a partner, 7.5 percent are legally married, and 7.8 percent have ever been married (See Table 3.2). In all cases, these figures are slightly higher for females than for males. For instance, while almost 11 percent of women in our sample are living with a partner, only 7 percent of men are. Another outcome we consider is whether individuals are in a relationship with someone from the same field of study. In our sample, 1.7 percent of college graduates are living with a partner who studied the same field, a figure that accounts for 18 percent of the couples we observe in the sample.

The typical college graduate included in our sample attended college in a field where 59 percent of students were females, as measured by the sex ratio at the year of graduation. Also, the average number of years since graduation is 1.82, and 49 percent of college graduates in our sample are currently studying (Table 3.2).

Finally, Tables A3.1 and A3.2 present the main descriptive statistics for our

province. Of the 52 Spanish provinces, there are 34 with only university, 13 with 2 universities, 2 with 3 universities, 1 (Valencia) with 4 universities, 1 (Barcelona) with 9 universities, and 1 (Madrid) with 13 universities.

two main explanatory variables. Table A3.1 shows statistics for the fraction of women by field of study, separately for short and long degrees. Table A3.2 reports statistics for the degree of balance in the sex-composition of the school class by field of study, and also separately for short and long degrees. This variable takes values between 0 (perfectly balanced sex-mix) and 0.5 (perfectly unbalanced sex-mix).

3.5 Results

The main results for the estimation of equation 3.1 are shown in Table 3.3. Columns 1 to 4 present the estimated effects for each of the four measures of partnership status. In all cases, the estimated effect is negative, as predicted by the theory. Individuals exposed to more sex-imbalanced classes are less likely to initiate a family shortly after finishing their university studies. However, in three out of the four outcomes, the effects are not estimated with precision. The only coefficient that is statistically different from zero, at the standard significance levels, is the one in Column 4, when the outcome variable is an indicator of partnership with an individual from the same field of study.

With respect to the magnitude of the effects, a one-standard-deviation increase in the degree of balance of the sex-composition of the school class increases the predicted probability of having a partner from the same field of study by 0.38 percentage points, which is equivalent to a 24 percent increase in the sample average of that probability.[9]

Table 3.4 reports the results of estimating equation 3.2. Now the main explanatory variable is the fraction of women individuals encounter during their period of education. As in the previous table, the four columns present

[9] Alternatively, for an individual moving from the 75th percentile to the 25th percentile of the degree of sex-imbalance of the school class, the probability of having a partner from the same field of study increases by 0.54 percentage points.

the results for the four different indicators of partnership status considered in the analysis. The top panel shows the estimated effects for females, while the bottom panel reports the results for males. The evidence shows that for both women and men the predicted probabilities of being in a partnership are increasing in the fraction of opposite sex classmates they are exposed to during their studies, as predicted by the theory.

In the case of women, the estimated coefficients are always negative and statistically significant, with the exception of Column 1 (when the outcome is the probability of residing with a partner) when the estimate is very close to 0. When the outcome is an indicator for being married (Column 2), a one-standard-deviation decrease in the fraction of females in the college class increases the predicted probability of being married by 0.87 percentage points, or 10 percent of the sample mean. Similar results are obtained when the dependent variable is a binary indicator for having ever been married (Column 3). When analyzing the probability of forming a partnership with someone who studied the same field (Column 4), the results indicate that a one-standard-deviation reduction in the fraction of females in the school class increases the predicted probability of having a partner from the same field by 0.53 percentage points, which is equivalent to 33 percent of the sample mean.

In the case of men, the coefficients have the expected sign but are not statistically significant at standard significance levels. The magnitudes of the point estimates are similar or slightly smaller than those found for women, depending on the outcome considered. For instance, a one-standard-deviation increase in the fraction of women in the college class increases the predicted probability of being married by 0.52 percentage points, or 10 percent of the sample mean. This effect is of similar size to the one found for women. Similarly, a one-standard-deviation increase in the fraction of women increases the predicted probability of having a partner from the same field by 0.16 percentage points, or 10 percent of the sample mean. In this case the effect is smaller than the one found for women (33 of the sample mean).

In Tables 3.5 to 3.7, we test the robustness of the baseline results by including in the regressions all possible pairwise interactions of the four sets of fixed-effects. For example, when we interact the field fixed-effects with the year of graduation fixed-effects (Column 1), we are controlling for unobserved differences across years of graduation within fields, and the model remains identified by the variation in the sex-composition of the school class within field and year of graduation across universities. The same reasoning applies to the remaining columns, where different fixed-effect interactions are included.

In Table 3.5, we test whether the effect of changes in the sex-imbalance of the college class on the predicted probability of having a partner from the same field (Column 4 of Table 3.3) is robust to the inclusion of these interaction terms. The results show that the effect of the degree of balance in the sex-composition of the school class on the predicted probability of being in a relationship with someone who studied in the same field remains statistically significant and of similar size across all five specifications.

In Table 3.6 we present the results of including all the interaction terms for fixed-effects in equation 3.2 and for the case when the dependent variable is an indicator for being legally married. The baseline result, reported in Column 2 of Table 3.4, was a coefficient of 0.055 (0.032) for women and of 0.025 (0.026) for men. In the case of women, after the introduction of all possible interaction of fixed-effects, the coefficient always maintains the negative sign, and remains statistically significant in 3 out 5 specifications.

Table 3.7 reports the results of adding the interaction terms in equation 3.2, when the outcome is an indicator for couples where both partners studied the same field. The baseline result in this case, shown in Column 4 of Table 3.4, was a coefficient of 0.033 (0.019) for women and of 0.0009 (0.017) for men. Again, in the case of women the coefficient maintains the negative sign and is statistically significant in 3 out of 5 specifications.

Finally, in Tables A3.3 and A3.4 we test the robustness of the results to different measures of the main explanatory variable. In the first case (Table A3.3), the main regressor is a dummy variable that takes value 1 if the individual was exposed during his or her studies to a high fraction of females, where high means a fraction of women higher than the sample average. In the second case (Table A3.4), we define a set of dummies for the quartiles of the fraction of women. In both cases the results remain essentially unchanged, that is, the higher the fraction of opposite-sex classmates, the lower the predicted probability of being in a stable relationship shortly after finishing school.

3.6 Conclusion

This paper explores the role of sex ratios in college in explaining family formation patterns. We investigate whether the sex-composition of the school class affects the family formation patterns of young adults after finishing their college education. We first construct the fraction of females among college graduates, by university and field of study, using administrative data for Spain. We then combine these data with a sample of recent college graduates containing information on field of study and partnership status, and test whether the sex-composition of the school class helps explain family formation patterns.

The evidence suggests that sex ratios in college matter. We find that the probability of starting a family shortly after school increases with the degree of balance in the sex-composition of one's school class. We also find this probability to depend on the fraction of opposite-sex school mates that individuals face during their period of study. Women who are exposed to a larger fraction of men during college are more likely to be married or residing with a partner from the same field of study, a few years after college. Similar results are found for men, although these estimates are not statistically significant. Moreover, we in general find sizable effects. For example, a one-standard-deviation increase in the fraction of males a woman faces in college increases her predicted probability of being married by 10 percent and the predicted probability of having a partner from the same field of study by 33 percent.

We contribute to the growing literature that examines the role of sex ratios in shaping marriage market outcomes. Our results suggest that school is a relevant marriage market in which individuals may find suitable partners, and the sex-composition in college is important in explaining the likelihood of this happening. These results are consistent with previous studies that used broader definitions of marriage markets. However, acknowledging that educational institutions are important places in which individuals meet potential partners, in a world characterized by a boom in higher education, opens new opportunities to study how sex ratios affect several marriage market outcomes.

Our results are subject to a number of limitations, some of them we hope to overcome in future research. First, our sample size is restricted to those individuals who graduated from college between 1998 and 2004.[10] We expect to revisit our results using a longer period in future versions of the paper. The additional time variation is expected to enrich the results and make them more precise.[11] Second, our results suffer from measurement error in our main explanatory variables for two reasons. One reason is that our survey data do not provide the exact university in which the individual graduated, so we proxy it using information on province of residence and field. The second reason is that we observe field of study in the survey

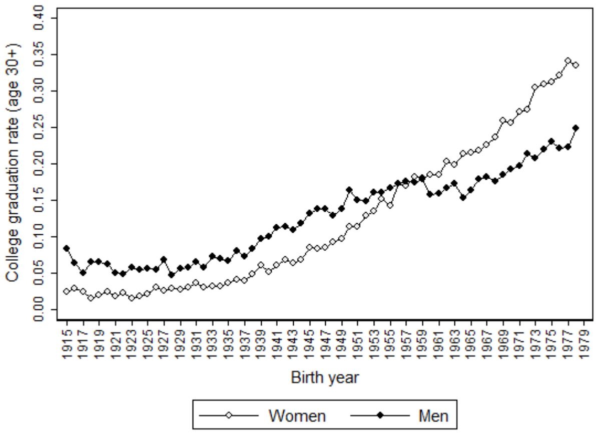
[10]On the one hand, the administrative data we use to construct the sex-composition by field is available since 1998. On the other hand, the public use microdata files of the labor force survey only contain information on field of study between 2000 and 2004. In both cases additional information exists and we will incorporate it into the analysis as soon as it becomes available to us.

[11]We also plan to enrich our measures of sex ratios using enrollment data (i.e. calculate sex ratios at freshman year instead of graduation year). In the current version of the paper we do not take advantage of enrollment data because this would imply to be left with an even smaller sample -i.e. those individuals who enrolled after 1998 and graduated before 2004.

data slightly more aggregated than in the administrative data. We plan to address these concerns in future research.

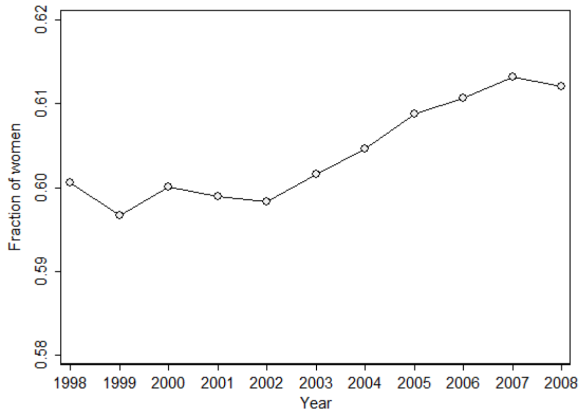
Figures

Figure 3.1: Fraction of college graduates by birth year and sex in Spain, 1915-1979



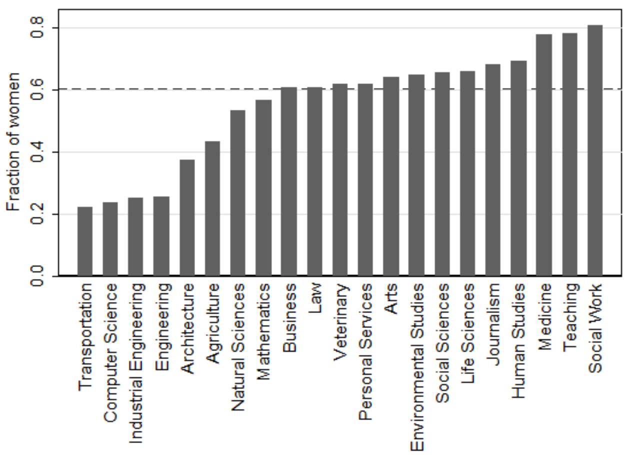
Note: The figure plots the fraction of females and males -aged 30 or more- by birth year who had finished college education in Spain. We used for the calculation the 2nd quarter samples of the Spanish Labor Force Survey for the years 2005 to 2009. Source: National Institute of Statistics of Spain.

Figure 3.2: Fraction of women among college graduates in Spain, 1998-2008



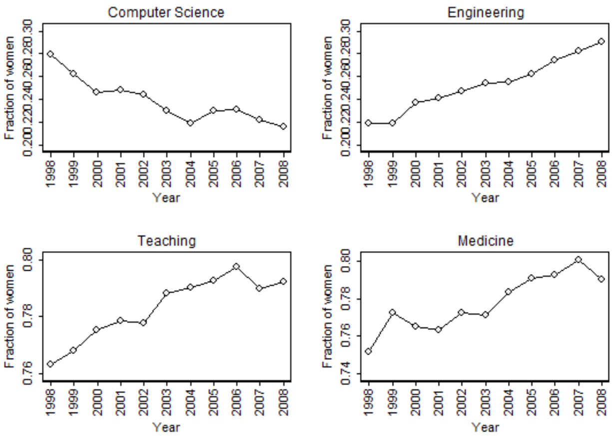
Note: The figure plots the fraction of females among those who had graduated from college in each year. Source: University Education Statistics, National Institute of Statistics of Spain.

Figure 3.3: Fraction of women among college graduates by field of study in Spain, 1998-2008



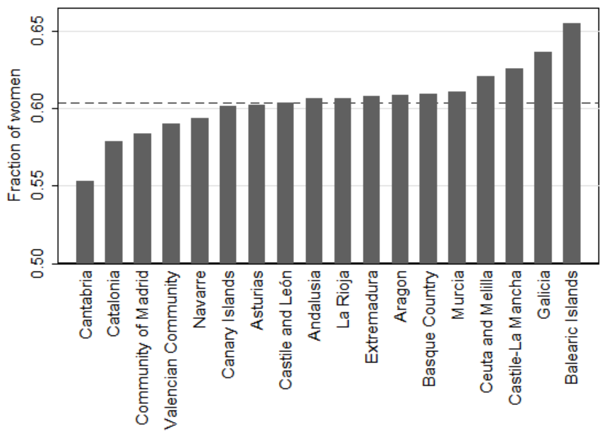
Note: The figure plots the fraction of females among those who had graduated from college in each field of study. The dashed line denotes the average fraction of females for all fields. Source: University Education Statistics, National Institute of Statistics of Spain.

Figure 3.4: Fraction of women among college graduates in Spain, 1998-2008. Selected fields



Note: The figure plots the fraction of females among those who had graduated from college in each year, for four selected fields of study (two with a relatively low fraction of women -Computer Science and Engineering- and two with a relatively high fraction of women -Teaching and Medicine). Source: University Education Statistics, National Institute of Statistics of Spain.

Figure 3.5: Fraction of women among college graduates by region in Spain, 1998-2008



Note: The figure plots the fraction of females among those who had graduated from college in each Autonomous Community in Spain. The dashed line denotes the average fraction of females at the country level. Source: University Education Statistics, National Institute of Statistics of Spain.

Tables

Table 3.1: New marriages by education of spouses, 2008-2010

Wife´s education	Total	Primary School or less	High School	Technical education	College or more	Missing
	Husband´s education					
Total	535,331	45,347	172,514	86,699	107,792	122,979
Primary School or less	36,094	18,583	12,222	3,332	1,168	789
High School	148,725	16,587	90,778	23,836	14,308	3,216
Technical education	81,221	5,671	30,442	29,956	13,088	2,064
College or more	148,725	3,398	35,716	27,992	76,394	5,225
Missing	120,566	1,108	3,356	1,583	2,834	111,685
	% Rows					
Total	100.0%	8.5%	32.2%	16.2%	20.1%	23.0%
Primary School or less	100.0%	51.5%	33.9%	9.2%	3.2%	2.2%
High School	100.0%	11.2%	61.0%	16.0%	9.6%	2.2%
Technical education	100.0%	7.0%	37.5%	36.9%	16.1%	2.5%
College or more	**100.0%**	**2.3%**	**24.0%**	**18.8%**	**51.4%**	**3.5%**
Missing	100.0%	0.9%	2.8%	1.3%	2.4%	92.6%
	% Columns					
Total	100.0%	100.0%	100.0%	100.0%	**100.0%**	100.0%
Primary School or less	6.7%	41.0%	7.1%	3.8%	**1.1%**	0.6%
High School	27.8%	36.6%	52.6%	27.5%	**13.3%**	2.6%
Technical education	15.2%	12.5%	17.6%	34.6%	**12.1%**	1.7%
College or more	27.8%	7.5%	20.7%	32.3%	**70.9%**	4.2%
Missing	22.5%	2.4%	1.9%	1.8%	**2.6%**	90.8%

Notes: Data come from the registry of all new marriages celebrated in Spain between 2008 and 2010 (spouses's education is only registered since 2008). Source: National Institute of Statistics of Spain.

Table 3.2: Descriptive statistics

	Total sample		Women		Men	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Main outcomes						
Living with a partner	0.0941	0.2919	0.1095	0.3123	0.0725	0.2594
Married	0.0749	0.2632	0.0883	0.2837	0.0562	0.2303
Ever-married	0.0778	0.2679	0.0924	0.2896	0.0575	0.2328
Same field couple	0.0171	0.1298	0.0168	0.1285	0.0176	0.1315
Main regressors						
Fraction of women	0.5926	0.1927	0.6451	0.1599	0.5194	0.2098
Sex imbalance	0.1850	0.1071	0.1874	0.1072	0.1816	0.1069
Control variables						
Age	25.95	2.96	25.71	2.87	26.29	3.06
Short degree	0.4341	0.4956	0.4355	0.4958	0.4321	0.4954
Currently studying	0.4909	0.4999	0.4928	0.5000	0.4883	0.4999
Years since graduation	1.82	1.46	1.85	1.46	1.78	1.45
Regional unemp. rate	0.1295	0.0471	0.1350	0.0446	0.1218	0.0494
Observations	19609		11413		8196	

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004 and were surveyed between 2000 and 2004. The main outcomes are four binary indicators for family status: being living with a partner, legally married, ever-married, or being living with a partner who studied the same field. Source: Microdata from the Spanish Labor Force Survey, National Institute of Statistics of Spain.

Table 3.3: Regression results for the degree of imbalance of sex composition within field

	Living with partner (1)	Married (2)	Ever-married (3)	Same field couple (4)
Imbalance	-0.016	-0.029	-0.036	-0.035***
	(0.027)	(0.025)	(0.025)	(0.013)
Female	0.049***	0.045***	0.049***	0.004*
	(0.004)	(0.004)	(0.004)	(0.002)
Age	-0.532***	-0.411***	-0.386***	-0.185***
	(0.086)	(0.085)	(0.083)	(0.052)
Age squared	0.017***	0.013***	0.012***	0.006***
	(0.003)	(0.003)	(0.003)	(0.002)
Age cubed	-0.000***	-0.000***	-0.000***	-0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Short degree	0.011**	0.015***	0.018***	0.004
	(0.006)	(0.005)	(0.005)	(0.003)
Currently studying	-0.037***	-0.029***	-0.030***	-0.005***
	(0.004)	(0.004)	(0.004)	(0.002)
Unemployment rate	0.073	0.013	0.019	0.084**
	(0.071)	(0.071)	(0.072)	(0.033)
Region FE	yes	yes	yes	yes
Year FE	yes	yes	yes	yes
Year of Grad FE	yes	yes	yes	yes
Field FE	yes	yes	yes	yes
Adj. R^2	0.185	0.171	0.184	0.035
Mean (depvar)	0.086	0.069	0.072	0.016
N	19035	19035	19035	19035

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Dependent variables are binary indicators for having a partner at home (column 1), being legally married (column 2), being ever-married (column 3), and being living with a partner from the same field of study (column 4). Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

Table 3.4: Regression results for the fraction of women in the same field of study

	Living with partner (1)	Married (2)	Ever-married (3)	Same field couple (4)
Panel A: Women				
Fraction of women	-0.005	-0.055*	-0.057*	-0.033*
	(0.034)	(0.032)	(0.032)	(0.019)
Adj. R^2	0.188	0.175	0.192	0.031
Mean (depvar)	0.101	0.083	0.087	0.016
N	11071	11071	11071	11071
Panel B: Men				
Fraction of women	0.022	0.025	0.018	0.009
	(0.031)	(0.026)	(0.026)	(0.017)
Adj. R^2	0.192	0.184	0.192	0.047
Mean (depvar)	0.065	0.051	0.052	0.016
N	7964	7964	7964	7964
Region FE	yes	yes	yes	yes
Year FE	yes	yes	yes	yes
Year of Grad FE	yes	yes	yes	yes
Field FE	yes	yes	yes	yes

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Dependent variables are binary indicators for having a partner at home (column 1), being legally married (column 2), being ever-married (column 3), and being living with a partner from the same field of study (column 4). Control variables include a third degree polynomial in age, a binary indicator for being studying, a binary indicator for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

Table 3.5: Regression results for the within-field sex composition. Different fixed-effects Specifications

	Dependent variable: Living with a partner from the same field				
	(1)	(2)	(3)	(4)	(5)
Imbalance	-0.039***	-0.037*	-0.038***	-0.038***	-0.036***
	(0.013)	(0.019)	(0.013)	(0.013)	(0.013)
Main fixed-effects					
Region	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
Year of graduation	yes	yes	yes	yes	yes
Field of study	yes	yes	yes	yes	yes
Interactions					
Field x year of graduation	yes	no	no	no	no
Field x region	no	yes	no	no	no
Year of graduation x region	no	no	yes	no	no
Field x year	no	no	no	yes	no
Region x year	no	no	no	no	yes
Adj. R^2	0.043	0.060	0.037	0.035	0.034
Mean (depvar)	0.016	0.016	0.016	0.016	0.016
N	19035	19035	19035	19035	19035

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Control variables include a third degree polynomial in age, binary indicators for females, for being studying, and for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

Table 3.6: Regression results for the fraction of women in the same field. Different fixed-effects specifications

	Dependent variable: Being married				
	(1)	(2)	(3)	(4)	(5)
Panel A: Women					
Fraction of women	-0.042	-0.037	-0.059*	-0.054*	-0.060*
	(0.033)	(0.044)	(0.033)	(0.032)	(0.033)
Adj. R^2	0.176	0.195	0.174	0.173	0.171
Mean (depvar)	0.083	0.083	0.083	0.083	0.083
N	11071	11071	11071	11071	11071
Panel B: Men					
Fraction of women	0.027	0.002	0.008	0.024	0.027
	(0.027)	(0.037)	(0.027)	(0.027)	(0.027)
Adj. R^2	0.188	0.207	0.185	0.183	0.178
Mean (depvar)	0.051	0.051	0.051	0.051	0.051
N	7964	7964	7964	7964	7964
Main fixed-effects					
Region	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
Year of graduation	yes	yes	yes	yes	yes
Field of study	yes	yes	yes	yes	yes
Interactions					
Field x year of graduation	yes	no	no	no	no
Field x region	no	yes	no	no	no
Year of graduation x region	no	no	yes	no	no
Field x year	no	no	no	yes	no
Region x year	no	no	no	no	yes

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Control variables include a third degree polynomial in age, a binary indicator for being studying, a binary indicator for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

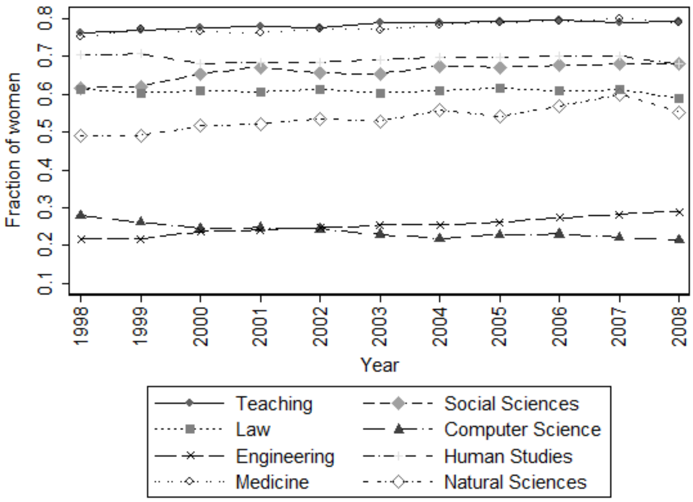
Table 3.7: Regression results for the fraction of women in the same field. Different fixed-effects specifications

	Dependent variable: Living with a partner from the same field				
	(1)	(2)	(3)	(4)	(5)
Panel A: Women					
Fraction of women	-0.024	-0.075***	-0.031	-0.033*	-0.032*
	(0.019)	(0.027)	(0.019)	(0.019)	(0.018)
Adj. R^2	0.041	0.068	0.027	0.029	0.025
Mean (depvar)	0.016	0.016	0.016	0.016	0.016
N	11071	11071	11071	11071	11071
Panel B: Men					
Fraction of women	0.011	-0.022	0.001	0.015	0.008
	(0.017)	(0.023)	(0.018)	(0.017)	(0.018)
Adj. R^2	0.064	0.074	0.053	0.047	0.042
Mean (depvar)	0.016	0.016	0.016	0.016	0.016
N	7964	7964	7964	7964	7964
Main fixed-effects					
Region	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
Year of graduation	yes	yes	yes	yes	yes
Field of study	yes	yes	yes	yes	yes
Interactions					
Field x year of graduation	yes	no	no	no	no
Field x region	no	yes	no	no	no
Year of graduation x region	no	no	yes	no	no
Field x year	no	no	no	yes	no
Region x year	no	no	no	no	yes

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Control variables include a third degree polynomial in age, a binary indicator for being studying, a binary indicator for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

Appendix A3. Additional Figures and Tables

Figure A3.1: Fraction of women among college graduates in Spain, 1998-2008



Note: The figure plots the fraction of females among those who had graduated from college in each year, for selected fields of study. Source: University Education Statistics, National Institute of Statistics of Spain.

Table A3.1: Fraction of college graduates who are females by field of study, 1998-2004

Field of study	Mean	St. Dev	Min	Max	Obs
Long degrees	**0.585**	**0.157**	**0.000**	**1.000**	**4681**
Teaching	0.835	0.050	0.571	0.947	313
Medicine	0.695	0.049	0.441	0.923	317
Human Studies	0.695	0.041	0.434	0.923	364
Journalism	0.652	0.058	0.333	0.875	160
Life Sciences	0.646	0.053	0.492	1.000	253
Social Sciences	0.646	0.072	0.167	0.890	335
Environmental Studies	0.636	0.071	0.426	0.850	114
Arts	0.629	0.068	0.381	0.760	162
Law	0.609	0.041	0.484	0.765	364
Veterinary	0.585	0.077	0.342	0.737	94
Mathematics	0.584	0.080	0.000	1.000	228
Business	0.569	0.055	0.278	0.775	351
Natural Sciences	0.520	0.071	0.143	1.000	363
Social Work	0.507	0.124	0.268	0.796	54
Agriculture	0.394	0.084	0.000	0.606	154
Architecture	0.383	0.087	0.000	0.592	182
Personal Service	0.291	0.076	0.000	0.684	141
Engineering	0.262	0.056	0.000	0.643	322
Industrial Engineering	0.244	0.072	0.000	0.600	68
Computer Science	0.240	0.078	0.000	1.000	272
Transport Studies	0.221	0.089	0.000	0.474	70
Short degrees	**0.619**	**0.223**	**0.000**	**1.000**	**3157**
Social Work	0.870	0.065	0.390	1.000	308
Medicine	0.819	0.056	0.699	1.000	356
Personal Service	0.810	0.058	0.556	1.000	199
Teaching	0.766	0.046	0.649	0.923	364
Journalism	0.746	0.064	0.603	1.000	114
Business	0.627	0.043	0.473	0.765	364
Mathematics	0.569	0.104	0.370	1.000	129
Agriculture	0.434	0.071	0.000	0.750	254
Architecture	0.342	0.059	0.000	0.667	213
Human Studies	0.332	0.256	0.000	0.833	27
Computer Science	0.249	0.066	0.000	0.489	328
Transport Studies	0.231	0.096	0.000	0.429	62
Engineering	0.224	0.049	0.000	0.537	326
Industrial Engineering	0.221	0.074	0.000	1.000	113

Notes: This table presents the main descriptive statistics for the fraction college graduates who are females by field of study, for the period 1998-2004. Fields of study observed in administrative registries are aggregated to match the classification of fields we observe in the Labor Force Survey. Source: University Education Statistics, National Institute of Statistics of Spain.

Table A3.2: Distance to balanced sex-composition by field of study, 1998-2004

Field of study	**Mean**	**St. Dev**	**Min**	**Max**	**Obs**
Long degrees	**0.156**	**0.087**	**0.000**	**0.500**	**4681**
Teaching	0.335	0.050	0.071	0.447	313
Transport Studies	0.279	0.089	0.026	0.500	70
Computer Science	0.260	0.076	0.026	0.500	272
Industrial Engineering	0.257	0.069	0.000	0.500	68
Engineering	0.238	0.054	0.000	0.500	322
Personal Service	0.211	0.072	0.048	0.500	141
Medicine	0.195	0.048	0.000	0.423	317
Human Studies	0.195	0.040	0.000	0.423	364
Journalism	0.153	0.057	0.000	0.375	160
Social Sciences	0.148	0.067	0.000	0.390	335
Life Sciences	0.146	0.053	0.004	0.500	253
Environmental Studies	0.141	0.061	0.009	0.350	114
Arts	0.131	0.064	0.008	0.260	162
Architecture	0.125	0.075	0.000	0.500	182
Agriculture	0.113	0.075	0.000	0.500	154
Social Work	0.112	0.050	0.017	0.296	54
Law	0.109	0.041	0.000	0.265	364
Veterinary	0.095	0.064	0.003	0.237	94
Mathematics	0.092	0.071	0.000	0.500	228
Business	0.071	0.051	0.000	0.275	351
Natural Sciences	0.056	0.048	0.000	0.500	363
Short degrees	**0.234**	**0.095**	**0.000**	**0.500**	**3157**
Social Work	0.372	0.049	0.056	0.500	308
Medicine	0.319	0.056	0.199	0.500	356
Personal Service	0.310	0.058	0.056	0.500	199
Industrial Engineering	0.280	0.068	0.045	0.500	113
Human Studies	0.278	0.121	0.000	0.500	27
Engineering	0.276	0.048	0.037	0.500	326
Transport Studies	0.269	0.096	0.071	0.500	62
Teaching	0.266	0.046	0.149	0.423	364
Computer Science	0.251	0.066	0.011	0.500	328
Journalism	0.246	0.064	0.103	0.500	114
Architecture	0.159	0.058	0.000	0.500	213
Business	0.127	0.043	0.003	0.265	364
Mathematics	0.098	0.076	0.003	0.500	129
Agriculture	0.082	0.052	0.000	0.500	254

Notes: This table presents the main descriptive statistics for the absolute value of the difference between the actual fraction of women and the one corresponding to a perfectly balanced sex-composition by field of study, for the period 1998-2004. Fields of study observed in administrative registries are aggregated to match the classification of fields we observe in the Labor Force Survey. Source: University Education Statistics, National Institute of Statistics of Spain.

Table A3.3: Regression results for the fraction of women in the same field of study

	Living with partner (1)	Married (2)	Ever-married (3)	Same field couple (4)
Panel A: Women				
High fraction of women	-0.007	-0.015**	-0.017**	-0.007*
	(0.008)	(0.007)	(0.007)	(0.004)
Adj. R^2	0.188	0.175	0.192	0.031
Mean (depvar)	0.101	0.083	0.087	0.016
N	11071	11071	11071	11071
Panel B: Men				
High fraction of women	0.008	0.004	0.003	-0.006
	(0.008)	(0.007)	(0.007)	(0.005)
Adj. R^2	0.192	0.184	0.192	0.048
Mean (depvar)	0.065	0.051	0.052	0.016
N	7964	7964	7964	7964
Region FE	yes	yes	yes	yes
Year FE	yes	yes	yes	yes
Year of Grad FE	yes	yes	yes	yes
Field FE	yes	yes	yes	yes

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Dependent variables are binary indicators for having a partner at home (column 1), being legally married (column 2), being ever-married (column 3), and being married to someone from the same field of study (column 4). “High fraction of women” is a dummy variable that takes the value 1 if the fraction of women in the field-university-year of graduation combination was above average. Control variables include a third degree polynomial in age, a binary indicator for being studying, a binary indicator for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

Table A3.4: Regression results for the fraction of women in the same field of study

	Living with partner (1)	Married (2)	Ever-married (3)	Same field couple (4)
Panel A: Women				
FW quartile 2	-0.013	0.000	0.000	-0.008
	(0.012)	(0.011)	(0.011)	(0.007)
FW quartile 3	-0.021*	-0.018	-0.017	-0.012*
	(0.013)	(0.012)	(0.012)	(0.007)
FW quartile 4	-0.011	-0.014	-0.014	-0.015*
	(0.015)	(0.014)	(0.014)	(0.008)
Adj. R^2	0.189	0.175	0.192	0.031
Mean (depvar)	0.101	0.083	0.087	0.016
N	11071	11071	11071	11071
Panel B: Men				
FW quartile 2	0.015	0.004	0.003	0.006
	(0.011)	(0.010)	(0.010)	(0.007)
FW quartile 3	0.016	0.009	0.008	0.001
	(0.011)	(0.010)	(0.010)	(0.007)
FW quartile 4	0.010	0.010	0.009	-0.001
	(0.015)	(0.014)	(0.014)	(0.009)
Adj. R^2	0.192	0.184	0.192	0.047
Mean (depvar)	0.065	0.051	0.052	0.016
N	7964	7964	7964	7964
Region FE	yes	yes	yes	yes
Year FE	yes	yes	yes	yes
Year of Grad FE	yes	yes	yes	yes
Field FE	yes	yes	yes	yes

Notes: The sample includes native individuals between 22 and 40 years of age, who graduated from college between 1998 and 2004. Dependent variables are binary indicators for having a partner at home (column 1), being legally married (column 2), being ever-married (column 3), and being married to someone from the same field of study (column 4). Control variables include a third degree polynomial in age, a binary indicator for being studying, a binary indicator for short degrees, and the region and field-specific unemployment rate. Robust standard errors are reported in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent levels, respectively. Source: Microdata from the Spanish Labor Force Survey, Spanish National Institute of Statistics.

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