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Legal Reforms, Conditional Cash Transfers, and Intimate Partner Violence: Evidence from Mexico

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Abstract

We study the relationship between divorce law reforms codifying intimate partner violence (IPV) as legal grounds for unilateral divorce, the Oportunidades conditional cash transfer program, and the incidence of IPV in Mexico. Using data from three nationally representative surveys in 2003, 2006, and 2011, we show the legal reforms lead to a 55 percent increase in annual divorce rates, concentrated among couples with a history of violence. Comparing groups of beneficiary and non-beneficiary households within villages, we find that IPV rates converge for these couples in the longer-run. Marital selection plays an important role in explaining the long-run relationships.

JEL Codes: J12, J16, K42

Keywords: divorce laws; conditional cash transfer programs; Oportunidades; divorce; intimate partner violence; marital selection

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1   Introduction

Intimate partner violence (IPV) is the most common form of violence experienced by women globally: approximately 30 percent of women experience this in some form during their lifetimes. IPV has been condemned internationally as a serious and costly human rights, public health, and women's personal security issue (Devries et al. 2013; Fearon and Hoeffler 2014). As a result, a growing number of national and subnational governments have instituted multiple efforts to promote women's ability to exit violent relationships as well as their empowerment within the household, partly with the intent of curbing intimate partner violence and other forms of violence against women.

Governments in many (mostly Western) countries have over the past 50 years introduced legal reforms that allow for unilateral divorce, which ease individuals' ability to exit marital relationships.[1] Strategies more commonly adopted across the developing world have focused on the implementation of legal reforms and social programs that aim to enhance women's options outside of existing marriages (i.e., property rights and inheritance reforms that favor women; cash and in-kind transfers-based poverty alleviation programs targeted to adult women), based on a growing consensus that targeting resources to women can promote their empowerment within the household.[2] A literature evaluating this broad set of programs and policy reforms in developing country settings tends to find moderate or substantial short-term reductions in the incidence of IPV as a result of both types of policy reforms (Buller et al., 2018). However, there is limited evidence on

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[1]These unilateral divorce law reforms, which allows a spouse to file for divorce unilaterally and without the other spouse having committed fault, have been instituted across states in the United States, Canada, and European countries since the late 1960s.

[2]For a survey of the broad literature on policies aimed at promoting female empowerment, see Duflo (2012). Buller et al. (2018) and Baranov et al. (2020) provide recent reviews of the work on cash transfers and intimate partner violence in low and middle-income countries. A growing body of research also studies how legal and judicial systems as well as access to justice programs can address gender inequities and improve women's rights and wellbeing, including exposure to intimate partner violence (see, for example, Anderson, 2018; Miller and Segal, 2019; Amaral, Bhalotra and Nishith, 2019; Kavanaugh, Sviatschi and Trako, 2019). Surveys of this small but growing literature can be found in Heise (2011); Doyle and Aizer (2018).

the important issue of how local legal institutions interact with social support programs to curb the incidence of IPV (Baranov et al., 2020).

The present paper investigates a series of questions regarding the long-run relationship between divorce legislation, conditional cash transfer programs and the incidence of IPV. Specifically, do legal reforms that help ease individuals' ability to exit violent relationships achieve *sustained* reductions in the incidence of IPV in developing country settings? If so, does this occur by inducing the marital dissolution of violent relationships? Do conditional cash transfer programs complement or substitute these legal efforts to curb intimate partner violence?

We study these questions in the context of rural Mexico. Starting in 1997, state governments across the country adopted a series of legal reforms formally including intimate partner violence as legal grounds for unilateral divorce. Concurrently, the federal government introduced in 1997 the landmark PROGRESA/Oportunidades/Prospera conditional cash transfer program – one of the earliest and most comprehensive poverty alleviation programs in Latin America. We exploit the introduction of these policies to provide evidence of the short and long-run relationship between these legal reforms, the Oportunidades CCT program, and the incidence of intimate partner violence. We use data from three nationally representative surveys, the National Surveys on the Dynamics of Household Relationships (ENDIREH) of 2003, 2006, and 2011. These surveys collect detailed information on IPV as well as households' program participation and women's marital histories, and thus provide rich data to document the consequences of the reforms for marital dissolution and patterns of IPV across program beneficiary and non-beneficiary households.

Our study uncovers several new findings regarding intimate partner violence in Mexico and its relationship with the instituted policies. The legal reforms, which reduce the cost of exiting violent relationships, lead to a substantial increase in divorce rates in the Mexican countryside. Exploiting the staggered adoption of these legal changes across states, we estimate a 0.35 percentage point (55 percent) increase in annual divorce rates following the passage of the reform. These effects are concentrated among Oportunidades program beneficiary households, those in which women reporting having experienced intimate partner violence, and those having experienced violence in

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their childhood, an important predictor of IPV incidence. These patterns suggest that divorce and marital selection can be an important driving force for reductions in the incidence of IPV.

Second, we study potential changes over time in the relationship between receipt of the Oportunidades CCT program and IPV, across the 2003, 2006, and 2011 surveys. We follow the observational analysis in Bobonis, González-Brenes and Castro (2013) comparing the incidence of women's experience of IPV across beneficiary and non-beneficiary households within villages (in 2003), and construct samples of comparable household for each of the two subsequent survey rounds (2006 and 2011). We first show that the short-run relationship between receipt of the Oportunidades CCT and reductions in IPV documented in Bobonis, González-Brenes and Castro (2013) is largely concentrated in states that had adopted the legal reforms to the divorce laws before 2003, early adoption states. The estimates suggest that beneficiary women in these states were 18 percentage points less likely to be victims of physical or sexual violence compared to non-beneficiary women in non-reform states in the short-run. This finding suggests a strong short-run complementarity between the two policies in reducing domestic violence. In contrast with these short-run outcomes, we find no differences in the incidence of IPV between women in beneficiary and non-beneficiary households in the longer-run. In combination with our results regarding marital dissolution patterns, the evidence is indicative of the idea that marital selection is an important mechanism of overall reductions in the incidence of IPV over time.

Finally, in the appendix we present a simple model of intra-household bargaining and conflict to help formalize the mechanisms via which both reductions in the cost of divorce and income support programs can affect equilibrium levels of marital dissolution and conflict within the household. Building on the work of Anderson and Genicot (2015) and Brassiolo (2016), our theoretical framework incorporates bargaining and conflict with incomplete information, together with endogenous divorce decisions, to formalize the effects that these two distinct sets of policies can have on the incidence of conflict and divorce. The first prediction we draw from the model shows how legal reforms, through a reduction in the costs of divorce, increase separations. More importantly, the model highlights the specific selection mechanisms through which the implemented policies, both

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CCTs and legal reforms, affect the *types* of couples which remain in union. In particular, the model reveals that relationships in which the private gains from marriage are lower and the cost of conflict experienced by the woman is higher, are more likely to dissolve once these policies are introduced. If these characteristics are associated with the occurrence of IPV, then these policies can increase the separation rate of relationships in which violence is prevalent.

Relevant Literature

Our paper contributes to the policy literature on intimate partner violence in a number of ways. While previous literature has highlighted how policy may affect the behaviour of violent partners, the potential of these to reduce violence by inducing divorce is an important and relatively unexplored channel through which this objective could be achieved. A growing body of work shows that unilateral divorce laws – instituted mainly in the context of developed countries – can help curb various forms of intimate partner violence (Dee 2003; Stevenson and Wolfers 2006; Brassiolo 2016). It is unclear whether the evidence provided by these studies is generalizable to developing countries, where differences in both socio-economic and health conditions as well as in cultural views towards IPV and social norms regarding gender roles (especially those that link notions of manhood to dominance and aggression) may limit the effectiveness of these legal rules in improving women's status and conditions within the household. García-Ramos (2018) explores the impact of unilateral divorce laws in Mexico and finds mixed evidence regarding the consequences of these for the incidence of physical, sexual, and emotional violence.[3] Our study complements this work: it examines a different set of (earlier) legal reforms – instituted in most (28 out of the 32) states in the country – that also enhance women's civil rights by creating a more equitable legal environment for marriage and divorce. Beleche (2017) studies the consequences of these reforms for other measures of violence within the household (i.e., female and male suicide rates) and finds no evidence of reductions in these extreme forms of violence and conflict resolution. In earlier

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[3]García-Ramos (2018) documents a positive relationship between the adoption of unilateral divorce laws and divorce rates in the Federal District (Mexico City) and the state of Hidalgo, which took place in 2008 and 2011 respectively.

work using aggregate administrative data on divorce rates, Beleche and Lew (2011) find no overall effect of this legal reform on divorce rates across Mexico.

In a recent review of the literature, Buller et al. (2018) summarize the existing evidence regarding the effects of conditional and unconditional cash transfer (CT) programs on intimate partner violence. While this body of work finds some subgroups to be at greater risk of some forms of violence, the overarching evidence is supportive of the view that CTs on average lead to short-run reductions in IPV. Few studies document the long-run consequences of the introduction of CT programs. Perova (2010), Ritter Burga (2014), and Díaz and Saldarriaga (Forthcoming) study the consequences of the introduction of the Juntos CT program in Peru; all find significant reductions in the incidence of physical and sexual violence over a period of up to five years following the introduction of the program. Roy et al. (2019a,b) study whether cash transfer programs lead to persistent reductions in IPV once these programs end. They show evidence based on a randomized controlled trial in rural Bangladesh that cash or food transfers to poor women alone do not lead to sustained reductions in IPV following the end of the program.[4]

Our findings suggest that the causal effects documented by the existing literature may not necessarily persist over time if changes in the legal environment affect household formation and dissolution patterns in such ways as to independently curb intimate partner violence. Finally, ours is the first study to explore the extent to which these legal reforms and social policies, which may both have independent effects on IPV, work as complements or substitutes. Our results are consistent with the view that policies to promote the improvement of women's livelihoods within the household can lead to sizeable reductions in IPV in the short-run, but that marital selection is an important mechanism explaining these effects in the longer-run.

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[4]They show strong evidence that a complementary nutrition behavior change communication intervention, together with cash transfers in particular, led to sustained reductions in IPV four years after the end of the program. Their evidence suggests that the combined intervention led to more sustained impacts through persistent increases in women's bargaining power, men's costs of perpetrating violence, and poverty-related emotional well-being.

2 Background

Family Violence Divorce Law Reforms

In Mexico, the laws regarding marriage, divorce, and family affairs are embedded in state civil codes; each of the 31 states has its own civil code regulating these proceedings. State civil codes have historically stipulated motives such as adultery, and out-of-wedlock births (among others) in addition to mutual consent, as grounds for divorce. In the late 1990s, states began adopting legal reforms with the objective of modernizing civil codes, one of which included adding family violence as a cause for divorce. The precise wording of the civil code varies across states, but it generally allows partners to unilaterally initiate divorce proceedings claiming family violence – the committal of violence by one of the partners against the other, their mutual children, or the children of one of them – as the cause. The specific dates at which these legal reforms were adopted across Mexican states are listed in Figure 1 (the source is the coding carried out by Beleche, 2017). The first state to adopt the reform was the Federal District (Mexico City) in December of 1997. By 2003, the majority of states had adopted the legal reform. In that year, domestic violence or threats as the cause for divorce accounted for about 1.6 percent of all divorces in urban areas, and about 2.2 percent in rural areas.[5]

Figure 2 depicts trends in the adoption of the family violence reforms and divorce rates since 1993, and suggests a relationship between the increase in divorce rate and the divorce law reform. In 1997 less than 10 percent of the states allowed domestic violence as grounds for divorce, but by 2001 this proportion had increased to almost 50 percent, and to almost 70 percent by the year 2006. Coinciding with the passage of divorce legislation, divorce rates also increased. In 1993 there were 0.41 divorces per thousand persons and this statistic rose to 0.76 divorces per thousand persons in 2005. According to Mexican statistical yearbook statistics, in 2001-2006 mutual consent accounted for over 70 percent of divorces, while separation or abandonment accounted for 5-10

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[5]Authors' own calculations based on data from INEGI: http://www.beta.inegi.org.mx/programas/nupcialidad/.

percent. In addition, while the majority of divorces are judicial, there has also been a rise in the number of administrative divorce filings over time. Relative to the other causes for divorce, the proportion of divorce filings listing domestic violence as the cause is slightly over one percent. IPV, however, is a major public health concern affecting Mexico. Based on household surveys in 2003, approximately 44% percent of women living with a partner reported having been a victim of domestic violence (ENDIREH, 2003). When intimate partner violence is allowed as grounds for unilateral divorce, an abused spouse may threaten to use it against the other in order to obtain mutual consent to dissolve the marriage.

Oportunidades Program

The Mexican government initiated in 1997 a conditional cash transfer program named PRO-GRESA, renamed *Oportunidades* in 2001 under the Fox Administration (and renamed PROS-PERA under the Peña Nieto administration), aimed at alleviating poverty and improving the human development of children in rural Mexico. The program targets the poor in marginal communities, where 40 percent of the children from poor households drop out of school after the primary level. The program has expanded considerably since its inception and has become an integral component of Mexico's social development and poverty reduction efforts. As of 2013, Oportunidades provided cash transfers to 6.5 million families, conditional on children school attendance, health checks, and participation in health clinics.

The targeting of the program was done at two levels. First, eligible localities were identified on the basis of a locality-level eligibility rule. Program officials used locality-level characteristics from the Mexican 1995 Mini-Census of Population to construct a marginality index for each locality that reflected its degree of marginalization and was correlated with the community's incidence of poverty.[6] Second, program enumerators conducted household surveys within eligible locali-

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[6]The variables used to construct this marginality index were: (i) the locality's population, (ii) the number of dwellings in the village, (iii) the proportion of the adult population who was illiterate, (iv) the proportion of adults working in the agricultural sector (in 1990), the proportion of households (v) without potable water, (vi) without drainage, (vii) without electricity, (viii) with a dirt floor (in 1990), and (ix) the average number of persons per room in each household (in 1990).

ties to identify households that would be classified as poor. Based on asset holdings used as proxy variables for poverty, the program administrators generated a proxy-means test.[7] Therefore, within each eligible community, only households below a threshold became program beneficiaries. The list of potential beneficiaries was then discussed in a community meeting and suggested revisions sent to the central Oportunidades office. In practice, very few changes were made to the list of targeted households (Skoufias, Davis and De La Vega, 2001). This targeting and program eligibility information is important in the construction of our sample of eligible women (see Sections IIII and IV.B).

Initially, a locality was eligible for Oportunidades if it was classified as "poor" (marginality grade 4) or "very poor" (marginality grade 5) out of a 1-5 scale based on the locality-level marginality index, and if it had access to a primary school, a secondary school, a health center, and was classified as rural (defined as inhabited by fewer than 2,500 people), but had at least 50 inhabitants (Skoufias, Davis and De La Vega, 2001). The last criterion was relaxed early on to incorporate some semi-urban localities (localities with between 2,500 and 14,999 inhabitants). The health center criterion was relaxed in 1998 when mobile health clinics were introduced. Since then, the inclusion of less marginal localities into the program has been gradually extended. Between 2000 and 2011, the program's coverage expanded from around 53,000 localities and 2.5 million families, to 97,000 localities and 5.8 million families. The program was phased-in through a different targeting design in urban areas starting in 2001. Since this targeting mechanism is very complex and substantially different to the one implemented in rural and semi-urban areas, and in order to maintain a sample comparable to that of the short-run study, we focus our analysis on rural households. We present additional information on the Oportunidades program in the appendix.

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[7]Within a sub-sample of communities, a poverty indicator was constructed using household income data collected from baseline surveys. A discriminant analysis was then separately applied in each region in order to identify the household characteristics that maximized the correct classification of as poor and non-poor (minimizing Type I and Type II targeting errors). Eligible households were identified on the basis of this welfare index (see Skoufias, Davis and De La Vega 2001 for a more detailed description of the targeting process).

3   Data, Measurement, and Summary Statistics

Description of the ENDIREH Surveys and Study Samples

We use data from Mexico's National Surveys on Relationships within the Household ("Encuesta Nacional sobre la Dinámica de las Relaciones en los Hogares", or ENDIREH) of 2003, 2006 and 2011. These are three cross-sectional, nationally representative household surveys measuring the prevalence and intensity of intimate partner violence, among other intra-household interactions. It contains data on household demographics, socio-economic characteristics, (limited) marital histories, household decision-making, marital conflict, and a module designed to measure the prevalence and severity of intimate partner violence. The 2003 survey was administered to 54,230 women 15 years or older living with a husband or partner, whereas the 2006 and 2011 surveys were administered respectively to 113,561 and 152,636 women in the same age range but independent of marital status.

We construct measures of incidence of violence that consist of dichotomous variables indicating whether the female partner had suffered physical or sexual violence from her spouse or partner in the past 12 months.[8] Physical violence includes pushing, kicking, throwing objects, hitting with hands or objects, choking, attacking with a knife or blade, and shooting. Sexual violence includes demanding sex against the woman's will, forced sexual acts, and forced sexual relations. A single incident reported within the past year is classified as violence.

Data on program participation comes from the ENDIREH surveys and is self-reported by women. The measure of program participation available in the ENDIREH 2003 is whether the woman receives benefits from any government support program at the time of the survey. Although Oportunidades is the largest and most generous cash transfer program, there are other small gov-

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[8]These definitions follow closely the documentation and results of the survey in Castro Roberto and Medina (2006). A previous version of this paper also included emotional abuse as an outcome, but analogous with the results for physical and sexual violence, we see no differences in the incidence of emotional violence between beneficiary and non-beneficiary women in the 2006 and 2011 surveys. These results are available upon request.

ernment programs that provide non-cash benefits. As a result, this measure may over-report the receipt of Oportunidades benefits. Nonetheless, although there is some noise in the data (since only ten households per village are randomly selected to participate in the survey) the correlation of the proportion of beneficiary households using the ENDIREH survey data with administrative data on the number of recipient households at the locality level in 2003 is 0.84 (not reported in the tables), which suggests that the information from the household survey closely represents receipt of Oportunidades benefits. In addition, since program receipt is measured at the time of the survey, it combines couples who have received the transfer for varying lengths of time (which we unfortunately do not observe) and our results should be interpreted with this in mind.

The ENDIREH 2006 and 2011 surveys ask women specifically whether they receive benefits from Oportunidades, and separately whether they are beneficiaries of other government support programs. In order for the analysis to be comparable to that using data from the ENDIREH 2003 survey, the measure we use is the analogous measure of being a beneficiary from any government support program (i.e., Oportunidades or other). In the samples for the ENDIREH 2006 and 2011 selected for our analysis, only between 1.6 and 3.0 percent of those who report being beneficiaries of any government support program, report not being beneficiaries of Oportunidades. These reliability checks suggest that the information from the household survey closely represents receipt of Oportunidades benefits.[9]

In order to minimize potential selection biases as a result of the targeting and endogenous take-up of the program, we restrict the analysis of the short-term relationship in 2003 to a particular subset of households as in Bobonis, González-Brenes and Castro (2013). The 2003 sample includes couples with women 25 years or older, with children younger than 11 years old, and who have been married since at least 1997. These restrictions result in a sample of 2,613 couples. For our analysis in subsequent years, we first construct a *pseudo-panel* of comparable households. We

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[9]We also estimate analogous models using the ENDIREH 2006 and 2011 data with the Oportunidades beneficiary indicator as the explanatory/treatment variable of interest. The results do not differ in any significant way from those reported in the tables. These are available from the authors upon request.

restrict the 2006 (2011) survey sample to couples with women 28 (33) years or older with children between the ages of 3 and 13 (8 and 18) years, who have been together since 1997. The resulting overall sample sizes for the pseudo-panel are 4,240 in the 2006 survey and 5,208 couples in the 2011 survey. As we will discuss below, these sample restrictions minimize potential confounding due to endogenous take-up of the program based on household socio-economic characteristics and preferences for human capital investments (see Section 4).

In an alternative and complementary exercise, we construct a *replication sample* for the subsequent surveys that consists of women with the same characteristics at the time of the survey, that is, women 25 years or older, with children younger than 11 years old, who have been in a relationship for at least six years (this last restriction is analogous to women being married since 1997 for the 2003 sample). Note that while the pseudo-panel approach tries to maximize the overlap of women across the samples, this alternative approach will include many new women and will exclude others that no longer meet the selection criteria (for instance, if the children are now out of primary school). The resulting sample sizes for the replication sample are 4,318 for the 2006 survey, and 5,931 for the 2011 survey.

It is worth highlighting that each of these samples suffers from specific limitations. In particular, the relationship between receipt of the CCT and IPV estimated using the pseudo-panel conflates the policy impacts with cohort effects and the time path of divorce, while that obtained using the replication sample conflates the policy impacts with changes in eligibility and the expansion of the program. However, as we discuss below, the descriptive patterns we observe are very similar in both subsamples.

Finally, to analyze the effect of the IPV divorce law reforms on the divorce rate in our sample, we use the 2011 ENDIREH survey and construct a *state-level panel*. We start with the 2011 pseudo-panel, the sample of women who remain in union in 2011 and who match our sample restrictions (5,208 women), and pool it with the sample of women who divorced between 1998 and 2011 but who share the same restricted characteristics as the initial sample (in 2003 these women were 25 years or older and had children younger than 11, consisting of 472 women). The

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pooled sample has 5,680 observations. We use data on the years of divorce in the pooled sample to estimate a divorce-rate for each state-year cell. We do this for the overall sample, as well as separately for both beneficiaries and non-beneficiaries of Oportunidades (at the time of the survey). The divorce-rate panel consists of 448 observation cells (32 states x 14 years).

Descriptive Statistics

Spousal violence remains a pervasive phenomenon in rural Mexico, but one that has decreased considerably throughout the period (Table 1). Whereas 16.5 percent of women in the sample reported experiencing some form of physical or sexual spousal violence in the year 2003, the incidence had decreased to 14 percent by the year 2006 and to 10.4 percent by 2011, in the pseudo-panel sample. The reductions are even larger for the younger cohorts; in the replication sample the incidence was 13.5 and 8.8 percent for 2006 and 2011, respectively. We observe a similar reduction in the incidence of sexual violence across survey rounds. The patterns in our selected sample are slightly more pronounced but generally consistent with overall trends in Mexico.[10]

Households in the sample are of relatively low socio-economic status. More importantly, we observe some stark differences in a number of dimensions of socio-economic status as we compare the samples of couples across survey years. A significant share of women report speaking an indigenous language (14 percent in 2003, 17/16 percent in 2006, and 21/19 percent in 2011, for the pseudo-panel/replication sample); this ethnic identity is highly correlated with low socio-economic status in Mexico (Table 2, Panel A). In addition, approximately 9 percent of women in 2003 have no schooling, and this figure increases to 15 percent among the pseudo-panel couples selected in 2011. The average age of women in the sample is 35.4 years in 2003, 38/35.2 years in 2006, and 43/35.1 years in 2011. The trend in age for the pseudo-panel is explained by the

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[10]According to estimates based on the ENDIREH 2003-2011 survey rounds, the overall incidence of physical violence among women was estimated at 9.3 percent in the year 2003, increased slightly to 10.3 percent in the year 2006, but decreased substantially to 4.4 percent in the year 2011. The overall incidence of sexual violence has decreased consistently across rounds – from 7.8 percent in the year 2003 to 6.0 and 2.8 percent respectively in the years 2006 and 2011 (Casique and Castro, 2014, p.193).

age restrictions imposed on the samples and is consistent with our attempt of following the same set of women over time. Male partners belong to the same age group (the average partner age is 38.4 years in 2003, 46.5/38.7 years in 2011), have similar schooling attainment levels, and are as likely to have an indigenous background (Table 2, Panel B). Households are relatively large, with more than 5 members on average, a statistic usually correlated with low socioeconomic status in the Mexican context.

The proportion of women who report having been exposed to intra-household violence between their parents during childhood is quite high, at approximately 9 percent in 2003, 11/10 percent in 2006, and 13/11 percent in 2011 (Panel A). Given the existing concerns and evidence regarding the intergenerational transmission of violent behavior, this suggests that women in this context may be at a particularly high risk of experiencing spousal violence, helping explain the prevalence of violence reported above. The proportion of women reporting that their male partners were exposed to spousal violence between their parents during childhood is also significant, but decreases, from 18 percent in 2003, to 13/12 percent in 2006, and to 12/10 percent in 2011. These are also important predictors of spousal abuse among current partners (e.g., Bowlus and Seitz 2006; Casique and Castro 2014).

4   Empirical Methodology

Relationship between Divorce Laws and Divorce Rates

To study the relationship between the introduction of the IPV divorce laws and the incidence of divorce we exploit the differential timing at which these reforms were introduced across Mexican states in a difference-in-differences framework. We use data on marital and family histories from the last wave of the ENDIREH survey (2011) to identify all women who matched our pseudo-panel characteristics (older than 25 and with young children in 2003, in union in 1997). With this information we construct a state-year panel of divorce rates as described in section 3. We then estimate regressions of the following form:

$$Y_{st} = \alpha + \theta PostDL_{st} + \gamma_s + \gamma_t + \varepsilon_{st} \qquad (1)$$

where Y_{st} is the divorce rate in state s and year t, and $PostDL_{st}$ is an indicator equal to 1 if the divorce law had been introduced in state s by year t. Our preferred specification includes state fixed effects γ_s to control for time-invariant heterogeneity across states and year fixed effects γ_t to control for time-shocks shared by all states. In some specifications we also include linear state time trends.

In addition, we use variation across states on whether the reform had been introduced in 2003 (year of the first survey wave) to investigate whether the introduction of these reforms led to differential divorce rates depending on respondent characteristics: violence in woman's childhood, violence in marriage, and Oportunidades beneficiary status. The baseline regression equation is the following:

$$Y_{is} = \alpha + \theta_0 DL2003_s + \theta_1 DL2003_s * V_i + \theta_2 V_i + X_{is}\beta + \varepsilon_{is} \qquad (2)$$

where Y_{is} is and indicator equal to 1 if individual i from state s was divorced in the measured year (2003, 2006, or 2011); $DL2003_s$ is equal to one if the IPV divorce law had been introduced by 2003 in state s, V_i is the covariate of interest, X_{is} are the pre-determined covariates, and ε_{is} are unobserved determinants of divorce. Regressions are weighted by inverse sampling weights aggregated at the state level. We cluster standard errors at the state level.

Relationship between Divorce Laws and Intimate Partner Violence

To investigate the relationship between spousal abuse and IPV divorce laws, we estimate models analogous to those from equation (2) but with IPV as our dependent variable. In addition, we estimate a similar model which includes village fixed-effects that absorb time-invariant characteristics at the village level. Because the village fixed effects are linearly dependent with the IPV divorce law indicator, in this case we can only identify the differential relationship between the reforms

with IPV depending on women's characteristics. Specifically, we estimate the following equation:

$$Y_{isv} = \theta_1 DL2003_s * V_i + \theta_2 V_i + \gamma_v + X_{isv}\beta + \varepsilon_{isv} \tag{3}$$

where Y_{isv} is and indicator equal to 1 if individual i from village v in state s was a victim of physical or sexual violence; $DL2003_s$ is equal to one if the IPV divorce law had been introduced by 2003 in state s; V_i is the covariate of interest, α_v are village fixed effects, X_{isv} is the vector of pre-determined covariates, and ε_{isv} are unobserved determinants of divorce. We do the analysis for each of the 2003 sample, and the pseudo-panel and replication samples for 2006 and 2011. We cluster standard errors at the state level.

Relationship between Oportunidades Beneficiary Status and Intimate Partner Violence

To obtain robust estimates of the relationship between Oportunidades beneficiary status and the incidence of spousal abuse, we estimate ordinary least squares models conditioning on a large set of pre-determined individual and household socio-economic characteristics as well as village fixed effects, in order to capture any village-specific unobserved heterogeneity influencing spousal abuse patterns (e.g., access to health clinics, community groups, village-level conditions affecting partners' socio-economic conditions and economic opportunities). The regression equation for outcome Y_{iv} is the following:

$$Y_{iv} = \theta T_{iv} + X_{iv}\beta + \alpha_v + \varepsilon_{iv} \tag{4}$$

where the treatment indicator T_{iv} equals one for beneficiary household i in village v and is zero otherwise; X_{iv} are the pre-determined covariates that are possibly significantly correlated with T_{iv} and Y_{iv}; α_v are village fixed effects, and ε_{iv} are unobserved determinants of domestic violence. Since T_{iv} is measured at the time of the survey, our estimates combine possible short-term relationships, among couples receiving program benefits for short periods, and those that have been receiving program benefits for longer. We cluster standard errors at the village level.

Dealing with Endogenous Selection into the Treatment

We follow Bobonis, González-Brenes and Castro (2013) in conducting the analysis using various strategies to minimize the extent of selection bias in our estimates.[11] First, as mentioned in Section III, our pseudo-panel uses a sub-sample of households with children ages between 0 and 10 in 2003 (ages between 3 and 13 in 2006, and ages 8 and 18 in 2011) and who have been in union since 1997. Second, we condition on a set of pre-determined individual and household socio-economic characteristics which are strongly correlated with determinants of program eligibility and likely capture a large component of the variation determining program take-up, thus minimizing concerns of endogenous program take-up.. In addition, we restrict the sample to women ages 25 and older in 2003 (28 and older in 2006, 33 and older in 2011). Most importantly, to address the endogenous targeting of the program to poor communities, we make comparisons of beneficiary and non-beneficiary households within villages in order to remove all selection based on the village-level targeting of the program. This within-village comparison dramatically reduces the observed selection into the program (Bobonis, González-Brenes and Castro, 2013).

In order to address the potential concerns of unobserved heterogeneity in the within-village household comparison, we pursue a set of tests and sensitivity analyses inspired by the work on diagnostics of selection on observable and unobservable variables (i.e., Imbens, 2003, 2004; Altonji, Elder and Taber, 2005). Essentially, we identify which observable characteristics (X_{iv}) that are correlated with treatment assignment (T_{iv}) - the woman's age, partner's age, partner's schooling, family size, and years in union - are also significant predictors of spousal abuse outcomes. These may plausibly be the covariates most correlated with the unobservable characteristics that jointly determine program eligibility/take-up and violence outcomes. For those identified variables, we evaluate the robustness of the results to flexible specifications that allow for high-order and interaction terms between these variables, and also include interactions with the woman's levels of education. The results obtained from this sensitivity analysis are qualitatively and quantitatively

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[11]In the context of IPV, Bajracharya and Amin (2013) highlight the importance of accounting for selection bias in a study of microcredit in Bangladesh.

similar across specifications.

5   Results

Relationship between Family Violence Divorce Laws and Divorce Rates

We begin our analysis by looking at the relationship between the legal reforms instituted across Mexican states and divorce rates. The results from estimating equation (1) are presented in Table 3. We present estimates for the overall sample, as well as for specific subsamples of women. Our preferred estimate includes both state linear time trends and other legal reforms related to domestic violence which were instituted around the same time.[12] The analysis suggests that legal reforms led to a statistically significant increase in divorce rates across adopting states. The average annual divorce rate in our sample is 0.64 per 100 couples. The estimated coefficient suggests an increase of 0.37 divorces (column 2), or about 60 percent. Women who are beneficiaries of Oportunidades have slightly lower divorce rates (0.59 per 100 couples), but experience a 0.49 rate increase following the reforms. Women who report violence in their marriage have significantly higher divorce rates on average, 2.48 for 100 couples per year. We also estimate larger increases in the divorce rates following the institution of the legal reform: a 1.39 divorce rate increase (significant at the 90 percent level), a 56 percent increase in proportional terms.

In a complementary exercise, we present estimates from an event-study design in which we allow the coefficients to vary flexibly relative to the timing of the reforms. We present these estimates in Figure 3, including an analysis in which we estimate divorce rates separately for beneficiary and non-beneficiary women. The estimates reveal a pattern consistent with those above and suggest that the effects of the legal reform on divorce was concentrated on beneficiary women.

The second part of our analysis presents estimates from the regression model outlined in equation (2). This analysis uses individual data and exploits one particular threshold regarding the

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[12]These include criminalization of domestic violence and the establishment of prevention and assistance programs for victims of domestic violence. These policies are documented and studied in Beleche (2017).

timing of the legal reforms, we compare states that had, by 2003 instituted the divorce law reforms, relative to those which had not. The advantage of this framework is that it allows us to include our control variables in the estimating equation. The results are presented in Table 4. Women in states that had instituted the reforms by 2003 were more likely to be divorced by 2006 (column 2, statistically significant at the 90 percent level) and by 2011 (column 6, not statistically significant), by about 1.3 percentage points. Interacting the presence of the reform with characteristics of these women suggests that the relationship between the legal reform and the propensity to divorce is heterogeneous across different subgroups of women. In particular, women who experienced violence in their households as children (columns 3 and 7), women who experience violence in their marriage (columns 4 and 8), and women who are beneficiaries of Oportunidades (columns 5 and 9), are more likely to be divorced in legal reform states by 2006 and by 2011, relative to their converse groups in non-reform states. The results from this complementary analysis confirm the findings from the state-panel regressions.

Relationship between Divorce Laws and Intimate Partner Violence

The results from this analysis are presented in Table 5. Columns 1, 4, and 7 evaluate whether women who lived in states that had passed the legal reforms by 2003 experienced spousal violence at different rates relative to women in non-reform states. We find no significant relationship between IPV divorce laws and spousal violence in this cross-sectional analysis. This suggests that the reforms by themselves did not substantially change IPV rates and/or that states with higher rates of IPV were not necessarily the first ones to pass these reforms.

We next test whether the relationship between IPV divorce laws and spousal violence is heterogeneous depending on whether women experienced violence in their homes as children. In 2003, women who were in reform states and who had experienced violence as children were between 14 and 17 percentage points more likely to be victims of spousal abuse (columns 2 and 3) relative to women who had not experienced violence as children. We also observe that the relationship between violence in childhood and domestic violence appears significant in the 2006 and

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2011 samples but is not restricted to reform states.

Relationship Between Oportunidades Beneficiary Status and Intimate Partner Violence

We start the presentation of the relationship between Oportunidades and IPV with a graphical analysis of the patterns in the data. Figure 4 shows the trends in physical violence among couples across the three survey years. The incidence of physical abuse among women in non-beneficiary households is quite high at almost 13 percent in 2003, and shows a downward trend over time to around 10 percent in 2006 and to around 7 percent in 2011 (significant at the 90 percent confidence level). In comparison, the incidence among beneficiary couples is around 9 percent, lower than that among non-beneficiary couples in 2003 (4 percentage points, not significant). However, the incidence among beneficiary couples hovers around 10.0 percent in 2006 and 7 percent in 2011, such that physical violence rates among these two groups of households converge in the longer-run. We observe a similar although less stark pattern of short-run differences (in 2003) and later convergence (in 2006, 2011) in the incidence of sexual violence.

Estimates of the overall five-year (2003), nine-year (2006), and thirteen-year (2013) relationship between program beneficiary status and spousal violence outcomes are displayed in Table 6. For purposes of comparison, we start the discussion with our preferred estimates of the short-run relationship. As documented in Bobonis, González-Brenes and Castro (2013), domestic violence incidence rates in the short-run are significantly lower among beneficiary couples than among non-beneficiary ones (column 1), the estimated difference in the incidence of physical or sexual abuse is 12.5 percentage points (75 percent). We then investigate whether this relationship is heterogeneous between states which had instituted the IPV divorce laws by 2003 and those that had not (column 2). Although we do not have the precision to detect a differential relationship (the interaction term between Oportunidades beneficiary and IPV divorce law states is large but not statistically significant), the overall relationship in IPV law states is statistically significant and large, suggesting that women in reform states who were beneficiaries of Oportunidades were 18.5 percentage points less

likely to be victims of physical or sexual abuse relative to non-beneficiary women in reform states. The analysis suggests that there is a strong short-run complementarity between the Oportunidades program and the legal reform in terms of reductions of IPV, and that the short-run reductions in domestic violence previously documented were concentrated in these reform states.

In contrast to the short-run relationship, domestic violence incidence rates do not vary significantly between beneficiary and non-beneficiary households in 2006 and 2011 (columns 3-6). Our preferred estimate for 2011 shows a statistically insignificant difference in the incidence of physical and sexual violence of 1 percentage point (column 5).[13] The heterogeneity analysis for 2006 suggests that beneficiary women in reform states were 4.1 percentage points *more* likely to be victims of physical or sexual abuse relative to non-beneficiaries in non-reform states (column 4). The finding could be partially explained by a combination of the strong selection mechanism combined with a "male backlash" effect. This relationship is, however, not present in neither the replication sample for 2006 (Panel B), or either of the 2011 samples (column 6). The interaction coefficient for 2011, though statistically insignificant, is, as in 2003, negative. In some of the specifications which include additional control variables, this coefficient becomes statistically significant (shown in Table 8 and discussed in the appendix), suggesting that to some extent, this complementarity between IPV divorce laws and the CCT program persists through to 2011.

Discussion and Empirical Extensions

Repeated Interactions and Marital Selection

The stark differences in the longitudinal pattern of the relationship suggests that the models of violence and household bargaining, in which male partners may use violence as instruments of coercion (see Bloch and Rao, 2002; Bobonis, González-Brenes and Castro, 2013; Anderson and Genicot, 2015) may correctly capture short-run interactions within the household but may do so

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[13]We also find no evidence of a significant difference in the incidence of violent threats or acts of emotional violence among beneficiary women, a finding documented in a previous version of this article (Bobonis, Castro and Morales, 2015).

poorly in the longer-run. In this class of models male partners are heterogeneous in their willing-ness to engage in violence and have private information regarding the "gains to marriage", such as their own private income or their status within the household based on traditional gender roles. However, once this information has been revealed through partners' actions, couples with violent types may dissolve such that couples in future periods are disproportionately composed of non-violent types. This selection patterns can lead to both a tendency for violence rates to decrease among couples remaining in union over time and for their relationship with program receipt status to be dampened.

The evidence we present is consistent with this interpretation. First, we have documented that the short-run relationship with Oportunidades receipt is concentrated in reform states. We have also shown that women in reform states experienced an increase in marital dissolution rates over this time period. Furthermore, the increase in marital dissolution rates was larger for beneficiary women and women who experienced violence in their marriage. This increase in divorce-based se-lection can help explain both the drop in the incidence of violence among these cohorts of women and the absence of a relationship between the Oportunidades program and the incidence of IPV in the subsequent surveys. In the appendix we present a formal model of spousal relations that high-lights the mechanisms through which legal reforms and conditional cash transfers affect divorce and conflict within the household, and reveals how different types of couples remain in union when the policies are implemented.

Second, a comparison of socio-economic and demographic characteristics of households across the three survey waves suggests that there are important changes in their distributions (see Table 2). Women are more likely to report speaking an indigenous language in the later survey waves (which is correlated with low socio-economic status in Mexico), they tend to have lower ed-ucational attainment levels, and report a higher prevalence of violence in their households during childhood. Moreover, their male partners are more likely to be indigenous themselves. However, they report a lower prevalence of violence in their male partners' households during childhood - decreasing from 18 percent in 2003 to around 12 percent in 2006 and 2011. Given the strong corre-

lation in the intergenerational transmission of violent behavior, the decrease in this statistic may be informative of the substantial drop in the incidence of violence observed in the sample across survey waves. The differences in the distribution of these pre-determined characteristics across waves reveal significant changes in sample composition, and most importantly, on the types of couples which remain in union.[14] Finally, levels of violence among couples in the replication sample tend to be lower than those in the pseudo-panel, consistent with the view that abuse in new couples may be lower. To the extent that changes in the composition of beneficiary households is driven by these marital selection dynamics, these sample selection and treatment effect heterogeneity patterns help explain the time path of spousal violence among beneficiary couples.

Increasing Rejection of Intimate Partner Violence

Recent research has documented a rapid global diffusion of norms regarding the unacceptability of spousal violence across a broad set of countries. Specifically, Pierotti (2013) uses nationally representative, repeated cross-sectional data from Demographic and Health Surveys (DHS) across a broad set of low and middle-income countries to document that women of reproductive age have increasingly rejected the justification of violence from intimate partners. She argues that new global cultural scripts rejecting violence against women - via international and national policies and discussions starting in the mid/late 1990s - may then be reflected in modifications of individual attitudes towards IPV across a large spectrum of societies. These new global scripts and norms may have also diffused across Mexican society in such ways as to decrease women's tolerance for IPV.
[15]

To evaluate this hypothesis, we use additional information available in the ENDIREH data.

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[14]See Casique and Castro (2014) for a rich analysis of changes in household socio-economic characteristics and patterns of intimate partner violence across the three survey waves.

[15]Suggestive of this phenomenon in the Mexican context is the passage of laws promoting gender equality and establishing the right of women to live free of violence in 2006 and 2007, respectively. Reports in the 2011 ENDIREH survey that 73 percent of women are knowledgeable of the gender equality legislation and 82 percent of women report being knowledgeable of the freedom from violence legislation are consistent with a strong dissemination of these scripts as embodied in national policy.

Following the analysis in Pierotti (2013), we construct an indicator variable that measures whether the woman believes an intimate partner is justified in hitting or beating his female partner when she does not meet her responsibilities.[16] These measures are imperfectly comparable to those from existing DHS data.[17] The proportion of women in our sample who reported a husband "has the right" to hit his wife decreased from approximately 20 percent in 2003 to approximately 9 percent in 2006 and 3 percent in 2011.[18] Figure 6 shows the trend in this measure among couples in the sample across the three survey years, by beneficiary status. Consistent with the cross-country evidence, the proportion of women reporting some justification of IPV shows a sharp reduction over this time period. This stark change in the justification of IPV occurs among women in both beneficiary and non-beneficiary households. For the pseudo-panel we observe a decrease of 12.4 percentage points (54 percent; significant at 95 percent confidence) - from 22.8 percent in 2003 to 10.4 percent in 2006 - among women in beneficiary households, and a similar change of 9.9 percentage points (58 percent; significant at the 95 percent confidence level) among those in non-beneficiary households. We estimate further decreases of proportional size between the 2006 and 2011 survey rounds. These patterns also hold for the replication sample (Figure 6, right), as the forces driving these patterns are present across Mexican society.

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[16]The 2003 and 2006 survey rounds ask the same question: "En su opinión, cuando la mujer no cumple con sus obligaciones, el marido tiene el derecho de pegarle?" [In your opinion, when a women does not meet her responsibilities, the partner has the right to hit her?]. In contrast, the question in the 2011 survey round is modified: "El hombre tiene el derecho de pegarle a su esposa?" [Does a man have the right to hit his partner?] Therefore, the responses in the 2011 survey round are not strictly comparable to those in earlier rounds. We report these in order to show a more complete picture, subject to this caveat.

[17]Specifically, Pierotti (2013) constructs outcome variables derived from questions that asked respondents whether it is okay for a man to hit or beat his wife under certain circumstances. Specifically, the most common form of the question asked, "Sometimes a husband is annoyed or angered by things which his wife does. In your opinion, is a husband justified in hitting or beating his wife in the following situations?" The five scenarios presented to respondents were (1) if she goes out without telling him, (2) if she neglects the children, (3) if she argues with him, (4) if she refuses to have sex with him, and (5) if she burns the food." (Pierotti, 2013, p. 248).

[18]However, the question was framed different in 2011. For 2003 and 2006, the question asked whether women agreed with the statement, "If his wife does not meet her duties, does the husband have the right to hit her?", in 2011, the question read "Does a man have the right to hit his wife?". In Spanish, "¿Cuando la mujer no cumple con sus obligaciones, el marido tiene el derecho de pegarle?" and "¿El hombre tiene el derecho de pegarle a su esposa?".

Table 7 presents estimates of the relationship between Oportunidades and Acceptability of IPV (analogous to those in Table 6 but with this measure as the dependent variable). We find that the relationship between the CCT program and acceptability of IPV is heterogenous across Mexican states. In particular, the interaction between Oportunidades and the IPV legal reforms is negative and significant, suggesting that beneficiary women in reform states are 14.5 percentage points less likely to report that husbands are justified in hitting their wives, relative to those in non-reform states (column 2). This evidence provides additional support to the complementary roles of legal reforms and the CCT program in reducing IPV. The negative relationship persists in 2006 but becomes statistically insignificant (column 4). In 2011, we find an overall negative relationship between the Oportunidades program and Acceptability of IPV for the replication sample (columns 5 and 6, Panel B).

Alternative Explanations

A first alternate explanation is that the program empowers women in the community and provides them with the instruments to prevent spousal abuse, directly via interactions with beneficiary women with higher levels of empowerment in the community, or indirectly via improved socio-economic conditions and options outside of current relationships or changes in the norms of intolerance of abuse, among other mechanisms.[19] Therefore, to the extent that these spillover effects reduce the incidence of abuse among non-beneficiary women and increase female partners' intolerance of abuse, this can help explain the patterns shown earlier.

We evaluate this alternate explanation empirically by estimating empirical models that capture spillover effects at the level of the village. Specifically, we estimate a variant of our main empirical model (1). The regression equation incorporating these effects is the following:

$$Y_{ivm} = \theta_1 T_{ivm} + \beta_1 E[T_{-i,v,m}] + X_{ivm}\beta_2 + \alpha_m + \varepsilon_{ivm} \tag{5}$$

[19]As shown by Angelucci and De Giorgi (2009) and Avitabile (2012), the program had spillover effects on the consumption levels and health behaviors (i.e., cervical cancer checks) of non-beneficiary households. Bobonis and Finan (2009), Lalive and Cattaneo (2009) show evidence of spillovers effects on middle school participation among children in non-beneficiary households.

where the treatment indicator T_{ivm} equals one for beneficiary household i in village v, municipality m and is zero otherwise; $E[T_{-i,v,m}]$ represents the proportion of beneficiary households in the sample in village v (excluding household i). This specification incorporates the possibility that local spillovers are a (linear) function of the proportion of beneficiary households in the village (e.g., Miguel and Kremer, 2004). These potential effects are captured by the β_1 term. We also estimate additional specifications that allow for heterogeneous spillover effects among beneficiary and non-beneficiary households by including an interaction term between T_{ivm} and $E[T_{-i,v,m}]$. Since the $E[T_{-i,v,m}]$ term is highly collinear with village fixed effects, we substitute these for municipality fixed effects in these specifications.

We report estimates of these models in the online appendix Table 8, columns 1-3; Panel A reports estimates for the 2006 data whereas Panel B reports analogous ones for the 2011 data. The estimates imply that a 10 percentage point increase in the proportion of beneficiary women leads to a statistically insignificant 0.3 percentage point increase in the incidence of physical or sexual abuse in 2006, and a 0.2 percentage point decrease in its incidence in 2011 (column 1). In the specification allowing for heterogeneous spillover effects by beneficiary status, the point estimates imply a statistically insignificant decrease in spousal violence for women in villages with larger shares of beneficiaries (column 3). Finally, it is worth noting that issues of unobserved heterogeneity generally cause upward bias (in absolute magnitude) in the estimates of the spillover effects; such that these can be considered overestimates of the true spillover or social interaction effects. We conclude that this alternative mechanism cannot explain the results.

In the appendix we evaluate and discuss other explanations potentially consistent with the evidence. In particular, we examine the role of both generalized social violence and women's labor opportunities exacerbating or mitigating the effects on spousal abuse. We do not find evidence supportive of these alternative explanations driving our results.

6 Conclusion

Our paper provides evidence of the relationship between the Oportunidades CCT program, legal reforms codifying intimate partner violence as legal grounds for unilateral divorce, and the prevalence of male-to-female spousal violence in rural Mexico. More broadly, our paper contributes to the debate on how the effects of cash transfer programs on IPV are mediated by their local context, an understudied aspect of empirical studies in the literature (Baranov et al., 2020). Our findings are consistent with the legal reforms both effectively increasing women's threat points in a household bargaining framework, therefore increasing the effectiveness of cash transfer policies in the short-run, but perhaps most importantly, also enhancing the extent to which marital selection shapes outcomes in the longer-run. In stark contrast to the short-run relationships documented in Bobonis, González-Brenes and Castro (2013), we find that, in the longer-run, women in beneficiary households are as likely to experience physical and sexual abuse as non-beneficiary women. We also document how the concurrent legal reforms that eased women's ability to exit relationships led to higher divorce rates, in particular for women who experienced violence in their marriage and for women who were beneficiaries of the Oportunidades program.

Together, the results suggest that the policies reinforced each others' capacity to reduce IPV. As further evidence of these policy complementarities, our analysis revealed that the short-run negative relationship between the CCT and IPV first documented in Bobonis, González-Brenes and Castro (2013) was mostly concentrated in states that were early adopters of the legal reforms. Our findings highlight that Oportunidades no longer appearing to protect women against violence in the longer-run can be explained by these violent relationships dissolving over time. We conclude that the absence of a relationship between Oportunidades and IPV in the more recent surveys should therefore be viewed as a success rather than as a failing of the policies.

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Tables and Figures

Figure 1: Date of introduction of legal reforms across Mexican states
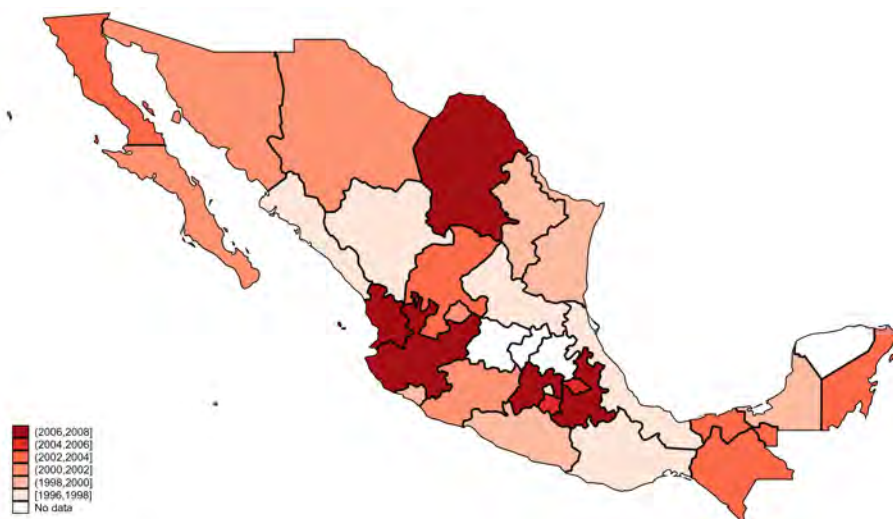
Notes: Map shows the years in which the legal reforms to state civil codes where introduced across Mexico.

Figure 2: Introduction of legal reforms and divorce in the ENDIREH pseudo-panel
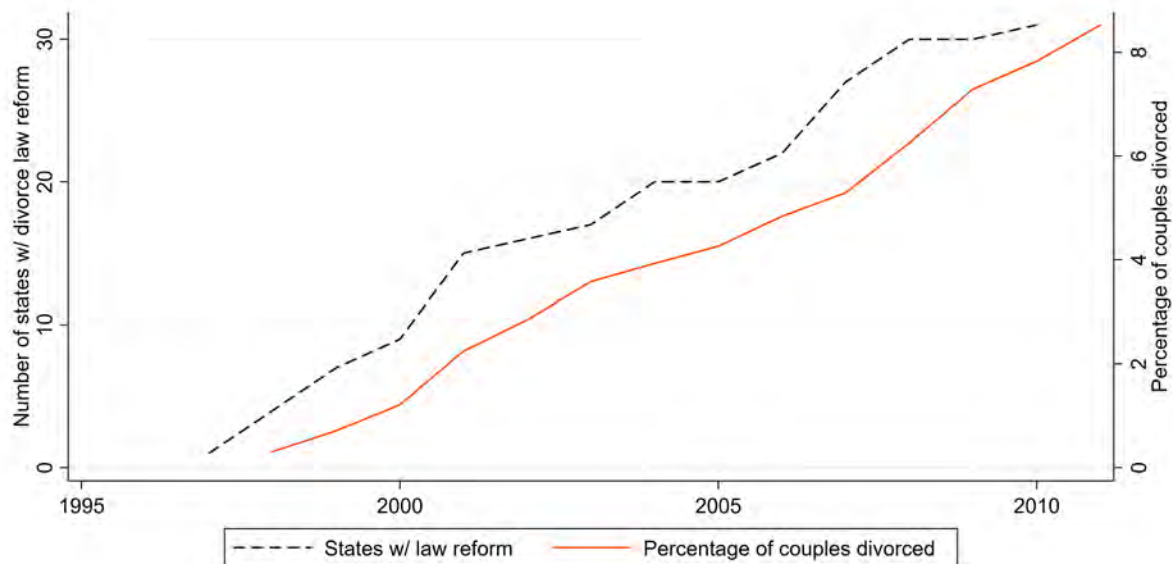
Notes: The figure shows the number of states that had passed the law reform and the percentage of couples in the ENDIREH 2011 pseudo-panel that had divorced between 1997 and 2008

Figure 3: Event study: Divorce laws and divorce rates for women in the ENDIREH 2011 pseudo-panel for all women (Top), beneficiaries of Oportunidades (middle), and non-beneficiaries (bottom)
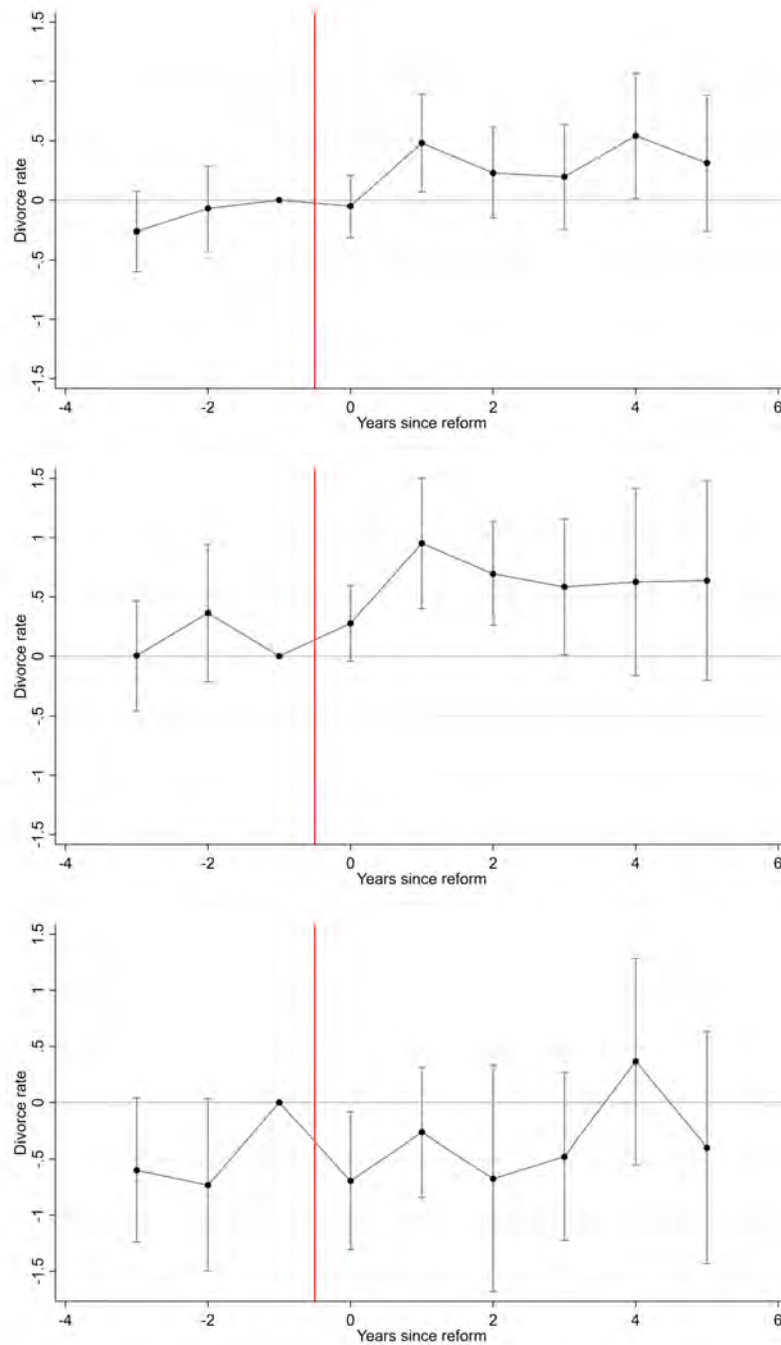
Notes: The figure shows the results of an event-study analysis in which coefficients capture changes in divorce rates relative to the year before the reform. Sample is a state-year panel of women in the ENDIREH 2011 pseudo-panel using inverse sampling weights aggregated at the state level. The first and last coefficients are binned to capture all pre and post observations respectively. Top figure shows all women, middle figure shows women who were beneficiaries of Oportunidades and bottom panel shows women who were not beneficiaries of Oportunidades.

Figure 4: Incidence of physical (top) and sexual violence (bottom) by beneficiary status and survey wave; pseudo-panel (left) and replication sample (right)
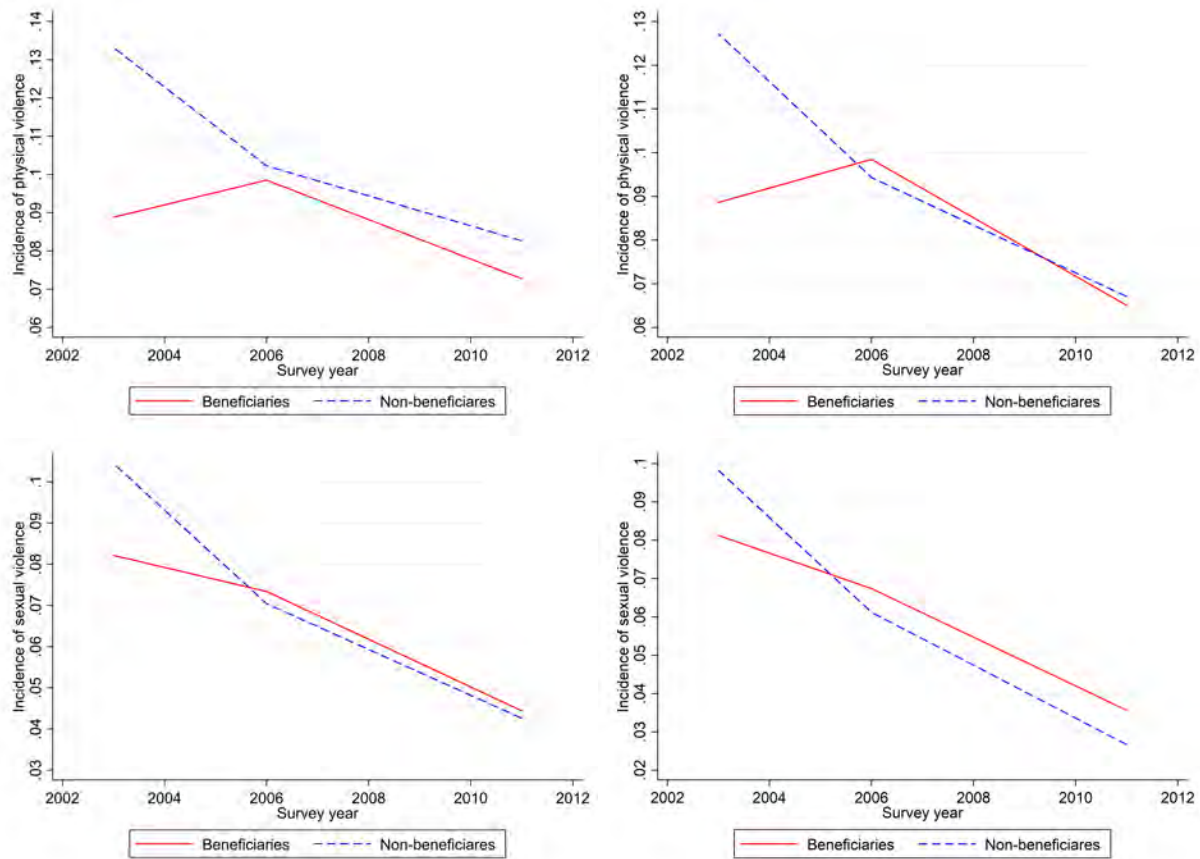
Notes: Shown are sample proportions of women reporting being victims of physical violence (top) and sexual violence (bottom) during the previous year, separately for beneficiaries and non-beneficiaries of government social programs. The pseudo-panel includes couples in rural villages with women ages 25/28/33 and older, with children aged 0-10/3-13/8-18, in a relationship since 1997. The replication sample includes couples in rural villages with women ages 25 and older, with children aged 0-10, who have been in a relationship for six years. Sample proportions weighted by inverse sampling weights.

Figure 5: Acceptability of IPV and survey wave; pseudo-panel (left) and replication sample (right)
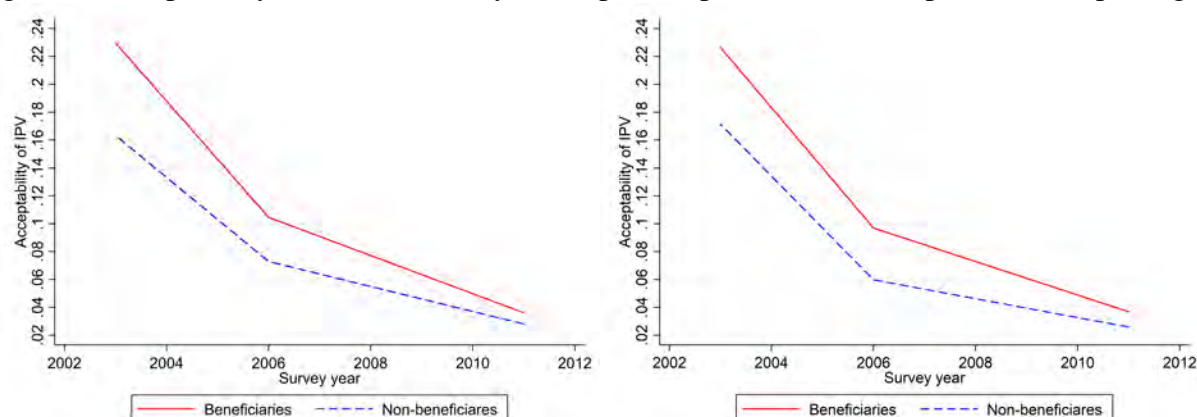
Notes: Shown are sample proportions of women stating that a husband "has the right to hit his wife", for each of the 2003/2006/2011 surveys. The pseudo-panel includes couples in rural villages with women ages 25/28/33 and older, with children aged 0-10/3-13/8-18, in a relationship since 1997. The replication sample includes couples in rural villages with women ages 25 and older, with children aged 0-10, who have been in a relationship for six years. Sample proportions weighted by inverse sampling weights. Sample proportions weighted by inverse sampling weights.

Table 1: Description of outcome variables: Intimate Partner Violence

Variable Name	Description	Sample Means				
			Pseudo-panel		Replication sample	
		2003	2006	2011	2006	2011
		(1)	(2)	(3)	(4)	(5)
Physical or sexual violence	Indicator for any occurance of physical or sexual abuse	0.165	0.140	0.104	0.135	0.088
Physical violence	Indicator for any occurance of physical abuse (e.g., push, beating, attack with blade)	0.110	0.100	0.076	0.097	0.069
Sexual violence	Indicator for any occurance of sexual abuse (e.g., use of force to have sexual relations)	0.093	0.072	0.044	0.068	0.032
Observations		2,613	4,240	5,208	4,318	5,931

Notes: Sample means weighted by inverse sampling weights. The 2003 sample (column 1) includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel (columns 2-3) includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample (columns 4-5) includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years.

Table 2: Descriptive Statistics – Woman, Partner, Couple, and Household Characteristics

	Sample Means				
	Pseudo-panel			*Replication sample*	
Variable Name	**2003**	**2006**	**2011**	**2006**	**2011**
	(1)	(2)	(3)	(4)	(5)
Panel A: Female Partner Characteristics					
Woman's age	35.46	38.02	42.99	35.20	35.08
Indigenous woman	0.14	0.17	0.21	0.16	0.19
No schooling	0.09	0.17	0.15	0.14	0.10
Primary school	0.67	0.58	0.57	0.56	0.52
Middle school	0.17	0.18	0.21	0.22	0.27
Secondary school	0.03	0.03	0.04	0.05	0.07
Violence in woman's childhood	0.09	0.11	0.13	0.10	0.11
Panel B: Partner and Couple Characteristics					
Partner's age	38.40	41.73	46.55	38.86	38.70
Indigenous partner	0.14	0.18	0.21	0.17	0.19
Partner's schooling attainment	5.57	5.01	5.45	5.51	6.38
Violence in partner's childhood	0.18	0.13	0.12	0.12	0.10
Cohabiting couple	0.16	0.21	0.19	0.21	0.28
Family size	5.90	5.99	5.58	5.88	5.48
Years in union	16.25	18.81	24.15	15.79	15.62
Panel C: Other Household Characteristics					
Dirt floor	0.29	0.26	0.12	0.26	0.12
Firm floor	0.63	0.64	0.78	0.64	0.77
Access to water - private tap	0.69	0.63	0.71	0.64	0.69
Access to water - public tap	0.04	0.06	0.04	0.06	0.06
Electricity	0.95	0.96	0.98	0.96	0.98
Computer	n/a	0.04	0.10	0.04	0.07
Telephone (Landline)	0.15	0.21	0.18	0.19	0.12
Radio	0.77	0.77	0.66	0.76	0.64
Drainage	0.69	0.74	0.85	0.73	0.84
Television	n/a	0.82	0.87	0.82	0.87
Num. of bedrooms	1.80	1.97	2.18	1.86	1.84
Num. of rooms	3.01	3.26	3.55	3.15	3.16
Observations	2,613	4,240	5,208	4,318	5,931

Notes: Sample means weighted by inverse sampling weights. The 2003 sample (column 1) includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel (columns 2-3) includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample (columns 4-5) includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years.

Table 3: IPV divorce law and divorce rates (State diff-in-diff)

	Overall		Beneficiaries		Non-beneficiaries		Violence in marriage		No violence in marriage	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
IPV Divorce Law * Post-Period	0.357* (0.178)	0.374** (0.170)	0.405* (0.235)	0.490** (0.238)	0.255 (0.248)	0.197 (0.259)	1.670 (1.002)	1.389* (0.681)	0.220 (0.164)	0.242 (0.203)
State Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State Linear Trends	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Include other reforms	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Observations	448	448	448	448	448	448	448	448	448	448
Mean of dep variable	0.636	0.636	0.589	0.589	0.725	0.725	2.480	2.480	0.456	0.456

Dependent variable: Divorce rate

Notes: The exercise uses the state-panel constructed using the ENDIREH 2011 survey for women in the pseudo-panel: women ages 33 and older, with children aged 8-18, and who were in union since 1997 or earlier. Other reforms controls include states' adoption of the Penal Code Reform, which criminalized domestic violence, and the Law of Access, Assistance and Prevention against Intra-Family Violence, which established government assistance programs to prevent domestic violence (Beleche, 2017). Regressions are weighted by sample size at the state level. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Table 4: Relationship between IPV Divorce Law and divorced status

Dependent variable: Divorced in Year X

	2003	2006				2011			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
IPV Divorce Law (2003)	0.006 (0.005)	0.013* (0.007)	0.000 (0.005)	0.006 (0.008)	-0.001 (0.012)	0.015 (0.012)	0.005 (0.011)	0.004 (0.010)	-0.007 (0.017)
Violence in woman's childhood × IPV Divorce Law (2003)			0.086** (0.031)				0.067 (0.043)		
Violence in marriage × IPV Divorce Law (2003)				0.056 (0.036)				0.091** (0.042)	
Beneficiary status × IPV Divorce Law (2003)					0.020 (0.013)				0.034** (0.016)
Violence in woman's childhood			0.045* (0.024)				0.094** (0.036)		
Violence in marriage				0.077** (0.029)				0.162*** (0.030)	
Beneficiary status					-0.007 (0.009)				-0.035*** (0.008)
Total Effect of IPV Divorce Law (for women in each category)			0.086** (0.031)	0.062* (0.032)	0.0193*** (0.007)		0.071* (0.042)	0.095** (0.045)	0.027** (0.011)
Village Fixed Effects	No	No	No	No	No	No	No	No	No
Individual & household controls (linear)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean of dependent variable	0.036	0.048	0.048	0.048	0.048	0.085	0.085	0.085	0.085
Observations	5,680	5,680	5,680	5,680	5,680	5,680	5,680	5,680	5,680

Notes: Coefficient estimates from OLS regressions weighted by survey sampling weights. Sample includes women both divorced and in union for the ENDIREH 2011 pseudo-panel: women ages 33 and older, with children aged 8-18, and who were in union since 1997 or earlier. IPV Divorce Law (2003) is an indicator equal to 1 if the state had passed the IPV Divorce Law by 2003. Individual controls include indicator variables for woman and partner's age, woman and partner's indigenous status, women's schooling-level indicators, the partner's schooling attainment level, household size, cohabiting couple indicator, years in union, and variables measuring reported histories of spousal abuse in parental household during childhood. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Table 5: Relationship between IPV divorce laws and physical or sexual violence

		2003			2006			2011	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					Panel A: Pseudo-panel				
IPV Divorce Law (2003)	0.026	0.012		-0.006	-0.008		-0.009	-0.008	
	(0.028)	(0.034)		(0.015)	(0.016)		(0.017)	(0.020)	
Violence in woman's childhood		0.166*	0.139*		0.017	0.051		-0.007	0.031
× IPV Divorce Law (2003)		(0.093)	(0.073)		(0.050)	(0.047)		(0.055)	(0.070)
Violence in woman's childhood		-0.028	0.007		0.145***	0.126***		0.117***	0.082
		(0.077)	(0.048)		(0.049)	(0.042)		(0.052)	(0.066)
Total Effect of IPV Divorce Law		0.178**			0.009			-0.015	
(for those w/ violence in childhood)		(0.071)			(0.046)			(0.046)	
Mean of dependent variable	0.165	0.165	0.165	0.140	0.140	0.140	0.104	0.104	0.104
					Panel B: Replication Sample				
IPV Divorce Law (2003)				0.001	0.003		0.003	-0.002	
				(0.015)	(0.017)		(0.009)	(0.013)	
Violence in woman's childhood					-0.018	-0.015		0.049	0.088
× IPV Divorce Law (2003)					(0.054)	-(0.052)		(0.049)	(0.074)
Violence in woman's childhood					0.149***	0.112**		0.048	0.029
					(0.051)	-(0.045)		(0.042)	(0.066)
Total Effect of IPV Divorce Law					-0.015			0.047	
(for those w/ violence in childhood)					(0.047)			(0.040)	
Mean of dependent variable				0.135	0.135	0.135	0.088	0.088	0.088
Village Fixed Effects	No	No	Yes	No	No	Yes	No	No	Yes
Individual & household controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations (Panel A / Panel B)		2867			4240 / 4318			5208 / 5931	

Notes: Coefficient estimates from OLS regressions weighted by survey sampling weights. The 2003 sample includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years. IPV Divorce Law (2003) is an indicator equal to 1 if the state had passed the IPV Divorce Law by 2003. Individual controls include indicator variables for woman and partner's age, woman and partner's indigenous status, women's schooling-level indicators, the partner's schooling attainment level, household size, cohabiting couple indicator, years in union, and variables measuring reported histories of spousal abuse in parental household during childhood. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Table 6: Relationship between Oportunidades and physical or sexual violence

			Dependent variable: Physical or sexual violence			
	2003		**2006**		**2011**	
	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Pseudo-panel					
Oportunidades beneficiary	-0.125**	-0.058	0.026	0.012	0.010	0.021
	(0.050)	(0.054)	(0.020)	(0.037)	(0.017)	(0.028)
Oportunidades beneficiary		-0.126		0.029		-0.022
× IPV Divorce Law (2003)		(0.083)		(0.043)		(0.030)
Total Effect		-0.184**		0.041**		-0.001
(Oportunidades beneficiary in IPV law state)		(0.068)		(0.019)		(0.013)
Mean of dependent variable	0.165	0.165	0.140	0.140	0.104	0.104
	Panel B: Replication Sample					
Oportunidades beneficiary			0.027	0.027	0.005	0.017
			(0.018)	(0.032)	(0.013)	(0.020)
Oportunidades beneficiary				0.000		-0.024
× IPV Divorce Law (2003)				(0.036)		(0.021)
Total Effect				0.028		-0.007
(Oportunidades beneficiary in IPV law state)				(0.017)		(0.012)
Mean of dependent variable			0.135	0.135	0.088	0.088
Village Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Individual & household controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,613		4240 / 4318		5208 / 5931	

Notes: Coefficient estimates from OLS regressions weighted by survey sampling weights. The 2003 sample includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years. IPV Divorce Law (2003) is an indicator equal to 1 if the state had passed the IPV Divorce Law by 2003. Individual controls include indicator variables for woman and partner's age, woman and partner's indigenous status, women's schooling-level indicators, the partner's schooling attainment level, household size, cohabiting couple indicator, years in union, and variables measuring reported histories of spousal abuse in parental household during childhood. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Table 7: Relationship between Oportunidades and Acceptability of IPV

	2003		2006		2011	
	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Pseudo-panel					
Oportunidades beneficiary	0.029	0.106	-0.008	0.012	-0.001	-0.004
	(0.046)	(0.070)	(0.021)	(0.035)	(0.010)	(0.016)
Oportunidades beneficiary		-0.145*		-0.039		0.006
× IPV Divorce Law (2003)		(0.075)		(0.038)		(0.019)
Total Effect		-0.039		-0.028		0.002
(Oportunidades beneficiary in IPV law state)		(0.032)		(0.017)		(0.010)
Mean of dependent variable	0.198	0.198	0.094	0.094	0.033	0.033
	Panel B: Replication Sample					
Oportunidades beneficiary			-0.015	0.004	-0.015**	-0.020**
			(0.016)	(0.023)	(0.006)	(0.010)
Oportunidades beneficiary				-0.036		0.010
× IPV Divorce Law (2003)				(0.028)		(0.012)
Total Effect				-0.032**		-0.010
(Oportunidades beneficiary in IPV law state)				(0.015)		(0.007)
Mean of dependent variable			0.084	0.084	0.034	0.034
Village Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Individual & household controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,613		4240 / 4318		5208 / 5931	

Dependent variable: Acceptability of IPV

Notes: Coefficient estimates from OLS regressions weighted by survey sampling weights. Acceptability of IPV is an indicator on whether the woman believes a partner is justified in hitting or beating his female partner. The 2003 sample includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years. IPV Divorce Law (2003) is an indicator equal to 1 if the state had passed the IPV Divorce Law by 2003. Individual controls include indicator variables for woman and partner's age, woman and partner's indigenous status, women's schooling-level indicators, the partner's schooling attainment level, household size, cohabiting couple indicator, years in union, and variables measuring reported histories of spousal abuse in parental household during childhood. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Appendix

Additional Information on the Oportunidades Program

The Oportunidades program promotes children's human development in education, nutrition, and health. Table 1 presents a summary of benefits for the years 2003, 2006, and 2011, the periods for which we have survey data on interactions among intimate partners. The education component of Oportunidades consists of subsidies typically provided to mothers, contingent on their children's regular attendance at school.[20] Although PROGRESA initially targeted only children in primary and middle school, Oportunidades was expanded to cover children in secondary school. In 1998, these ranged from 70 to 255 pesos per month (approximately 7 to 25 USD), depending on the gender and grade level the child is attending, with a maximum of 625 pesos (62.5 USD) per month per family. Scholarship amounts have gradually increased, and in 2011 these ranged from 150 pesos per month (approximately 12 USD), up to 960 pesos (77 USD).[21] Families also receive yearly benefits for the purchase of school supplies of between 200 and 400 pesos (16 and 32 USD). In a further expansion of the program in 2009, it now offers a cash transfer of approximately 4,200 pesos to youth graduating from high school before age 22 (Jóvenes con Oportunidades).

The health and nutrition components consist of both cash transfers and nutritional supplements. Supplements are targeted at infants 6-months to 23-months old, pregnant and breast-feeding women, and children aged 2-5 years who exhibit signs of malnutrition. Monthly cash transfers for beneficiary families expanded throughout 1997-2011, by 2011 these benefits included: nutritional

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[20]Receipt of the education-specific benefits is contingent on children attending school, which is verified by school personnel. For primary and secondary school, the child becomes ineligible for support if he or she misses school 4 times in a month without justification, or 12 times during the school year. High school students become ineligible if they are not certified as active during the school semester, defined according to the regulations of the institution they attend.

[21]This nominal average value of transfers has gradually increased since the start of the program, and its purchasing power has varied (depending on price levels in these areas and relative price changes with respect to foreign currencies, i.e., USDs) throughout the 1997-2011 period. Given these fluctuations, we opt to report the figure valid at the date of the most recent ENDIREH survey, 2011.

support (Alimentario), 225 pesos (18 USD), originally part of PROGRESA; energy support (Energético), 60 pesos, established in 2007 to help families pay for energy costs (electricity, gas, firewood, etc.); compensated nutritional support (Alimentario Vivir Mejor), 120 pesos, established in 2008 to compensate families for rising food prices; child support (Infantil Vivir Mejor), 105 pesos for every child aged between 0-9, established in 2010; elderly support (Adultos Mayores), 315 pesos for every adult aged 70 or over, established in 2006. These benefits are contingent on participation by mothers in monthly health talks with the local health care provider, the vaccination of family members, health checks of all children under 5 years old, and biannual health checks of all household members. Overall, the program transfers are important, representing approximately 10 percent of the average expenditures of beneficiary families (Skoufias, 2001). Maximum benefit levels have increased by approximately 20 percent over time for families with children in only elementary or middle school, but have almost doubled for those with children in secondary school (see Figure A1).

Alternative Explanations

We evaluate here other explanations that can potentially help us understand the empirical patterns we documented. In particular, we examine the role of generalized social violence and women's labor opportunities exacerbating or mitigating the effects on spousal abuse. For brevity, we present the results from these analyses using the pseudo-panel.

Generalized Social Violence

Another potential alternate explanation we consider is changes in the incidence of abuse or in its reporting due to the marked increase in social violence. As is well known, Mexico has seen a surge in homicide rates since 2007, and it has been concentrated in particular regions of the country. Many analysts attribute this drastic change in the level of violence to consequences of the federal government's anti-crime policies meant to combat drug cartels (e.g., Astorga and Shirk, 2010; Dell, 2015).

We consider the potential for this surge in generalized social violence across municipalities and or states to affect the trends in spousal abuse and the relationship with program beneficiary status. On one hand, to the extent that the surge in homicides can impinge on partners' stress levels or their levels of emotional health more broadly, this could lead to greater conflict-related abuse. On the other hand, if this generalized conflict negatively impinges on women's willingness to report events of abuse, this would be consistent with the significant drop in reported abuse rates in 2011, although not so in the year 2006. These are two potential mechanisms that would induce heterogeneity in the relationship, among others.

We evaluate this idea empirically by estimating models that capture these factors at the level of the municipality or state. The regression equation incorporating these factors is the following:

$$Y_{ivm} = \theta_1 T_{ivm} + \theta_2 T_{ivm} H_{m(s)} + \beta_1 H_{m(s)} + X_{ivm} \beta_2 + \varepsilon_{ivm} \tag{6}$$

The $H_{m(s)}$ variable measures the homicide rate per hundred thousand individuals in municipality m (or alternatively, state s); the other variables are defined as above.[22] The homicides measures are included for the calendar year preceding the household survey (2005 and 2010, respectively) since the surveys are conducted over a long time period and we aim to ensure that the timing of measured homicides predates that of abuse outcomes.[23] The β_1 coefficient captures the partial correlation between homicides and spousal abuse rates among non-beneficiary couples, whereas the θ_2 term captures the differential correlation among beneficiary ones. In our main specification, because the homicide rate is measured at the municipality level, we do not include village fixed effects in this

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[22]The homicide data is available from Mexico's National Statistics and Geography Institute (INEGI). We follow the standard specifications in the literature and estimate the relationship between violence and individual/household outcomes with measures of violence at the municipality level (e.g., Camacho, 2008; Leon, 2012). Empirical models of this sort find strong relationships with adult labor force participation (BenYishay and Pearlman, 2013) and student achievement Michaelsen and Salardi (2020) in Mexico.

[23]The results are robust to using contemporaneous year measures (2006 and 2011, respectively). These are also robust to the use of gender-specific homicide rates, in spite of the different trends among the victim's gender, shown in Valdivia and Castro (2013). Estimates are available upon request.

specification. In a second specification with village fixed effects, we can identify the differential effect for beneficiary couples.

We report estimates of these models fin Table 8, columns 4-6. The estimate for 2006 implies that a one standard deviation increase in the municipality-level homicide rate (10.18 deaths per 100,000 individuals) is associated with a 1.6 percentage point decrease in spousal abuse (column 4, 0.01018×-1.615), and although the relationship is stronger for beneficiary women (column 6), this is not differentially so. Analogous estimates for survey year 2011 imply that a one standard deviation increase in the homicide rate is associated with a 0.5 percentage decrease in IPV, not statistically significant (0.01971×-0.303). We conclude that, although some evidence is suggestive of generalized social violence being associated with overall changes in IPV, and in particular for 2006, it does not explain the loosening of the relationship between program beneficiary status and spousal abuse.[24]

Improvement in Women's Labor Market Opportunities

An extensive literature documents that increases in a woman's relative labour opportunities, possibly by increasing her bargaining power within the household by means of an improvement in her outside option, can lead to lower levels of violence (e.g., Bowlus and Seitz, 2006; Aizer, 2010; Anderberg et al., 2016). If women's relative income-generating opportunities have improved in Mexico over the last decade, this may help explain the strong decline in the incidence of violence. To evaluate this hypothesis, we estimate models analogous to Aizer (2010)'s that capture the mitigating effects of the gender wage gap at the state level. The regression equation incorporating these effects is the following:

$$Y_{ivs} = \theta_1 T_{ivs} + \theta_2 T_{ivs} W_s + \beta_1 W_s + X_{ivs} \beta_2 + \varepsilon_{ivs} \tag{7}$$

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[24]These results are somewhat consistent with recent work presented in Tsaneva, Rockmore and Albohmood (2018) that finds that the effects of violent crime on female decision-making in Mexico are small and tend to be short-lived.

The W_s variable measures the female/male wage ratio in state s relative to the average gender wage gap for the sample of women and men in our study; the other variables are defined as above. The β_1 coefficient captures the partial correlation between the female/male wage ratio and spousal abuse rates among non-beneficiary couples, whereas the θ_2 term captures the differential correlation among beneficiary ones. We use the state-level rural wage gap measure because the surveys are representative at the state level and thus the lowest level of aggregation at which these measures can be consistently estimated is at this level.[25] Moreover, because the female/male wage ratio rate is measured at the state level, we do not include village or state fixed effects in this specification. In a second specification with state fixed effects, we can identify the differential effect for beneficiary couples.

We report estimates of these models in Table 8, columns 7-9. The estimate for 2006 implies that a 10 percentage point increase in the female/male wage ratio is associated with a decrease in spousal abuse by 0.2 percentage points (Panel A, column 7). The relationship for 2011 implies that an analogous increase in the later period is associated with a 2.1 percentage point decrease in spousal abuse (Panel B, column 7).[26] Neither of these estimates is statistically significant. Moreover, if we estimate a model that allows for a heterogeneous response by couples' beneficiary status, it implies that beneficiary women in states with lower female/wage ratios observe larger reductions in spousal abuse, relative to beneficiary women in states with higher female/male wage ratios. Though the estimates are not statistically significant for 2006 (column 9, Panel A), they are significant at the 90 percent level for 2011 (column 9, Panel B). The coefficients imply, on one end of the spectrum, that beneficiary women in legal reform states with the lowest female/male wage ratios (0.79), are 4.5 percentage points less likely to be victims of spousal abuse (0.79×0.319 - 0.264 - 0.035). On the other hand, beneficiary women in non-reform states with the highest

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[25]Using a state-level female wage gap measure may be somewhat restrictive for purposes of the analysis, as it may not appropriately capture the relative labor market opportunities women face across distinct municipalities and villages within the state. However, it should capture broad differences at the state level in these relative labor market opportunities.

[26]In 2006, the average wage gap for our sample was 24 percentage points, and by 2011, the average wage gap was substantially smaller, at 12 percentage points.

female/male wage ratios (0.98), are 4.5 percentage points more likely to be victims of spousal abuse (0.98×0.319 - 0.264). These heterogeneous effects imply a greater reduction in spousal abuse rates among beneficiary households in contexts of larger gender wage gaps. It also suggests that the complementarity between IPV divorce laws and the CCT program, though much weaker than in the short-run, tends to persist in the longer-run.

Figure 6: Oportunidades Program Maximum Monthly Benefits (in Real 1998 Mexican Pesos)
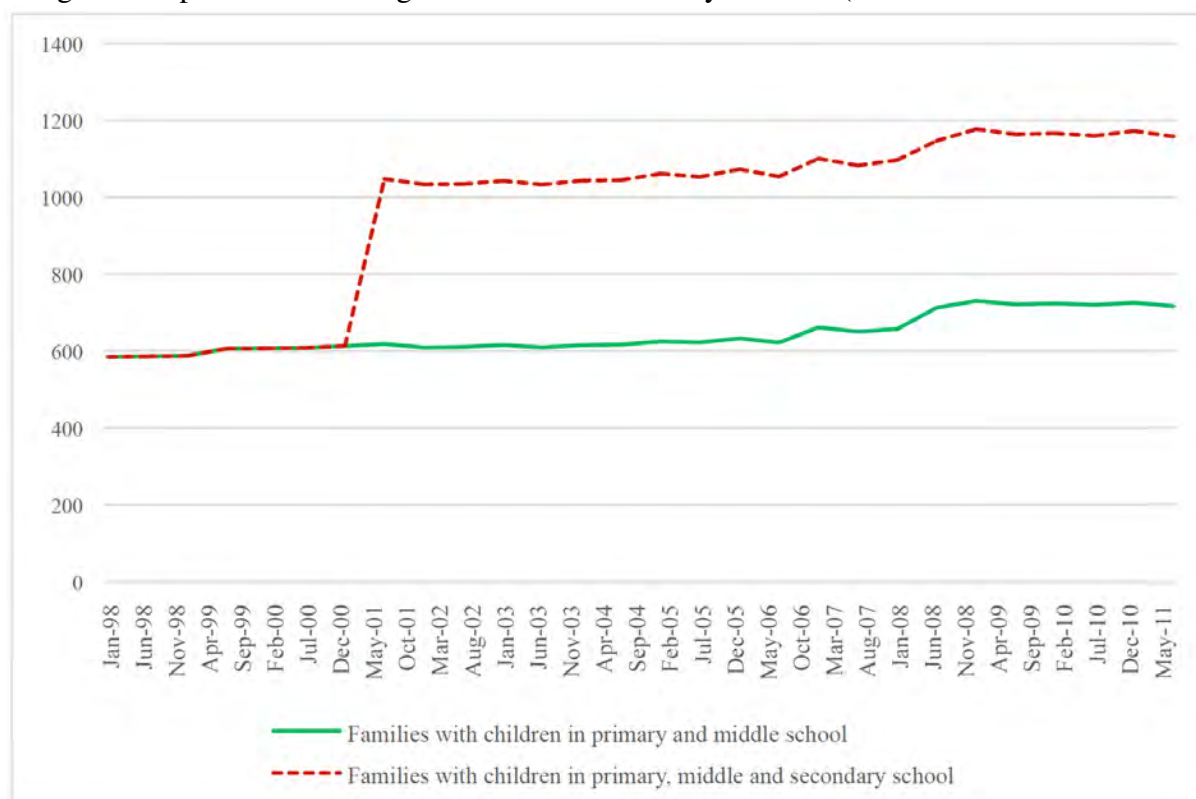
Source: Authors' calculations based on data from Secretariat of Social Development (SEDESOL) Mexico and the Bank of Mexico.

Table 8: Alternative explanations

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Dependent variable: Physical or sexual violence — Pseudo-panel								
Panel A: 2006									
Oportunidades beneficiary	0.033*	0.023	-0.071	0.017	0.019	0.026	0.010	0.001	-0.091
	(0.017)	(0.030)	(0.068)	(0.016)	(0.025)	(0.031)	(0.016)	(0.025)	(0.114)
Oportunidades beneficiary × IPV Divorce Law (2003)		0.020	0.019		-0.005	-0.004		0.017	-0.001
		(0.035)	(0.033)		(0.030)	(0.029)		(0.029)	(0.028)
Robustness variable	0.030	0.030	0.138	-1.615**	-1.618**	-1.069	-0.020	-0.021	
	(0.046)	(0.046)	(0.087)	(0.686)	(0.692)	(1.317)	(0.070)	(0.070)	
Oportunidades beneficiary × Robustness variable			-0.178			-0.769			0.137
			(0.117)			(1.792)			(0.130)
Total Effect (Op. beneficiary in IPV law state)		0.043**			0.014			0.019	
		(0.017)			(0.019)			(0.017)	
Panel B: 2011									
Oportunidades beneficiary	-0.003	0.013	0.004	-0.004	0.014	0.013	-0.007	-0.001	-0.264*
	(0.013)	(0.020)	(0.035)	(0.011)	(0.010)	(0.012)	(0.010)	(0.013)	(0.135)
Oportunidades beneficiary × IPV Divorce Law (2003)		-0.031	-0.031		-0.036**	-0.036**		-0.013	-0.035**
		(0.023)	(0.023)		(0.014)	(0.014)		(0.017)	(0.014)
Robustness variable	-0.023	-0.020	-0.008	-0.303	-0.285	-0.321	-0.213	-0.225	
	(0.036)	(0.035)	(0.054)	(0.573)	(0.564)	(0.522)	(0.200)	(0.204)	
Oportunidades beneficiary × Robustness variable			-0.019			0.048			0.319*
			(0.052)			(0.446)			(0.156)
Total Effect (Op. beneficiary in IPV law state)		-0.017			-0.022*			-0.013	
		(0.013)			(0.012)			(0.012)	
Fixed Effects level	Municipality			State			None		State
Individual & household controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Robustness variable	Village share beneficiary			Homicide rate (muni)			Female/male wage ratio (state)		
Observations	4035 / 5005			3571 / 4519			4240 / 5208		

Notes: Coefficient estimates from OLS regressions weighted by survey sampling weights. The 2003 sample includes couples with women ages 25 and older, with children aged 0-10, and who have been in union since 1997 or earlier. The pseudo-panel includes couples with women ages 28/33 and older, with children aged 3-13/8-18, and who have been in union since 1997 or earlier, respectively for the 2006/2011 surveys. The replication sample includes women ages 25 and older, with children aged 0-10, and who have been in union for at least six years. IPV Divorce Law (2003) is an indicator equal to 1 if the state had passed the IPV Divorce Law by 2003. Individual controls include indicator variables for woman and partner's age, woman and partner's indigenous status, women's schooling-level indicators, the partner's schooling attainment level, household size, cohabiting couple indicator, years in union, and variables measuring reported histories of spousal abuse in parental household during childhood. Robust standard errors in parentheses, clustered at the state level; significant at (*) 90 percent, (**) 95 percent, (***) 99 percent confidence levels.

Theoretical Framework

We present a model of spousal relations to highlight the mechanisms through which legal reforms and conditional cash transfers could affect divorce and conflict within the household. Our theoretical framework incorporates bargaining and conflict with incomplete information, together with endogenous divorce decisions, to formalize the effects that these two distinct sets of policies can have on the incidence of conflict and divorce.

As is common in the literature on intra-household bargaining, partners use their resources to generate a marital surplus and bargain over its allocation; who owns the resources in the household matters by affecting the spouses' outside options. Building on the work of Anderson and Genicot (2015) and Brassiolo (2016), in order for bargaining to fail some of the time we assume that spouses derive some private satisfaction with the marriage, whose magnitude is unknown to their partner. We further assume that when an offer is rejected, conflict ensues. This comes at a cost to each spouse, and a cost whose magnitude is realized only at the time of the conflict. The cost of conflict in our model can be thought of as the physical and or psychological pain that each spouse endures during an episode of violence. Its magnitude is uncertain ex-ante since it depends on many factors, including each spouse's ability to cope with it. Separation can only be achieved after going through a period of marital conflict; after observing her private cost of conflict, the wife may decide to end the relationship.

We model the legal reform as a reduction in the cost to divorce faced by the wife, as these policies allow women to unilaterally initiate divorce proceedings if they were victims of family violence. We also model the introduction of the CCT as an increase in the wife's income, consistent with the idea that these transfers were targeted towards the female partner. The model therefore allows us to think about the effects of these two policies and their interaction on the incidence of conflict and divorce. Most importantly, the model allows us to capture the selection mechanisms through which the policies may affect the types of couples that remain married and those that choose to divorce.

Preferences

We assume three possible states in a relationship: cooperation, conflict, and divorce. The utility functions of the partners depend on the status of their marriage.

Cooperation

The utility function in the cooperation state depends on the share of household resources each partner gets and their private gains to marriage. Specifically, we assume that preferences for the wife and husband are represented by the utility functions:

$$V^w(x, \theta_w; I_w) = u(x(I_h + I_w)) + \theta_w \ \& \ V^h(x, \theta_h; I_h) = u((1 - x)(I_h + I_w)) + \theta_h, \quad (8)$$

where $u(.)$ is a concave function representing the utility from consumption, i.e. $u'(.) \geq 0$ and $u"(.) \leq 0$, and θ_i, $i \in \{w, h\}$ represent a personal satisfaction from marriage. Partners enjoy consumption from their shared resources, where I_i, $i \in \{w, h\}$ is the individual income each partner contributes to the marriage and x captures the share of resources that are allocated to the wife's consumption . We follow Bloch and Rao (2002); Bobonis, González-Brenes and Castro (2013); Friedberg and Stern (2014) in assuming that each spouse's personal utility from marriage is private information; θ_i, $i \in \{w, h\}$ are independent and follow a distribution $\theta_w, \theta_h \sim G(.)$ with support $[\underline{\theta}, \bar{\theta}]$, such that $\underline{\theta} > 0$, i.e. both husband and wife enjoy positive utility from marriage.

Conflict

Households in the conflict state do not share resources; they enjoy utility of consumption only from their own income, and incur a cost of conflict κ_i, $i \in \{w, h\}$; this captures both a private cost of conflict as perceived by each partner as well as the intensity of the conflict. For instance, a more violent husband will impose a higher κ_w on his wife. Specifically, the utility functions in this state are represented by:

$$U^w(\theta_w, \kappa_w; I_w) = u(I_w) + \theta_w - \kappa_w \quad (9)$$

for the wife, and

$$U^h(\theta_h, \kappa_h; I_h) = u(I_h) + \theta_h - \kappa_h \tag{10}$$

for the husband. We assume $\kappa_w, \kappa_h \sim F(.)$ with support $[\underline{\kappa}, \bar{\kappa}]$, and $\underline{\kappa} \geq 0$, i.e. conflict will impose a loss in utility for both husband and wife.[27]

Divorce

Partners who separate enjoy only their own income for consumption:

$$U_D^w(I_w, \rho) = u(I_w) + \rho \tag{11}$$

for the wife, and

$$U_D^h(I_h) = u(I_h) \tag{12}$$

for the husband. In addition, the parameter ρ captures the additional utility or outside option from divorce for the wife such that $\rho < 0$ would represent a net loss from divorce. A reduction in divorce costs are represented by an increase in ρ.

Timing

Once married, each partner observes his/her level of private satisfaction from the marriage. The husband then proposes a division of household resources such that the wife gets share x (and the husband gets share $1 - x$) of total household income. The wife then chooses to accept or reject the husband's offer. If the wife accepts, they enter the cooperative state and enjoy V^w and V^h as their utility. If the wife rejects the offer, they enter the conflict state and learn their costs κ. The wife can then choose to divorce the husband. If she decides to stay married, they stay in the conflict state and get U^w and U^h, otherwise if they divorce, they get U_D^w and U_D^h. For simplicity, we assume

---

[27]One important distinction in our setup relative to that in Anderson and Genicot (2015) is that couples in the conflict state continue to enjoy the private gains to marriage.

husbands choose x, while wives choose whether to divorce or not.

Decisions

We use backward induction to present the decisions at each stage:

Wife's divorce decision: The wife chooses to divorce if $U^w < U^w_D$, or equivalently, the couple remains married if $\kappa_w \leq \theta_w - \rho$.

Partners' cooperation or conflict decision: If the wife rejects an offer, her expected utility is given by:

$$
\begin{aligned}
E_w(\theta_w; I_w, \rho) &= F(\theta_w - \rho)[u(I_w) + \theta_w] - \int_{-\infty}^{\theta_w - \rho} \kappa dF(\kappa) + [1 - F(\theta_w - \rho)][u(I_w) + \rho] \\
&= u(I_w) + \rho + F(\theta_w - \rho)(\theta_w - \rho) - \int_{-\infty}^{\theta_w - \rho} \kappa dF(\kappa)
\end{aligned}
\tag{13}
$$

The expected utility captures both possible states after rejecting an offer, conflict or divorce. A wife accepts the offer x if:

$$
V^w(x, \theta_w; I_w) \geq E_w(\theta_w; I_w, \rho)
\tag{14}
$$

Husband's offer decision: Let $\tilde{\theta}(x; I_w, \rho)$ be the value of θ_w such that $V^w(x, \theta_w; I_w) = E_w(\theta_w; I_w, \rho)$, then a wife accepts an offer if and only if $\theta_w \geq \tilde{\theta}(x; I_w, \rho)$, where $\tilde{\theta}(x; I_w, \rho)$ solves

$$
u(x(I_h + I_w)) + \theta_w = u(I_w) + \rho + F(\theta_w - \rho)(\theta_w - \rho) - \int_{-\infty}^{\theta_w - \rho} \kappa dF(\kappa)
\tag{15}
$$

Hence, $G[\tilde{\theta}(x; I_w, \rho)]$ is the probability that an offer x is rejected from the husband's perspective. The husband's expected utility if the wife rejects the offer is:

$$
E_h(\theta_h; I_h, \rho) = u(I_h) + [\theta_h - E(\kappa)] \frac{1}{G[\tilde{\theta}(x; I_w, \rho)]} \int_{-\infty}^{\tilde{\theta}(x; I_w, \rho)} F(\theta_w - \rho) dG(\theta_w)
\tag{16}
$$

Anticipating the potential rejection of the offer by the wife, the husband chooses an offer

$x^*(\theta_h; I_h, I_w, \rho)$ that maximizes his expected utility:

$$
\begin{aligned}
h(x, \theta_h; I_h, I_w, \rho) &= (1 - G[\tilde{\theta}(x; I_w, \rho)])V^h(x, \theta_h; I_h + I_w) + G[\tilde{\theta}(x; I_w, \rho)]E_h(\theta_h; I_h, \rho) \\
&= [1 - G(\tilde{\theta})][u((1-x)(I_h + I_w)) + \theta_h] + G(\tilde{\theta})u(I_h) \qquad (17) \\
&\quad + (\theta_h - E(\kappa)) \int_{-\infty}^{\tilde{\theta}} F(\theta_w - \rho)dG(\theta_w).
\end{aligned}
$$

Our main objects of interest are the ***divorce rate*** – the proportion of partners who choose to divorce – and the ***conflict rate*** – those who remain in the conflict state. The divorce rate (D) is given by the probability that a wife rejects an offer times the probability that she chooses to divorce in the case of rejection. Accordingly, it is given by:

$$
D(I_h, I_w, \rho) = \int_{-\infty}^{\infty} \int_{-\infty}^{\tilde{\theta}(x^*(\theta_h; I_h, I_w, \rho); I_w, \rho)} (1 - F[\theta_w - \rho])dG(\theta_w)dG(\theta_h) \qquad (18)
$$

The conflict rate is given by the probability that the wife rejects an offer conditional on not divorcing:

$$
C(I_h, I_w, \rho) = \frac{\int_{-\infty}^{\infty} \int_{-\infty}^{\tilde{\theta}(x^*(\theta_h; I_h, I_w, \rho); I_w, \rho)} F[\theta_w - \rho]dG(\theta_w)dG(\theta_h)}{1 - \int_{-\infty}^{\infty} \int_{-\infty}^{\tilde{\theta}(x^*(\theta_h; I_h, I_w, \rho); I_w, \rho)} (1 - F[\theta_w - \rho])dG(\theta_w)dG(\theta_h)}. \qquad (19)
$$

Predictions

We are interested in studying how changes in the two parameters of interest, a reduction in divorce costs due to the legal reform, ρ, and the CCT program (represented by an increase in the wife's income), affect the two outcomes of interest: the divorce rate (D) and the conflict rate (C). For simplicity, we assume that once an offer x^* has been made by the husband, it remains fixed. We summarize the main results here and present derivations and extensions in the theoretical appendix.

Result 1:

As divorce costs decrease, the equilibrium divorce rate increases ($\frac{\partial D}{\partial \rho} > 0$). There are two forces driving this result: a direct effect in which the decrease in ρ (i.e. an increase in the outside option)

makes couples who are in conflict more likely to choose to divorce; and an indirect effect by which the wife's expected utility from rejecting an offer becomes higher and hence it is more likely for wives to choose to reject an offer and enter the conflict or divorced state.

Result 2:

The change in the equilibrium conflict rate with respect to a decrease in divorce costs is ambiguous ($\frac{\partial C}{\partial \rho} <$ or > 0). *Women with sufficiently high costs of conflict* (κ_w) *or sufficiently low private gains from marriage* (θ_w) *are more likely to exit the relationship* ($\frac{\partial C}{\partial \rho} < 0$ if $\kappa_w > \bar{\kappa}_w$ or $\theta_w < \underline{\theta_w}$). Again, there are two main forces driving this result. On one hand, after the offer is rejected by the wife she is more likely to choose to divorce instead of remaining in the conflict state. This effect makes couples more likely to get divorced and less likely to be in conflict. On the other hand, given a higher expected utility in the conflict state ($E_w(\theta_w; I_w, \rho)$), the more likely the wife is to reject her husband's offer, such that the proportion of households in the cooperative state decreases. The overall magnitude of the effect depends on the distribution of the wives' private gains from marriage (θ_w).

The types of couple remaining in the conflict state changes as ρ increases. In particular, women with high costs of conflict κ_w and lower private gains from marriage θ_w are more likely to exit the relationship (as highlighted by the direct effect) relative to women with lower κ_w and higher θ_w. If conflict manifests itself in different forms across different levels of κ, then this particular selection mechanism, through which different types of couples remain in marriage after changes in divorce laws, is an important channel through which the potential effects arise.

Result 3:

The effect of an increase in the wife's income on the equilibrium divorce rate depends on the wife's ex ante outside option and her partner's private gains to marriage. Specifically, there exist thresholds ρ_H, ρ_L and $\hat{\theta}$, such that:

$$\frac{\partial D}{\partial I_w} \begin{cases} > 0 & \text{if} \quad \rho > \rho_H \\ \\ < 0 & \text{if} \quad \rho < \rho_L \end{cases} \tag{20}$$

How the wife's income affects divorce rates depends on the ex-ante allocation of resources x^* proposed by the husband and the utility function from consumption u. Two important determinants of x^* are the wife's initial outside option, which depends on ρ, and the husband's private utility from the marriage θ_h. Women facing very low divorce costs ($\rho > \rho_H$) are compensated by husbands through a high share x^* of household resources. However, since utility is diminishing in consumption, an increase in I_w increases the utility of rejecting the husband's offer relative to the utility of staying together, such that $\frac{\partial D}{\partial I_w} > 0$. On the other hand, if women face very high divorce costs ($\rho < \rho_L$), the reverse is true.

For, $\rho_L < \rho < \rho_H$, changes in the divorce rate will depend on the distribution of types F and G. Husbands who enjoy a high private utility from marriage θ_h offer favourable allocations x^* to their wives such that divorce is less likely for these couples, and vice versa. Couples in which the husband enjoys greater benefit from the marriage are more likely to remain together after the wife receives the cash transfer. If husbands with higher gains from marriage are also less likely to perpetrate violence, then this highlights one important selection mechanism by which CCTs can affect the composition of married couples.

Result 4:

The effect of an increase in the wife's income on the equilibrium conflict rate is ambiguous. The change in the conflict rate with respect to the income of the wife $\frac{\partial C}{\partial I_w}$ is more negative for women with relatively higher costs of conflict κ_w and for women with relatively lower utility of marriage θ_w.

Even though the overall effects of a CCT for women is ambiguous for both conflict and divorce, this selection mechanism suggests that different types of women will be in each of these states. As income increases, women become more likely to reject the offers from their husbands.

Once women are in the conflict state, those facing high costs of conflict choose to divorce, while those facing relatively lower costs remain in conflict.

The results here follow from the same analysis as in the previous point. Note in particular that the wife's income does not shift her relative preferences between the conflict and divorce states. However, changes in the wife's income affect the *types* of couples that remain in marriage, as given by their utility from marriage θ and their costs of conflict κ.

With respect to the differential effect of female income for conflict among women with relatively higher costs of conflict, imagine two sets of women H ($\kappa_w \geq \tilde{\kappa}$) and L ($\kappa_w < \tilde{\kappa}$) facing different costs of conflict stratified by the threshold level $\tilde{\kappa} = \theta_w - \rho$. The conflict rate in the high cost group decreases or increases at a slower rate than that for women in the low costs group, as income changes. That is, $\frac{\partial C_H}{\partial I_w} < \frac{\partial C_L}{\partial I_w}$. Note that the inverse is true for divorce, that is, $\frac{\partial D_H}{\partial I_w} > \frac{\partial D_L}{\partial I_w}$.

The change in the conflict rate with respect to the income of the wife $\frac{\partial C}{\partial I_w}$ is smaller. As before, with two sets of women H ($\theta_w \geq \tilde{\theta}$) and L ($\theta_w < \tilde{\theta}$) facing different utility of marriage divided by $\tilde{\theta} < \kappa_w + \rho$, then conflict in the low utility group decreases or increases at a slower rate than that for women in the high utility group, as income changes. That is, $\frac{\partial C_L}{\partial I_w} < \frac{\partial C_H}{\partial I_w}$. Note that the inverse is true for divorce, that is, $\frac{\partial D_L}{\partial I_w} > \frac{\partial D_H}{\partial I_w}$. As before, this selection mechanism suggests that different types of women will be in each of these states. Once women are in the conflict state, those facing low utility from marriage choose to divorce, while those facing relatively higher utility choose to remain in the conflict state.

This simple model of household relationships outlines the mechanisms through which the policies of interest may affect conflict and divorce. The model highlights some of the tradeoffs faced and suggests that both legal reforms and CCTs lead to important selection mechanisms which affect the *types* of couples which remain in union. In particular, couples in which the partners enjoy greater utility from marriage θ and in which wives face lower costs from conflict κ_w are more likely to stay together following the introduction of these policies.