"Till death do us part": Easier Access to Divorce and Femicide*

Francesca Agosti† Davide Cipullo‡ Elisa Muscarella§

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Abstract

This paper estimates the causal effect of easing access to divorce on domestic violence. We study two legislative reforms in Italy that sharply cut the monetary costs and mandatory waiting periods for divorce, triggering a large and heterogeneous rise in divorce rates nationwide. We combine administrative data on the universe of divorces with a novel dataset on femicides, and we exploit plausibly exogenous variation in municipalities' exposure to the reforms, arising from municipalities' age composition interacted with national age-specific divorce rates. We find that easier access to divorce increases the probability of femicide, an effect primarily driven by domestic cases. We also find a reduction in measures typically associated with non-lethal domestic violence, indicating that lethal and non-lethal violence need not respond in the same way to an improvement in women's outside options.

Keywords: Divorce, femicide, intimate partner violence, intrahousehold bargaining

JEL: J12, D13, J16, K42

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†Università Cattolica del Sacro Cuore, INSPIRE-LEAP (e-mail: francesca.agosti@unicatt.it).

‡Università Cattolica del Sacro Cuore, CESifo, CIFREL, Uppsala Centre for Fiscal Studies (e-mail: davide.cipullo@unicatt.it).

§Universitat de Barcelona, IEB, IPERC, INSPIRE-LEAP (e-mail: elisa.muscarella@ub.edu).

1 Introduction

Violence against women (VAW) affects millions of women worldwide and poses substantial challenges for public policy. Globally, nearly 1 in 3 women are estimated to have experienced physical or sexual violence by an intimate partner, non-partner sexual violence, or both, over their lifetime (World Health Organization, 2021). Its consequences are severe for victims and exposed children alike: adverse health outcomes (Aizer, 2011; Campbell, 2002), lost productivity (Duvvury et al., 2023), and intergenerational transmission of violence (Carrell and Hoekstra, 2010; Gutierrez and Molina, 2021; Lloyd, 1997; Pollak, 2004). Beyond these individual costs, VAW imposes a substantial economic burden, estimated at around 5% of global GDP through lost productivity, healthcare, and justice system costs (EIGE, 2023; Fearon and Hoeffler, 2014; Peterson et al., 2018; Shah and Barski, 2025).

Among the many forms VAW can take, femicide – the killing of a woman for gender-related reasons – represents its most extreme and irreversible outcome. Globally, an estimated 50,000 women and girls were killed by an intimate partner or family member in 2024 alone, around 60% of all female homicides worldwide and equivalent to one woman killed roughly every ten minutes (UN Women and UNODC, 2025). The intimate partner is by far the most common perpetrator within this domestic sphere: in Italy, 106 femicides were recorded in 2024 out of 116 homicides with a female victim, 93% of them committed by a partner, ex-partner, or relative (ISTAT, 2025b). Despite the severity of the problem, femicide became an autonomous criminal offense under Italian law only in 2025, spurred by media attention on high-profile cases and public mobilization, punishable by life imprisonment when motivated by gender-based hatred, discrimination, or control.

Since the couple is the setting in which most femicides take place, institutional changes that reshape women's outside options within relationships – such as divorce law reforms approved by many countries in recent decades (Brassiolo, 2016; Calabresi, 2024; Corradini and Buccione, 2023; Friedberg, 1998; García-Ramos, 2021; Wolfers, 2006) – can affect violence through two distinct but potentially simultaneous channels (Chiappori et al., 2002). First, easier divorce strengthens a woman's outside option and makes the threat of separation more credible, changing the constraints faced by husbands who wish to preserve the relationship. This can discipline violence, although some husbands may instead respond by intensifying control when sufficiently high violence can deter separation. Second, separation is itself a period of acute risk: spousal homicide risk is substantially higher among estranged than cohabiting couples (Anderson and Genicot, 2015; Campbell et al., 2003; Einiö et al., 2023; Molina et al., 2025; Wilson and Daly, 1993, 1992). These two channels are not mutually exclusive: easier divorce changes violence within continuing relationships while also inducing some couples to enter the separation process and become exposed to its associated risks. Reconciling these two forces is a central goal of this paper.

This paper studies the effect of divorce institutions on intimate partner violence in Italy, focusing on its most extreme and irreversible outcome: femicide. We study two legislative reforms that substantially eased marital dissolution. The first, enacted in 2014, introduced out-of-court procedures for separation and divorce, including assisted negotiation between lawyers and agreements filed directly before the civil registrar. The second, enacted in 2015, shortened the mandatory separation period preceding divorce from three years to six months in consensual separations and twelve months in contested ones. By reducing the procedural costs and frictions associated with marital dissolution, these reforms generated a sharp increase in divorce filings: the number of new divorces rose from about 52,000 in 2014 to roughly 82,000 in 2015 (ISTAT, 2022).

We exploit cross-municipality variation in exposure to these reforms by constructing a measure of reform intensity that is based on a shift-share intuition. Age plays a central role in this construction: because the reforms lowered the cost of divorcing, we expect their impact to be strongest among age groups in which the divorce is more likely to occur, particularly middle-aged individuals. More specifically, we combine variation across municipalities in the pre-reform population age structure with aggregate differences across age groups in the likelihood of divorcing.

Our empirical analysis relies on a novel dataset that covers all femicides committed in Italy between 2007 and 2021. We complement this dataset with administrative records on all divorce applications and procedures concluded in Italy. We find that the reforms significantly increased the probability of observing a femicide in municipalities more exposed to them. The effect is concentrated among cases involving ex-partners and relatives, consistent with relationship dissolution representing a particularly vulnerable phase, and is robust to alternative outcome definitions and to using administrative mortality records instead of our manually compiled dataset.

To shed light on the mechanisms, we explore several dimensions of heterogeneity. We find that the effect is stronger in municipalities with more favorable attitudes toward divorce and higher average income, as well as in municipalities where pre-reform domestic violence and homicide rates were higher.

Finally, we investigate whether the reforms also affected reported offenses commonly associated with non-lethal domestic violence. Because these administrative crime measures are available only at the province level, we reconstruct our shift-share measure of exposure by combining national age-specific divorce rates with pre-reform differences in provincial age composition. We find that greater exposure to the reforms reduces reported threats, offenses, and a composite index of crimes commonly associated with domestic violence, while the positive effect on femicides remains statistically significant when the analysis is aggregated to the province level. Because beatings, threats, and offenses are not disaggregated by the victim's sex, we interpret this evidence as complementary rather than as a direct measure of IPV against women. The resulting divergence between lethal and non-lethal violence motivates the conceptual framework described in our first contribution below.

Our paper contributes primarily to the literature on women's outside options and intimate partner violence. A large economic literature shows that improvements in women's outside options can reduce violence by strengthening their bargaining position within the household (Aizer, 2010; Angelucci and Heath, 2020; Calabresi and Rodríguez-Planas, 2025; Haushofer et al., 2019; Hidrobo and Fernald, 2013; Perova et al., 2021), while another strand documents retaliatory or backlash responses when women's economic or legal position improves (Bergvall, 2024; Bhalotra, Brito et al., 2021; Bhalotra, Kambhampati et al., 2021; González and Rodríguez-Planas, 2020; Guarnieri and Rainer, 2021; Heath, 2014). Related work on divorce institutions likewise finds both reductions and increases in intimate partner violence following reforms that strengthen the option to leave a marriage (Brassiolo, 2016; Calabresi, 2024; Corradini and Buccione, 2023; Dee, 2003; García-Ramos, 2021; Silverio-Murillo, 2019; Stevenson and Wolfers, 2006). Our contribution is to show that these apparently conflicting responses need not characterize different institutional environments: the same policy can generate both at the same time because it may change behavior at different stages of the relationship.

Our conceptual framework formalizes this idea by distinguishing couples according to how lower divorce costs change their equilibrium relationship status. Among couples that remain together under both regimes, a more credible threat of exit can discipline violence, although the model also allows sufficiently coercive husbands to respond by increasing violence to deter separation. Among couples that enter the separation process under both regimes, lower divorce costs induce de-escalation: equilibrium non-lethal violence falls and peaceful dissolution becomes more likely, reducing femicide risk. Crucially, however, lower divorce costs also induce additional couples to switch from staying together to attempting separation. These couples move from zero equilibrium femicide risk to a strictly positive risk associated with the separation stage, while their non-lethal violence may either rise or fall depending on the husband's type. This framework rationalizes our central empirical finding that the same improvement in women's outside options can be associated with less non-lethal violence but more femicide.

Second, we contribute to the literature on lethal violence against women by showing that lethal and non-lethal violence need not represent different points along a common severity continuum. A key distinction between the two forms of violence concerns the role that violence can play within a relationship. Non-lethal violence can be instrumental: it can be used to extract resources, enforce compliance, or maintain control over a partner (Aizer and Dal Bó, 2009; Bloch and Rao, 2002; Bobonis et al., 2013). It can also take the form of coercive control, whereby violence or the threat of violence restricts the partner's behavior and ability to leave. Lethal violence is fundamentally different. By definition, killing the partner cannot serve as an instrument for shaping her subsequent behavior within the relationship. Instead, lethal violence can arise when the relationship itself is breaking down and the perpetrator's ability to exercise control is being lost. Consistent with this distinction, evidence from criminology documents that lethal violence is disproportionately concentrated around relationship dissolution (Campbell et al., 2003; Monckton Smith, 2020; Spencer and Stith, 2020; Wilson and Daly, 1993, 1992). Our paper brings this difference into a common economic framework and empirical design. The same policy that strengthens women's outside options can reduce the scope for instrumental or coercive violence within ongoing relationships while increasing exposure to lethal violence by inducing additional couples to enter the separation stage. Empirically, using the same reforms and the same source of identifying variation, we find an increase in femicide alongside a decline in reported offenses commonly associated with non-lethal domestic violence.[1]

Third, our paper contributes to the literature on the consequences of divorce law. Family economics has long studied how divorce institutions affect marital dissolution (Friedberg, 1998; Wolfers, 2006), intra-household bargaining and intertemporal behavior (Chiappori et al., 2002; Stevenson, 2008; Voena, 2015), and children's and family outcomes (Gruber, 2004). Most closely related to our paper is the literature on divorce law and intimate partner violence. In the United States and Spain, easier divorce is associated with declines in domestic violence, female suicide, and spousal homicide (Brassiolo, 2016; Dee, 2003; Stevenson and Wolfers, 2006), and a subsequent Spanish reform strengthening fathers' custody rights similarly reduced IPV by shifting intra-household bargaining positions (Fernández-Kranz et al., 2026). In Egypt, unilateral no-fault divorce reduced domestic abuse (Corradini and Buccione, 2023). Evidence from Mexico instead points to increases in IPV following unilateral divorce legalization (Calabresi, 2024; García-Ramos, 2021; Silverio-Murillo, 2019), while Hoehn-Velasco and Silverio-Murillo (2020) find no significant effects on female suicide and homicide. Our results provide a way to reconcile these findings: easier divorce changes both bargaining within ongoing relationships and the extensive margin of couples entering separation, so bargaining and backlash forces can

[1] A small literature in economics studies femicide directly but falls short in conceptualizing its differences with respect to non-lethal forms of violence. Gutiérrez-Romero (2024) examines divorce liberalization and femicide in Mexico; Asik and Ozen (2021) studies female homicides during COVID-19 restrictions in Turkey; and Frisancho et al. (2022) studies female political representation and femicide in the United States. In Italy, recent work studies the provision of anti-violence centers (Cerqua et al., 2026), local fiscal responses to femicides (Buonanno et al., 2025), and their electoral consequences (Buonanno et al., 2026). Closest to our setting, Pisanelli (2025) compares the national incidence of femicides before and after one of the reforms we study and documents an increase in femicides together with a reduction in IPV help-seeking. Relative to this literature, our contribution is to provide a subnational causal design and to connect lethal violence to the broader economic mechanism through which outside options reshape both behavior within ongoing relationships and exposure to the separation stage.

operate simultaneously within the same institutional setting and on different margins of violence.

The remainder of the paper is organized as follows. Section 2 describes the institutional background, detailing the evolution of divorce legislation and the legal framework surrounding femicide in Italy. Section 3 introduces the data sources used in the analysis. Section 4 presents the empirical design. Section 5 reports the main findings, while Section 6 provides a series of robustness checks. Section 7 explores heterogeneous treatment effects across different institutional and sociodemographic dimensions. Section 8 examines the effect of the reform on non-lethal domestic violence. Section 9 presents a conceptual framework that organizes our empirical findings. Lastly, Section 10 concludes.

2 Italian institutional setting

2.1 The evolution of Divorce Law in Italy

The right to legal divorce was first introduced in Italy in 1970, amid significant debates in the society.[2] Distinctive features of the Italian legal framework are that divorce must be preceded by a period of formal separation (henceforth, legal separation), and that it can follow either a *consensual* or a *judicial* procedure depending on whether the spouses agree on the terms.[3] Under a consensual divorce, spouses submit an agreement on the division of jointly owned assets, child custody, and any monthly allowance to a court, which approves it if it protects the rights of the children and complies with legal principles. Under a judicial divorce, a judge determines these terms. The choice between the two procedures does not depend on whether the decision to file was unilateral or joint: unilateral requests are admissible and may follow either route. Both procedures require couples to appear before a court (typically assisted by a lawyer), implying that legal separation and divorce come with a significant financial, administrative, and time burden. Under the original 1970 law, the minimum mandatory length of legal separation before a divorce was five years.[4] This requirement was reduced to three years in 1987.

In this paper, we focus on two reforms to divorce law approved between 2014 and 2015, in close succession and as part of the same agenda aimed at making divorce more accessible. The reform agenda was motivated both by the desire to align Italian family law with other European

[2] The introduction of divorce sparked intense public and political debates. The 1974 referendum resulted in a clear majority in favor of maintaining the legislation, thereby consolidating the legal legitimacy of divorce in Italy.

[3] The legal separation suspends marital obligations (e.g., cohabitation and fidelity) but does not terminate the marriage: for example, unless the spouses later pursue the divorce procedure, the surviving spouse retains full inheritance rights.

[4] In specific circumstances, such as separation resulting from the exclusive fault of one spouse and a lack of mutual agreement, this waiting period could be extended up to seven years.

countries and by the recognition that the length and inefficiency of the existing procedure induced many couples to remain legally separated indefinitely, rather than proceeding to divorce, unless one of them intended to eventually get married again with a new spouse. Indeed, as we document in Panel (a) of Figure 1 below, prior to the reforms the average time elapsed between legal separation and the divorce application exceeded 70 months, well above the 36 months that were formally required.

The first reform was Law No. 162/2014 (in force from December 2014), which allowed spouses without minor or dependent children to submit their consensual separation or divorce agreement before a civil servant (*Ufficiale di Stato Civile*) of the municipality where the spouses live or of the municipality in which they got married instead of being forced to go before a court. Thus, the reform waived the obligation to appear in court and to be assisted by a lawyer, though spouses retained the right to hire a lawyer if they wished so.

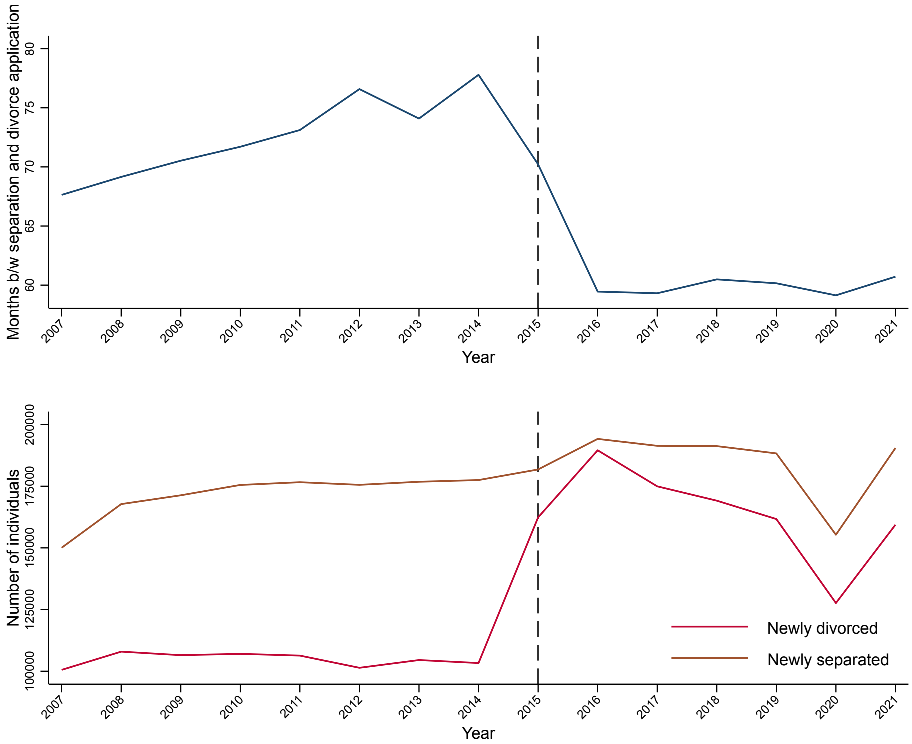
The second reform is Law No. 55/2015, known in the Italian public debate as the "short divorce law" (*Legge sul divorzio breve*), effective from May 26, 2015. The reform reduced the minimum mandatory length of legal separation from three years to six months for consensual separation and twelve months for judicial separation.

Taken together, the two reforms substantially reduced the financial and time costs of obtaining a divorce. As a result, as shown in Figure 1, the average number of months between separation and divorce request declines sharply in 2015–2016 (panel (a)) and the number of individuals obtaining a divorce nearly doubled within two years since the approval.[5] This pattern shows that the increase in divorces is not merely the mechanical effect of shortening the mandatory separation period: absent a behavioral response, the reform would have produced a short-lived peak followed by an equally large decline. Instead, we interpret the observed pattern as reflecting both a stock effect – previously separated couples accelerating divorce filings once procedural constraints were relaxed – and a flow effect, whereby lower legal and time costs induced additional couples to initiate separation and divorce. The behavior of legal separations, also reported in Panel (b), supports this interpretation. Separations were already far more numerous than divorces before the reform; yet they also increased after 2015. This indicates that also the propensity to separate increased, which corroborates the hypothesis that the jump in divorces discussed above reflects a behavioral response rather than a mere mechanical effect.[6]

[5] Compared to a pre-reform stable trend (100,000 newly divorced individuals per year), the reforms lead to an immediate increase to 165,000 in 2015 and 200,000 in 2016. The number of divorces declined somewhat in subsequent years but remained well above pre-reform levels. The number of individuals who obtained a divorce during 2020 and 2021 is also above the pre-reform trend despite the COVID-19 lockdown in Italy, which suspended or slowed down administrative procedures for many months during both years.

[6] Figure A.1 in the Appendix documents that the reforms changed the composition of divorces too. Before 2014, consensual procedures outnumbered judicial procedures although both required a court hearing. After the reforms, the number of divorces increased markedly across procedures: the increase in the number of consensual divorces

Figure 1. Trends in divorce



Notes: Panel (a) plots the average duration (in months) between legal separation and the divorce application in Italy, 2007–2021. Panel (b) plots the number of newly divorced and newly separated individuals over the same period, based on official ISTAT data. The vertical dashed line marks 2015, the year in which both divorce reforms were fully in effect.

2.2 Femicide in Italy: definition, measurement, and legal framework

Violence against women (VAW) and intimate partner violence (IPV) are widespread phenomena in Italy: survey-based data indicate that 31.9% of Italian women aged 16–75 have experienced at least one episode of physical or sexual violence during their adult life (ISTAT, 2025a), a share broadly in line with that reported for other Western countries (European Institute for Gender Equality (EIGE), 2024). Violence occurs both within and outside intimate relationships: 12.6% of women report having been abused by a current or former partner, while 26.5% report violence by a relative, friend, colleague, acquaintance, or stranger.[7] Administrative data corroborate this picture: in 2023, the Department of Public Security registered almost 14,000 requests for help and police intervention related to episodes of domestic or gender-based violence against

can be explained by the compound effect of shorter waiting times and the new extrajudicial channel because its rise is driven mainly by cases handled by a municipal officer, while the increase in the number of judicial divorces is perhaps helpful to capture the effect of shortening the legal separation window.

[7] This figure is based on preliminary results from ISTAT's ongoing survey *Sicurezza delle donne* (2023–2026).

women. This figure likely underestimates the true underlying scale of the issue, as suggested by the number of calls to the national public-utility helpline against violence and stalking (*1522*) – which have been steadily growing over time, reaching 65,000 in 2024 alone – and by the share of women who call seeking help but are not willing to file a formal complaint (72%). Both sources indicate that the problem is concentrated primarily within the domestic sphere: 50% of victims identify a current partner (cohabiting or not) as the perpetrator, 21% a former partner, and 11% another family member. The support infrastructure has grown over time, reaching 409 anti-violence centers in 2024, which together assisted over 26,000 women; yet Italy remains far from the standard of one center per 50,000 women set out in the Istanbul Convention.

The most extreme manifestation of violence against women is femicide, defined by the United Nations as "the killing of women and girls because of their gender" (United Nations, 2012). According to the most recent measurement guidelines[8], this definition covers killings by (i) current or former intimate partners, (ii) other family members, and (iii) other perpetrators when accompanied by gender-related motives or circumstances (UNODC and UN Women, 2022).[9] Most recent Italian data confirm that the large majority of female homicides fit the femicide definition: ISTAT estimates that 106 of the 116 intentional killings of women recorded in 2024 qualify as femicides (91.4%), so that only a small residual share of cases falls outside the definition. Of these 106 cases, 62 were committed by a current or former partner (58.5%), 37 by another family member (34.9%), and the remaining cases by another perpetrator with an explicit gender-related motive (6.6%) (ISTAT, 2025b). Figure A.2 shows that this is not a feature of the most recent year alone: killings perpetrated by a current or former partner or another family member have constituted the main component of female homicide throughout the 2002–2022 period, and have not declined at the same pace as total female homicides. Because this measure excludes gender-motivated killings by other perpetrators, it represents a lower bound on the true extent of femicide in Italy.

The topic has gained increasing prominence in public opinion and in both the legislative process and the policy agenda in Italy. Media coverage of the phenomenon has increased markedly over time. According to Dow Jones Factiva, the term "femicide" appeared in only 37 articles published in Italy in 2007, whereas by 2021 this figure had risen to almost ten thousand articles, a number that doubled again by 2025. From a legislative standpoint, this growing attention has been accompanied over time by a series of interventions, including the establishment of the 1522 helpline and the introduction of specific criminal offenses and aggravating

[8]In 2022, the UN Statistical Commission approved the *Statistical Framework for Measuring the Gender-Related Killing of Women and Girls (also referred to as femicide/feminicide)*, developed by the United Nations Office on Drugs and Crime (UNODC) and UN Women.

[9]The third category includes, among others, cases involving prior abuse, exploitation or trafficking, abduction, sexual violence, gendered forms of post-mortem violence, or explicit gender-based hate motives.

circumstances,[10] and, most recently, into the introduction of femicide as a stand-alone criminal offense under Law No. 181/2025, punishable by life imprisonment when the killing is motivated by the victim's gender[11] – a significant institutional advancement after decades in which the legal framework lagged behind the structural dynamics of gender-based violence.

2.3 Statistical and administrative sub-national levels of government

As will become apparent in the next sections, our empirical analysis leverages territorial variation in exposure to the two divorce law reforms. For this reason, it is essential to provide the reader with the necessary details about subnational statistical units of observation and administrative layers. Italy has three sub-national levels of government. From the bottom to the top, these are the municipalities ($\approx$ 8,000, with non-negligible variation in their count over time due to splits and mergers), the provinces (107), and the regions (20). Municipalities are nested into provinces which in turn are nested into regions. In the remainder of the paper, we will use subscript r when referring to a region, subscript p when referring to a province, and subscript m when referring to a municipality, respectively. When no subscript is used, the variable refers to the national level.

3 Data

3.1 Data: Divorces

We utilize restricted-access individual-level administrative data collected by the Italian National Institute of Statistics (ISTAT) from courts and municipalities, accessed through the ADELE laboratory (*Scioglimenti e cessazioni degli effetti civili del matrimonio*). The data report information from virtually the universe of divorces obtained in Italy from 1990 to 2021.[12] For the

[10] The Italian legislative response to gender-based violence has evolved gradually and often in a fragmented manner. A first turning point occurred in 2013 with the ratification of the Istanbul Convention through Law No. 77/2013, the first legally binding international instrument establishing a comprehensive framework to prevent and combat violence against women and domestic violence. In the same year, the so-called "femicide decree" (Decree-Law No. 93/2013, later converted into Law No. 119/2013) introduced significant amendments to the Criminal Code and Criminal Procedure Code, strengthening aggravating circumstances for domestic abuse, sexual violence, stalking, and the misuse of digital tools. It was at that stage that the term *femminicidio* appeared for the first time in Italian legislation, albeit only as an aggravating circumstance rather than as a distinct criminal offense. Subsequent reforms, including Law No. 69/2019 (the so-called *Codice Rosso*), further reinforced procedural safeguards for victims and increased penalties for related crimes.

[11] i.e., hatred, discrimination, control, or domination toward the victim because she is a woman; refusal of an affective relationship; or limitation of her personal freedom.

[12] According to our conversations with ISTAT staff, some courts or municipalities did not respond to ISTAT's data requests in some of the more recent years; in those cases, ISTAT reconstructed the universe with oversampling techniques.

purposes of this analysis, we restrict the time period to divorces occurring between 2007 and 2021.[13] The microdata include detailed sociodemographic information on both spouses (age, gender, residence, nationality, education, occupation), characteristics of the marriage (place of celebration, presence of children, marital property regime), and information on the divorce settlement (child custody, alimony).

Because microdata accessed through the ADELE laboratory cannot be extracted from the secure research environment and combined with external data sources that are not publicly available (as it is the case for our hand-collected femicide data, as it will become apparent in the next sub-section), we requested a number of customized extractions of municipality-level aggregates. To construct the variable used in the empirical analysis – the number of newly divorced individuals in each year as a percentage of the municipal population – we divide these municipal counts by the corresponding municipal population in each year, obtained from ISTAT. According to these extractions, the share of newly divorced individuals increased from approximately 0.13% in the pre-reform period to 0.24% after the reform.

3.2 Data: Femicides

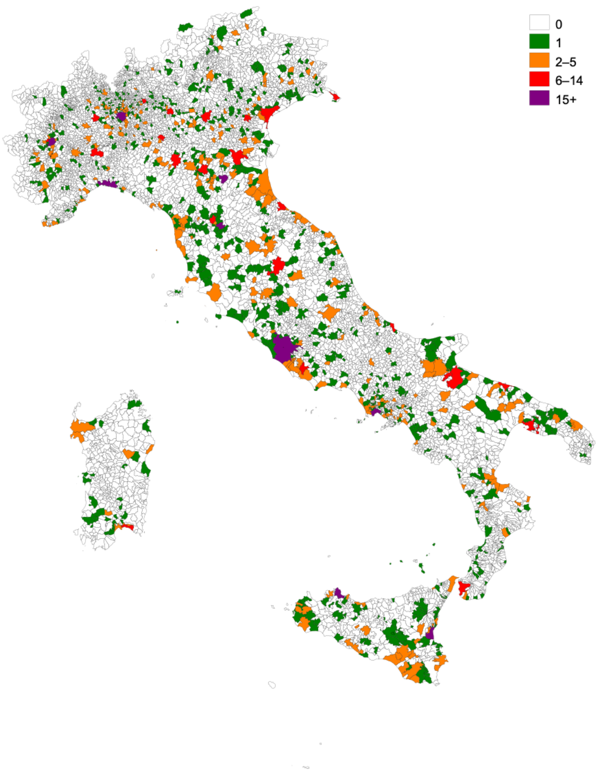
We constructed a novel dataset covering virtually the universe of femicides that occurred in Italy since 2007 by systematically collecting and cross-validating information from multiple independent sources. The primary source is *La Casa delle Donne per non Subire Violenza*, a non-governmental organization based in Bologna and Reggio Emilia that has monitored femicide events nationwide since 2005. This is the only source currently used in the literature on femicides in Italy (Buonanno et al., 2026, 2025; Cerqua et al., 2026; Colagrossi et al., 2023; Pisanelli, 2025), owing to its comprehensive coverage and its unique ability to provide consistent data going back further in time than any alternative source. We complement it with two additional independent sources: *La 27esima Ora*, an online portal published by the most prominent Italian newspaper *Corriere della Sera* (coverage starts in 2012), and with the independent website *femminicidioitalia.info* (systematic coverage since 2017 plus information about the most salient femicides occurred earlier). We complement these three sources with additional information retrieved from other national and local newspapers. As a result, our data provides more comprehensive coverage of femicide cases in Italy than the previous literature. For example, Cerqua et al. (2026) (who primarily relies on *La Casa delle Donne per non Subire Violenza*)

[13]This choice is motivated by the enactment of Law No. 54/2006, which reformed child custody arrangements in cases of separation and thus may have affected family dynamics and divorce behavior. To avoid confounding the effects of the divorce reforms we study with any effects stemming from the 2006 custody reform, we exclude earlier years from the analysis.

covers 1,730 cases between 2007 and 2021, as compared to our coverage of 1,881 femicides.[14]

We identify a total of 1,881 femicide cases, corresponding to an average of approximately 126 cases per year. Figure 2 provides an overview of the spatial distribution of these cases across Italy. The map displays the municipalities in which at least one femicide occurred over the period (around 13% of total municipalities), with different colours indicating the number of cases recorded in each. The map reveals a concentration of cases in large urban areas, alongside a broad geographic dispersion across municipalities of various sizes and regions. The large majority of affected municipalities — approximately 70% — record only a single case over the entire period; these single-case municipalities alone account for roughly 9% of all Italian municipalities.

Figure 2. Geographic distribution of femicides



Notes: The figure plots the total number of femicides recorded in each Italian municipality over the period 2007–2021. Municipalities are grouped into five bins based on their cumulative femicide count, as shown in the legend.

For each femicide, we recover detailed information on both victims and perpetrators, including age, nationality, residence, occupation, and any history of mental health issues or substance abuse. Additionally, we obtain data on the relationship between victim and perpetrator, the

[14]Notice that Cerqua et al. (2026) identified four of their 1,730 cases using other newspapers.

presence of children, and prior incidents of violence, as well as case-specific details such as the date and location of the crime, the weapon used, and the likely motive. Table B.1 provides descriptive evidence on these characteristics. As expected, in 45.9% of cases the perpetrator is the victim's current partner, and in an additional 12.6% an ex-partner, implying that nearly 60% of femicides are committed within an intimate relationship. Killings perpetrated by relatives of the victim or of her partner/ex-partner account for 22.2% of cases, while cases involving professional relationships, stalking, or situations in which the perpetrator does not fall into the previous categories or remains unidentified are comparatively rare. In sum, more than 80% of femicides occur within the family sphere (in what follows, we refer to these cases as the domestic femicides).

Understanding the circumstances that trigger a femicide is informative for interpreting our results and for relating them within the existing literature. We classify femicide cases into categories based on the precipitating motive.[15] Figure A.3 reports this classification separately for domestic and non-domestic femicides. Among domestic cases, femicides triggered by separation or by the victim's decision to end the relationship account for nearly 30% of cases, making it the single most common precipitating motive. This finding is consistent with a well-established result in the criminological literature: women in abusive relationships face a particularly high risk of lethal violence in the period surrounding separation and divorce. Non-domestic femicides, by contrast, are more frequently associated with the rejection of romantic or sexual advances, prior stalking or harassment, or occur in the context of another crime, such as sexual assault or robbery (Tripi et al., 2026). The prominence of separation and divorce among the precipitating motives of domestic femicide is precisely what motivates our focus on studying the effects of a divorce law reform on this outcome.

We validate our manually constructed dataset by comparing it with administrative data on female homicide victims provided by the Italian Ministry of the Interior and ISTAT. First, we use a customized statistical elaboration from the administrative ISTAT Cause of Death registry (*Indagine sulle cause di morte*), which reports the number of deaths at the municipal level, by sex and cause of death, based on the International Classification of Diseases (ICD) – the WHO's standard system for classifying causes of death and disease using alphanumeric codes. For privacy protection reasons, the registry does not disclose the cause of death when fewer than three deaths are recorded in a given municipality-year. This measure is necessarily an upper bound on femicide, since not all female homicide victims are killed for gender-related reasons, but provides a useful alternative proxy. Nonetheless, if our manually collected data are accurate, one would expect it to correlate strongly with our main measure. Figure A.4 in the Appendix presents a binned scatter plot of the correlation between the number of femicides recorded

[15] See Figure A.3 for further details on the construction of these categories.

in our dataset and the number of female homicide victims derived from the ISTAT registry. The figure reveals a strong positive relationship between the two measures: municipalities with higher numbers of female homicide victims in the official registry also exhibit higher numbers of femicides in our dataset, suggesting that our dataset accurately captures the spatial distribution of lethal violence against women.

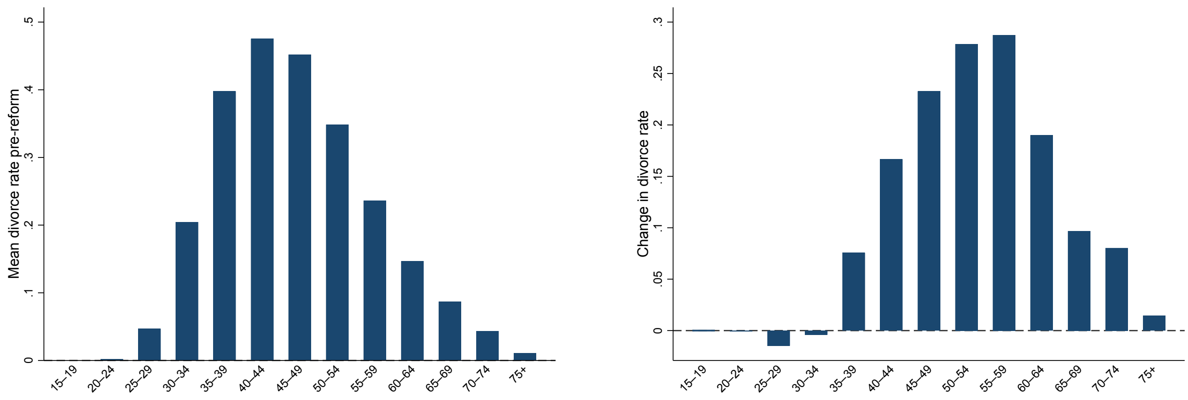
Second, we compare the aggregate time series of our manually collected data with the time series of female victims of homicide according to administrative data. Figure A.5 reports the evolution of female homicide victims recorded by the Ministry alongside femicides identified in our dataset. The addition of new sources over time improves the precision of our data in capturing officially recorded female homicides: starting in 2012, the number of femicides identified in our dataset closely mirrors the number of female homicides reported in official statistics. As discussed above, not all homicides involving women are classified as femicides, which accounts for the small but persistent gap between the two series. Overall, this approach allows us to construct a comprehensive and harmonized dataset of femicide incidence that minimizes undercounting and measurement error while providing granular case-level information unavailable in standard public data sources.

4 Empirical Strategy

The goal of the paper is to identify whether and how easier access to divorce, induced by the two reforms introduced in 2014 and 2015, affects the risk of femicide. As shown in Figure 1, these reforms induced an immediate, large, and persistent increase in the number of new divorces, suggesting that indeed the reforms succeeded in their intended aim of reducing the financial and time costs attached to the decision to divorce.

A key empirical challenge is that both reforms were introduced simultaneously across Italy for all individuals, so that there are no easily defined treated and control regions (or individuals). To address this issue, we leverage the fact that the propensity to divorce varies substantially across age groups. While understanding the underlying economic and sociological reasons why such differences emerge is beyond the scope of our work, we argue that they reflect how the expected costs and benefits of divorce depend on age: few young people divorce – for example because individuals tend to avoid breaking up during the early years of marriage or because they have young kids – and few old people divorce – for example, because the chances of finding a new partner deteriorate over time. Conversely, most divorces are concentrated around mid-age, sufficiently away from the wedding and from the kids' birth but while (re-)entering the marriage market is still feasible. Figure 3a documents this pattern in the pre-reform period while Figure 3b shows that the reform-induced change in divorce rates between 2014 and 2015

Figure 3. Divorce rates



(a) Mean divorce rate pre-reform by age group

(b) Change in divorce rate by age group

Notes: Panel (a) plots the mean annual age-specific divorce rate over the pre-reform period (2007–2014), computed as the number of newly divorced individuals in each age group divided by the corresponding population in that age group. Panel (b) plots the change in this rate for each age group between 2014 and 2015; positive values indicate an increase in the divorce rate in 2015 relative to 2014. In both panels, values are expressed in percentage terms.

largely follows the same age gradient.

We exploit this pre-existing demographic heterogeneity to construct a difference-in-differences design where we predict the local exposure to the two reforms by constructing a shift-share exposure measure. The aggregate incidence of divorce by age bracket (measured as of the 2011 Population Census, i.e., prior to the reform) represents the "shift" component while the local exposure to the common age-specific national shock stems from variation across municipalities in the 2011 Population Census population age structure (the "shares"). Formally, we define:

$$d_m = \sum_{a=1}^{A} \frac{Pop_{a,m,2011}}{Pop_{m,2011}} \times \frac{D_{a,2011}}{Pop_{a,2011}} \tag{1}$$

where $\frac{Pop_{a,m,2011}}{Pop_{m,2011}}$ denotes the share of individuals in age group a in municipality m at the time of the 2011 census[16] (the "share"), and $\frac{D_{a,2011}}{Pop_{a,2011}}$ represents the national number of newly divorced individuals relative to the population for age group a in 2011 (the "shift").[17] Prior to estimation, we standardize d_m to have unit variance, so that coefficients throughout the paper are interpreted as the effect of a one-standard-deviation increase in exposure to the reform. Intuitively, d_m captures the share of individuals in a municipality who belong to the age brackets in which

[16]By construction, for each municipality m, $\sum_{a=1}^{A} \frac{Pop_{a,m,2011}}{Pop_{m,2011}} = 1$. This is important to obtain unbiased shift-share estimates (Borusyak et al., 2022).

[17]D_a captures the national number of newly divorced individuals for age group a in 2011 while Pop_a represents the national population in 2011.

divorce is more likely to occur; that is, these are the individuals more likely to benefit from the two reforms. Figure A.6 and Figure A.7 in the Appendix illustrate, respectively, the geographic distribution of the instrument, d_m, and of the shares component.

Both components of the exposure measure are predetermined with respect to the reforms. The shares capture municipalities' age composition in the 2011 Census, while the shifts capture national age-specific divorce rates in the same pre-reform year. Because our specification includes municipality fixed effects, any time-invariant differences across municipalities associated with their demographic composition are absorbed. Identification therefore requires that, absent the reforms, municipalities with different predicted exposure would not have experienced systematically different changes in femicide risk. We assess the plausibility of this assumption through dynamic specifications documenting the absence of differential pre-trends, by allowing a rich set of predetermined municipal characteristics to have year-specific effects, and through alternative constructions of both the share and shift components reported in the robustness section.

Upon constructing a panel of municipalities (current boundaries) for the period 2007–2021, we estimate the following reduced-form regression model:

$$F_{m,t} = \beta_{RF} * (d_m \times PostReform_t) + X_{m,2011} \sum_t \lambda_t + \eta_m + \theta_t + \varepsilon_{m,t} \tag{2}$$

We focus on a number of main dependent variables. First, *At least one femicide (any killer)* is an indicator variable equal to one if at least one femicide is recorded in the municipality-year. Second, *At least one domestic femicide* restricts this to femicides perpetrated by a current partner, an ex-partner, or a relative.[18] As a complementary check, we construct *At least one femicide (other killers)*, capturing all remaining cases, to verify that our results are not driven by femicides outside the domestic sphere. We also produce results where we construct counts by municipality-year of the former variables. The coefficient of interest is β_{RF}, which identifies, after the introduction of the two reforms, the effect of the change in the share of individuals at risk of divorce on the dependent variable. η_m denotes municipality fixed effects, while θ_t denotes the year fixed effects. In some specifications, we also control for a vector of predetermined characteristics measured at the 2011 Population Census, $X_{m,2011}$, which includes the share of women in the population, the share of foreign-born in the population, the share of newborns (aged 0-4), and the share of college-educated residents, all interacted with time dummies. We cluster the standard errors at the municipality level. Equation 2 estimates an Intention-To-Treat model: we focus on the fact that the reform increases the risk of divorcing by reducing its financial and time costs, rather than on the individuals who actually decide to divorce because

[18] We focus on the domestic subsample in several specifications, both because it can be unambiguously classified as femicide, minimizing concerns about subjective interpretation, and because it enables meaningful comparisons with official pre-2019 statistics.

of the reform.

To document that our shift-share exposure measure indeed predicts the increase in the number of divorced individuals induced by the reform, we also estimate a first-stage specification of the form:

$$D_{m,t} = \pi * (d_m \times PostReform_t) + X_{m,2011} \sum_t \lambda_t + \eta_m + \theta_t + \upsilon_{m,t} \tag{3}$$

where $D_{m,t}$ is the number of individuals who reside in municipality m and obtain divorce in year t, scaled by the total municipal population in the same year, the coefficient π captures the differential increase in divorces after the reform among municipalities with higher exposure as measured by the shift-share variable d_m.[19]

For completeness, we also report two-stage least squares (2SLS) estimates as a convenient rescaling of the reduced-form magnitude by the reform-induced change in the municipal divorce rate. Our preferred causal estimand remains the reduced-form effect of differential exposure to the reforms. A strict interpretation of the 2SLS coefficient as the causal effect of divorce incidence itself would require the exclusion restriction that the reforms affect femicide only through completed divorces, an assumption that is unlikely to hold in our setting because the reforms also alter bargaining, separation incentives, and the composition of divorces.

Formally, in our 2SLS model the second-stage takes the form:

$$F_{m,t} = \beta_{2SLS} * \hat{D}_{m,t} + X_{m,2011} \sum_t \lambda_t + \eta_m + \theta_t + \varepsilon_{m,t} \tag{4}$$

where $\hat{D}_{m,t}$ denotes the share of newly divorced individuals as a share of the population in municipality m and year t obtained from estimating Equation 3.

5 Results

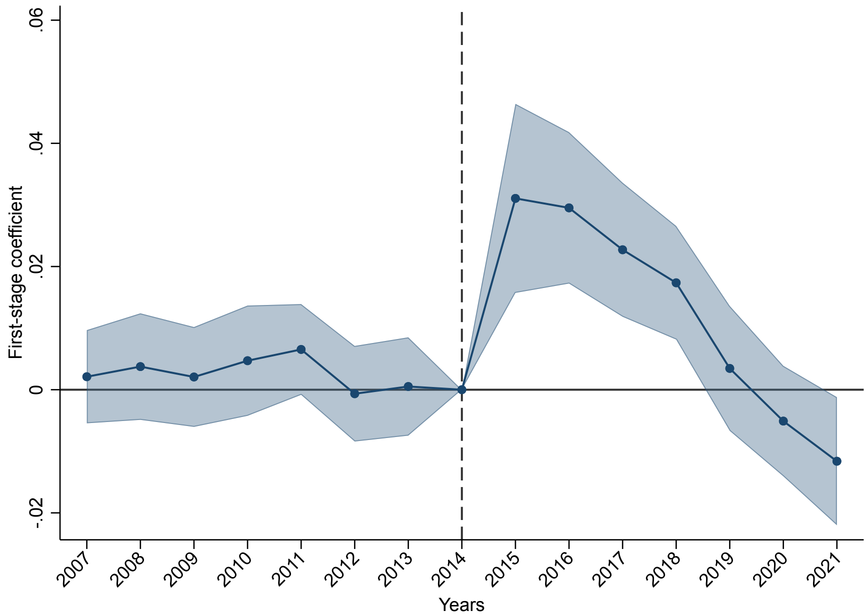
5.1 First-stage: Effect on Divorces

Divorce rates. The first-stage specification is given in Equation 3. The coefficient of interest, π, captures whether municipalities with greater pre-reform exposure to the reform – as measured by our shift-share instrument d_m – experienced a differentially larger increase in the share of newly divorced individuals after the reform.

Results are reported in Table B.2. The coefficient π is positive and statistically significant across all specifications, and remains stable as we progressively add controls: column (1) re-

[19] We obtain yearly municipal population counts from ISTAT's *Ricostruzione della popolazione* for the inter-censal years and from *Popolazione residente* for the most recent years.

Figure 4. Parallel trends



Notes: The figure plots the estimated coefficients ω_y from Equation 5, with 2014 as the omitted baseline year. The shaded area denotes 95% confidence intervals, based on standard errors clustered at the municipal level. The vertical dashed line marks the baseline year, 2014.

ports the specification without controls, column (2) adds a control for municipal population, and column (3) further adds controls for the female share, the foreign-born share, the share of newborns (aged 0-4), and the share of college-educated residents. Quantitatively, a one-standard-deviation increase in our standardized measure of reform exposure raises the municipal share of newly divorced individuals by approximately 0.0101 percentage points (column (3)), corresponding to a 7.4% increase relative to the pre-reform mean of 0.13%. The first-stage F-statistic is 19.375, above the conventional threshold of 10 (Stock and Yogo, 2005).

We next examine the dynamic relationship between predicted reform exposure and divorce rates. This exercise assesses whether municipalities with different levels of predicted exposure exhibited differential trends before the reforms and documents when the first-stage relationship emerges relative to reform implementation. To this end, we estimate the following dynamic specification:

$$D_{m,t} = d_m \times \sum_{k \neq 2014} \omega_k 1(Year_t = k) + X_{m,2011} \sum_t \lambda_t + \rho_m + \gamma_t + \upsilon_{m,t} \tag{5}$$

In Equation 5, we interact the shift-share instrument d_m with year dummies for each year k, excluding the pre-reform year 2014, which we take as the omitted baseline. Here, ρ_m denotes municipality fixed effects, γ_t year fixed effects, and $\upsilon_{m,t}$ the error term.

Figure 4 reports the estimated coefficients for each year between 2007 and 2021. The pre-reform coefficients fluctuate around zero and are statistically indistinguishable from zero, providing no evidence of differential pre-trends prior to the reform. This pattern supports the validity of the parallel-trends assumption underlying our research design. Starting in 2015, coinciding with the implementation of the reform, the coefficients increase markedly and then gradually decline over time. The only significant and negative coefficient in the post-reform period is observed in 2021, which likely reflects the impact of COVID-19-related lockdowns and disruptions. Taken together, the evidence in Figure 4 validates our research design and shows that our strategy effectively captures territorial differences in exposure to the divorce law reforms.

Divorce application rates. We also establish that our shift-share measure of exposure to the divorce law reforms significantly increases the probability of new divorce applications. To this end, in Table B.3 we use as the dependent variable the share of residents who file a divorce application in a given year, regardless of whether the divorce was subsequently granted. The results document that the first stage remains strong, with first-stage F-statistics well above conventional thresholds across all specifications.

Composition of divorces. The results above document that the reform increased the overall number of divorces. However, as discussed in Section 4, the reform may have also affected which couples divorce, altering the composition of divorces along dimensions such as the type of procedure, the marital property regime, and the terms of the settlement. To assess this directly, we estimate a reduced-form specification analogous to Equation 2, but where the dependent variable is the incidence of a number of divorce characteristics.[20]

Table B.4 in the Appendix reports the results. Municipalities more exposed to the reform experience a significant shift in the composition of divorces: the share of judicial divorces rises by 1.7 percentage points, the share involving a shared property regime (*comunione dei beni*) increases by 1.5 percentage points, and the shares awarding joint custody and spousal transfers decline by 1.8 and 0.6 percentage points, respectively. By contrast, the share of divorces awarding custody to the father and the share involving children show no significant change.

Taken together, these results indicate that the reform did not simply scale up the pre-existing pool of divorces, but changed its composition: divorces occurring after the reform are relatively more likely to require judicial rather than consensual proceedings, to involve shared assets, and to end without joint custody or spousal transfers – patterns broadly consistent with divorces that

[20]ISTAT granted access to this information only as pre- and post-reform totals, rather than at the yearly level, as part of the customized statistical elaboration described in subsection 3.1. Therefore, we collapse the panel into two periods, pre- and post-reform, rather than exploiting yearly variation.

Table 1. Probability of femicide - reduced-form

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable =	At least one femicide (any killer)		At least one domestic femicide		At least one femicide committed by other killers	
Post-reform × Instrument (std.)	0.0011**	0.0011**	0.0013***	0.0012**	-0.0000	0.0000
	(0.0005)	(0.0006)	(0.0004)	(0.0005)	(0.0002)	(0.0003)
Obs	118410	118410	118410	118410	118410	118410
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.013	0.013	0.011	0.011	0.003	0.003
R-squared	0.182	0.183	0.168	0.169	0.163	0.176

Notes: OLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *Post-reform* is a dummy equal to 1 from 2015 onward and 0 otherwise; *Instrument (std.)* is the standardized measure defined in Equation 1. The dependent variable is an indicator equal to one if at least one femicide occurred in the municipality, committed by any killer in columns (1)–(2), by a partner, ex-partner, or relative (domestic femicide) in columns (3)–(4), and by another type of perpetrator (e.g., professional relation, acquaintance, stalker, or unknown) in columns (5)–(6). Within each pair of columns, the first specification excludes pre-determined controls and the second includes them: population, the shares of women, children aged 0–4, individuals with tertiary education, and foreign residents, all measured in the 2011 Census and interacted with year fixed effects.

are, on average, more contentious or involve a less favorable bargaining position for one spouse. This is consistent with the reform lowering the cost of divorce enough to induce previously-reluctant couples – those facing higher-conflict separations – to complete the separation-to-divorce process.[21]

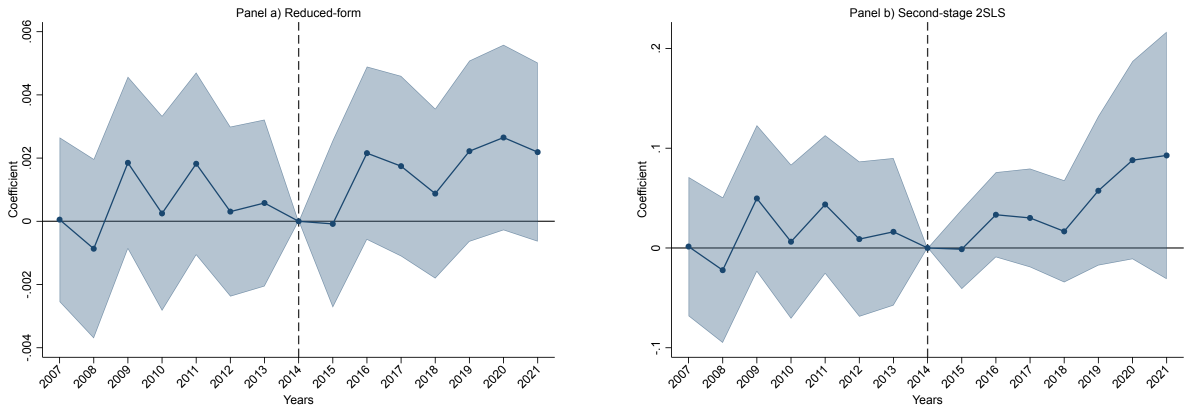
5.2 Exposure to Divorce Law Reforms and Probability of Femicide

In this section, we present our estimates of the causal effect of exposure to the divorce law reform on the probability that at least one femicide is perpetrated in the municipality. As our main specification, we use the reduced-form regression in Equation 2. This allows us to capture not only the effect on individuals who actually divorce because of the reform, but also on those who remain married yet benefit from an improved outside option. We complement this with the 2SLS estimates from Equation 4, to ease the interpretation of the magnitude of our results.

Table 1 reports the reduced-form estimates for the probability that at least one femicide occurs. Columns (1)-(2) consider all femicides, columns (3)-(4) focus on domestic femicides (i.e., cases committed by a partner, ex-partner, or relative), and columns (5)-(6) examine femicides committed by other perpetrators. The latter outcome may be interpreted as a *placebo* because divorce law reforms should not impact the probability of femicides perpetrated outside of the family sphere. Odd-numbered columns report estimates without the inclusion of predetermined

[21]These findings reinforce the discussion in Section 4 regarding the interpretation of our 2SLS estimates: because the reform affects not only the number of divorces but also their composition, the exclusion restriction required for a strict LATE interpretation is unlikely to hold. We therefore continue to treat the reduced-form coefficient as our preferred causal estimand throughout the paper.

Figure 5. Dynamic treatment effects



(a) Reduced-form

(b) Second-stage 2SLS

Notes: The figure plots the estimated coefficients from the dynamic specifications of Equation 2 (Panel (a)) and Equation 4 (Panel (b)). The dependent variable is a dummy equal to one if at least one domestic femicide occurred in municipality m in year t. The shaded area denotes 95% confidence intervals, based on standard errors clustered at the municipal level. The vertical dashed line marks the last pre-reform year, 2014.

covariates, while even-numbered columns include the interaction between a number of predetermined control variables interacted with year dummies. We find that exposure to the reform causes a statistically significant increase in the probability of femicide and that the estimates are remarkably stable to the inclusion of controls. According to our preferred specification (column (2)), a one-standard-deviation increase in reform exposure raises the probability of observing at least one femicide by approximately 0.11 percentage points. Given a pre-reform mean of 1.3%, this corresponds to an increase of about 8.5% relative to baseline levels. The effect is driven exclusively by an effect on femicides committed within the family sphere: column (4) implies that a one-standard-deviation increase in reform exposure increases the probability that a domestic femicide is perpetrated in the municipality by 0.12 percentage points, or roughly 10.9% relative to the pre-reform mean. By contrast, our estimates are quantitatively negligible and statistically indistinguishable from zero for femicides committed by other perpetrators (columns (5)-(6)).

For completeness, Table B.5 in the Appendix reports 2SLS estimates that rescale the reduced-form magnitude by the reform-induced increase in the municipal divorce rate. The results document that a one-percentage-point increase in the share of newly divorced individuals raises the probability of observing at least one femicide by approximately 0.112 percentage points (column (2)), with a similar magnitude among domestic femicides (column (4): 0.117 percentage points) and no detectable effect for femicides committed by other perpetrators (column (6)).

We report the dynamics of both the reduced-form and 2SLS specifications jointly in Figure 5. In the pre-reform years, coefficients fluctuate around zero and are statistically insignificant,

Table 2. Number of femicides - reduced-form

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable =	Total number of femicides		Total number of domestic femicides		Total number of femicides (other killers)	
Post-reform × Instrument (std.)	0.0016***	0.0016**	0.0016***	0.0014**	0.0000	0.0001
	(0.0006)	(0.0007)	(0.0005)	(0.0006)	(0.0002)	(0.0003)
Obs	118410	118410	118410	118410	118410	118410
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.017	0.017	0.013	0.013	0.004	0.004
R-squared	0.366	0.383	0.289	0.314	0.220	0.259

Notes: OLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is the total number of femicides in the municipality in columns (1)–(2), the number of domestic femicides (committed by a partner, ex-partner, or relative) in columns (3)–(4) and the number of femicides committed by other killers (e.g., professional relation, acquaintance, stalker, or unknown) in columns (5)-(6). Control variables are described in Table 1.

supporting the absence of differential pre-trends. Beginning in 2016, the year following the reform, the estimated effects turn positive and remain consistently above zero for the rest of the period. Despite a loss in precision relative to our main regressions, the pattern is broadly consistent across the reduced-form and 2SLS specifications and further corroborates the validity of our empirical design.

5.3 Exposure to Divorce Law Reforms and Number of Femicides

In the previous section, we focused on the extensive margin, that is, how exposure to the reforms shapes the probability that a femicide occurs in a municipality. We now turn to analyzing the intensive margin, estimating whether the reform also affected the number of femicides per municipality.

As shown in Table 2 and in Table B.6 in the Appendix, an increase in exposure to divorce law reforms causes a statistically significant increase in the number of femicides. The reduced-form estimates imply that a one-standard-deviation increase in the instrument increases the number of femicides by approximately 0.0016 cases, a 9.4% increase relative to the pre-reform mean of 0.017.[22] The effect remains positive and statistically significant when restricting the outcome to domestic femicides (0.0014, a 10.8% increase relative to a pre-reform mean of 0.013). No significant effect is found for femicides committed by other perpetrators (columns (5)-(6)). The corresponding second-stage 2SLS estimates reported in Table B.6 in the Appendix broadly confirm this pattern.[23]

[22] Taking into consideration that the population of the average municipality as of the 2011 census was 7529 inhabitants, 0.0016 cases correspond to an increase of 0.02 femicides per 100,000 inhabitants.

[23] Dynamic estimates from both the reduced-form and the 2SLS specifications are plotted in Figure A.8 in the

In this analysis, the dependent variable features a large number of zeros because no femicides occur in the majority of municipality-year pairs. In Table B.7 in the Appendix, we follow standard practices in the empirical literature and report both reduced-form (Panel (a)) and second-stage 2SLS (Panel (b)) estimates using alternative functional transformations, to rule out the possibility that our results are driven by the estimation of a linear model in a setting with many zeros. First, we use the inverse hyperbolic sine transformation of the number of femicides, which allows us to retain zero observations in the sample while approximating a logarithmic scale for positive values, thereby mitigating the influence of extreme observations. Second, we estimate a pseudo-Poisson maximum likelihood model, which is particularly suited for count data characterized by many zeros (Chen and Roth, 2024; Silva and Tenreyro, 2006), both in its standard specification and in a rate specification that includes yearly municipal population as an exposure offset.[24] All three approaches yield results consistent with our main findings, with positive and statistically significant estimates across all specifications.

5.4 Exposure to Divorce Law Reforms: Heterogeneity by Victim-Killer Relationship and Motive

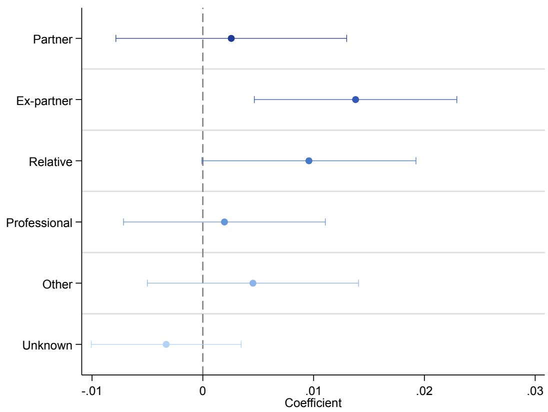
We decompose the overall effect of the reform along two dimensions: the relationship between victim and perpetrator, and the reported motive of the femicide. To this end, we redefine the dependent variables to capture municipality-year indicators for only selected types of femicide. **Victim-killer relationship.** We distinguish between (i) current partners, (ii) ex-partners, (iii) relatives, (iv) professional acquaintances, (v) other known individuals, and (vi) unknown perpetrators. As shown in Figure 6a, the largest and most precisely estimated effects emerge for ex-partners and relatives, while no significant effect emerges for the remaining categories. The prominence of the ex-partner category, which includes divorced and legally separated individuals as well as former cohabiting partners, is consistent with the reform primarily affecting violence arising in connection with relationship dissolution. The effect on relative-perpetrated femicides is less directly connected to marital dissolution and should therefore be interpreted cautiously. In our data, a substantial share of these episodes involve adult sons killing an elderly mother. Changes in household composition following separation may generate spillovers within the broader family, but our data do not allow us to establish the underlying mechanism directly.

Reassuringly and consistently with columns (5) and (6) of Table 1, we cannot detect any

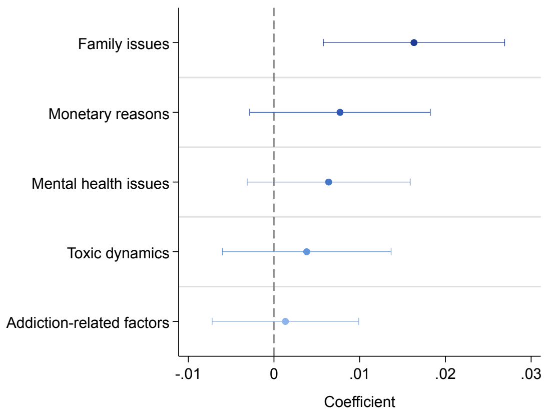
Appendix.

[24] It should be noted that the number of observations decreases substantially when estimating the pseudo-Poisson specifications, as the model automatically drops from the observations count the units that do not contribute to the estimation. This reduction is expected and does not compromise the robustness or significance of the estimated effects.

Figure 6. Effect decomposed by victim-killer relationship and motive



(a) Victim-killer relationship



(b) Motive of the femicide

Notes: The figure plots the estimated coefficients from Equation 2 using alternative femicide classifications. Panel (a) classifies femicides based on the relationship between victim and perpetrator: current partner, ex-partner, relative, professional acquaintance, other known individuals not falling into the previous categories, and unknown perpetrators (either undefined or not personally known to the victim). Panel (b) classifies femicides based on the reported motive, as documented in media sources. *Family issues* include cases in which one of the following motives was reported: illness, abandonment, betrayal, quarrels, jealousy, honor-related conflicts, domestic abuse, or disputes involving children or other relatives. *Monetary reasons* include cases involving debt, extortion, broader financial distress, inheritance disputes, or conflicts over separation payments or wages. *Mental health issues* include cases citing depression, psychological distress, or other mental health conditions. *Toxic dynamics* include cases involving sexual assault, revenge, raptus, unrequited love, or possessive behavior. *Addiction-related factors* include cases involving alcohol abuse, drug abuse, or gambling. For each classification, the dependent variable is an indicator equal to one if at least one femicide of that type (Panel (a)) or with that reported motive (Panel (b)) occurred in the municipality-year, standardized using its pre-reform mean and standard deviation. Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Control variables are defined in Table 1.

statistically significant effect on femicides perpetrated outside of the family sphere.

Motive. We next examine which motives – as reported in news coverage based on police investigations – underlie the reform's effect on femicide. We aggregate motives into five categories: *Family issues* (illness, abandonment, betrayal, quarrels, jealousy, honor-related conflicts, domestic abuse, or disputes involving children or other relatives); *Monetary reasons* (debt, extortion, financial distress, inheritance disputes, or conflicts over separation payments or wages); *Mental health issues* (depression, psychological distress, or other mental health conditions); *Toxic dynamics* (sexual assault, revenge, raptus, unrequited love, or possessive behavior); and *Addiction-related factors* (alcohol abuse, drug abuse, or gambling). Because reported motives are reconstructed from news coverage and police investigations, we interpret this exercise as a descriptive decomposition rather than a direct test of mechanism. The exercise is still helpful because divorce law reforms should affect femicides that occur within the family sphere rather than other types of femicide. Figure 6b shows that the positive effect is concentrated among cases classified as involving family-related motives, while the estimated effects for the remaining categories are small and statistically insignificant. Reassuringly about the overall validity of our empirical design, this pattern is consistent with the main effect being concentrated within the family sphere.

6 Robustness checks

In this section, we conduct a series of robustness checks to assess both the validity of our data and the stability of the main results.

6.1 Robustness to Alternative Data Sources

Administrative Records. Table 3 and Table B.8 report reduced-form and second-stage 2SLS estimates, respectively, that replicate our main results using the administrative data on female victims of homicide described in Section 3 to construct the dependent variables. The results using administrative records in column (2) are remarkably similar to our baseline estimates (replicated in column (1) for convenience) both in magnitude and statistical significance. Column (3) restricts the sample to the balanced panel of municipalities that are never subject to privacy-related data disclosure suppression. Despite losing around one-third of observations, the estimated effect remains similar to our baseline results and, if anything, larger in magnitude. Columns (4) and (5) replicate columns (2) and (3), respectively, using the number of female homicide victims as the dependent variable instead of an indicator; the results are qualitatively similar to our baseline estimates.

Placebo Causes of Death. We construct four placebo outcomes using the ISTAT *Cause di morte* administrative registry to verify that the reform did not affect female mortality for causes that are arguably unrelated to family dynamics. First, we construct an indicator equal to one if at least one woman dies in a transport accident in the municipality-year.[25] We then construct three additional placebo outcomes capturing deaths from diseases that are unlikely to be directly affected by changes in family dynamics: indicators equal to one if at least one woman dies due to (i) blood diseases, (ii) respiratory diseases, or (iii) musculoskeletal diseases.[26] As

[25] ICD-10 codes for transport accidents are V01-V99. These codes include accidents involving pedestrians, cyclists, motor vehicles, buses, trains, as well as water and air transport.

[26] Blood diseases correspond to ICD-10 codes D50–D89 and include conditions such as iron deficiency anemia, hemolytic anemias, coagulation disorders (e.g., hemophilia), and immune disorders. Respiratory diseases correspond to ICD-10 codes J00–J99 and include acute respiratory infections, chronic obstructive pulmonary disease (COPD), pneumonia, and asthma. Musculoskeletal diseases correspond to ICD-10 codes M00–M99 and include conditions such as rheumatoid arthritis, osteoarthritis, osteoporosis, and connective tissue disorders.

Table 3. Probability and number of female homicides - reduced-form

	(1)	(2)	(3)	(4)	(5)
Dependent variable =	At least one femicide (any killer)	At least one female homicide	At least one female homicide (balanced sample)	Number of female homicides	Number of female homicides (balanced sample)
Post-reform × Instrument (std.)	0.0011**	0.0014***	0.0030***	0.0014**	0.0029**
	(0.0006)	(0.0005)	(0.0011)	(0.0006)	(0.0012)
Obs	118410	117859	80925	117859	80925
Municipalities	7894	7894	5395	7894	5395
Year and Mun. FE	✓	✓	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓	✓	✓
Mean dep. var pre-reform	0.013	0.013	0.019	0.015	0.022
R-squared	0.183	0.171	0.170	0.357	0.360

Notes: OLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable in column (1) is the main dependent variable described in Table 1. In column (2), the dependent variable is an indicator equal to one if at least one female homicide occurred in the municipality, constructed using the ISTAT *Cause di morte* registry, which classifies deaths by underlying cause; homicides are identified using ICD codes X85–Y09 and Y87.1. Column (3) restricts the sample to municipalities in which the exact count of female homicide victims is never suppressed for privacy reasons. Columns (4) and (5) replicate columns (2) and (3), respectively, using as the dependent variable the number of female homicide victims rather than an indicator variable. Control variables are described in Table 1.

shown in Figure A.9, the reduced-form estimates are small and statistically insignificant across specifications for all four outcomes.

6.2 Robustness to Sample Restrictions

Excluding 2021. In Figure 4, we observe that 2021 is the only year in which the coefficient turns negative and statistically significant. This anomaly likely reflects the disruption caused by the COVID-19 pandemic: divorces finalized in 2021 largely stem from separations initiated in 2020, a year in which lockdowns and restrictions on mobility and court activity led to a sharp decline in the number of separations. To rule out the possibility that our results are confounded by this sign inversion, we assess whether our results are sensitive to the inclusion of 2021. Table B.9 in the Appendix reports the results. Reassuringly, the estimates remain positive and statistically significant, indicating that our findings are not driven by the anomalous behavior of the instrument in 2021.

Leave-One-Region-Out. In Figure A.10, we replicate our reduced-form estimates by excluding one region at a time. The coefficients remain positive and of similar magnitude throughout, and remain statistically significant in the vast majority of specifications, indicating that our results are not driven by municipality-level idiosyncrasies concentrated in any single region.

6.3 Robustness to Instrument Construction

Alternative Instrument Constructions. We assess the robustness of our shift-share instrument to several alternative construction choices, summarized in Figure A.11 in the Appendix. First, motivated by concerns about the exogeneity of the share component, we recompute municipal age-group population shares using the 2001 census instead of the 2011 census. Second, we compute the age-specific shares using the 2014 population structure (i.e., immediately prior to the reforms), which more closely reflects the age composition of municipalities at the time of treatment.[27] Third, in addition to using 2014 shares, we also fix the national shift component at its 2014 level, so that both components of the instrument are measured immediately before the reforms. Fourth, we construct a leave-out version of the shift by excluding, for each municipality, divorces and population counts from the region to which the municipality belongs when computing the national age-specific divorce rates, in the spirit of the leave-out shift-share instruments proposed by Borusyak et al. (2022, 2025). Finally, we implement the shock-level estimator recommended by Borusyak et al. (2022, 2025), aggregating the analysis to the level of the shocks (age groups) using the `ssaggregate` Stata command. Across all these alternative specifications, the estimated coefficients remain positive and of a similar order of magnitude to our baseline estimate; all are statistically significant at conventional levels, with the exception of the shock-level estimate, which is significant at the 10% level. Overall, this indicates that our results are not an artifact of particular choices in the construction of the exposure measure.

7 Heterogeneous Treatment Effects

In this section, we investigate whether the estimated effect of the reform on femicide conceals meaningful heterogeneity across a number of municipal characteristics that allow us to better understand the mechanisms through which the reform operates. We consider four broad dimensions: (i) attitudes toward divorce and the composition of pre-existing divorces; (ii) religious presence and social capital; (iii) pre-existing levels of intentional homicides and domestic violence; and (iv) income levels and gender gaps in employment and educational attainments. For each dimension, we construct indicator variables equal to one if the relevant municipal characteristic, measured prior to the reform, is above the national median, and zero otherwise. We then interact the indicator with our main treatment variable and estimate:

$$\begin{aligned} F_{m,t} &= \beta_1 \left(d_m \times PostReform_t\right) + \beta_2 \left(d_m \times PostReform_t \times Z_{m,2011}\right) \\ &\quad + \beta_3\, PostReform_t \times Z_{m,2011} + X_{m,2011} \sum_t \lambda_t + \eta_m + \theta_t + \varepsilon_{m,t} \end{aligned} \tag{6}$$

[27]We obtain yearly municipal population counts from ISTAT's *Ricostruzione della popolazione* for the intercensal years and from *Popolazione residente* for the most recent years, which we use to construct all outcome and treatment variables expressed as shares of the municipal population.

where $F_{m,t}$ is our main dependent variable, an indicator for at least one domestic femicide in municipality m and year t, and $Z_{m,2011}$ denotes the above-median indicator for the relevant heterogeneity dimension. The coefficient of interest, β_2, captures how the effect of the reform varies with the baseline characteristic Z. All estimates reported in this section are reduced-form.

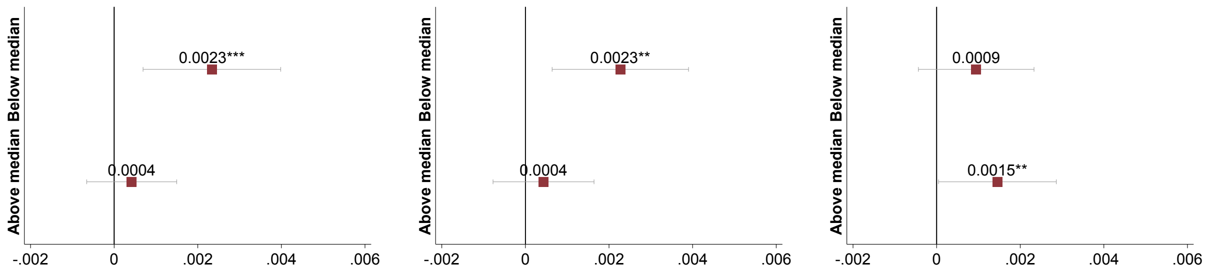
7.1 Divorce Attitudes and Composition

Divorce Attitudes. We first examine whether the effect of the reform varies with local attitudes toward divorce, which proxy for the presence of social or normative barriers to marital dissolution prior to the reform. We exploit administrative data from the Italian Ministry of Internal Affairs (*Eligendo*) containing municipality-level voting outcomes of the 1974 Italian divorce referendum, in which voters were asked whether to repeal the 1970 law that had introduced divorce in Italy. A vote in favor of repeal ("Yes") implied opposition to divorce, while a vote against repeal ("No") implied support for the divorce law. We construct an indicator equal to one if the municipal-level share of "Yes" votes is above the national median, capturing relatively stronger attitudes against marital dissolution.

The results, reported in column (1) of Table B.10 and Panel (a) of Figure 7, show that the interaction term is negative and statistically significant: the effect of the reform on femicide is concentrated in municipalities with historically weaker opposition to divorce, and is not statistically distinguishable from zero in municipalities above the median. This pattern is consistent with the reform's legal change translating into a more credible threat of separation where social norms are less likely to limit the feasibility of marital dissolution. In municipalities traditionally experiencing weaker support for divorce, the legal and financial burdens that the two reforms addressed are plausibly not the main constraint faced by couples who may otherwise decide to break up.

Divorce Composition. As shown in Table B.4, the reform itself shifts the composition of divorces toward relatively more judicial proceedings. Motivated by this finding, we examine whether the effect of the reform on femicide varies with the pre-existing composition of divorces in the municipality, focusing on two dimensions: the share of divorces settled consensually between spouses, and the share involving kids. Both variables plausibly relate to the potential for conflict surrounding marital dissolution: a lower share of consensual divorces may signal a more contentious divorce environment, while cases involving children are more likely to generate disputes over custody and child support. Importantly, the two variables also give us the opportunity to discriminate between the two reforms: indeed, the 2014 reform allowed couples to divorce in front of a municipal official only in the case of a consensual divorce with no minor-age kids involved. We construct indicator variables equal to one if the pre-reform

Figure 7. Heterogeneity: divorce attitudes and characteristics



(a) Divorce attitudes (b) Consensual divorces (%) (c) Divorces with children (%)

Notes: Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Each panel splits municipalities using an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise. In Panel (a), the dimension is divorce attitudes, measured as the share of voters who voted "Yes" in the 1974 divorce referendum. In Panel (b), the dimension is the pre-reform share of consensual divorces out of total divorces. In Panel (c), the dimension is the pre-reform share of divorces involving children out of total divorces. Control variables are defined in Table 1.

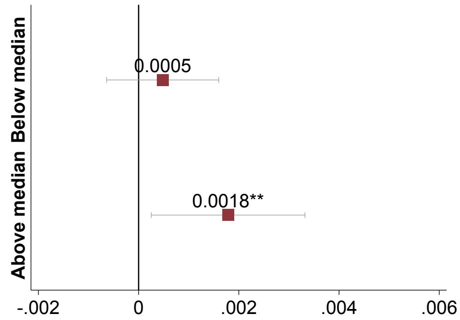
municipal share of consensual divorces, and correspondingly of divorces involving children, is above the national median.

The results, reported in Panels (b) and (c) of Figure 7 and in columns (2)-(3) of Table B.10 in the Appendix, show that the effect of the reform is larger and more precisely estimated in municipalities with a below-median share of consensual divorces, consistent with a negative and marginally significant interaction coefficient. This evidence suggests that conflictual separations are more likely to escalate into lethal violence.

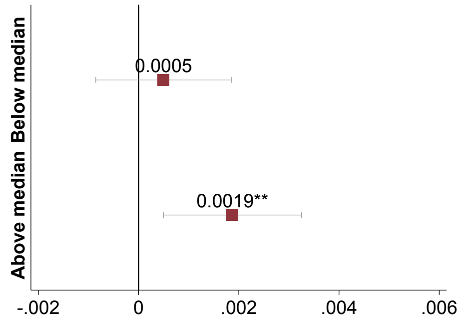
For divorces involving children, the interaction term is not statistically significant on its own; however, the total effect in municipalities with an above-median share of parents among individuals who divorce is positive and statistically significant, while the effect of higher exposure to the reform is not distinguishable from zero in municipalities where divorces are comparatively more likely to occur among childless couples. This result indicates that custody arrangements may trigger lethal violence.

Taken together, these two heterogeneity exercises strongly suggest that the 2015 law that reduced the mandatory period of legal separation from three years to six months appears to be the key driver of our main results. Indeed, the 2014 reform that allowed couples to appear in front of a municipal officer instead of a court required the couple to both reach a consensual divorce agreement and not to have any minor-age dependent kid. The evidence presented in Panels (b) and (c) of Figure 7 and in columns (2)-(3) of Table B.10 in the Appendix documents the opposite pattern.

Figure 8. Heterogeneity: crime



(a) Intentional homicides



(b) Domestic violence

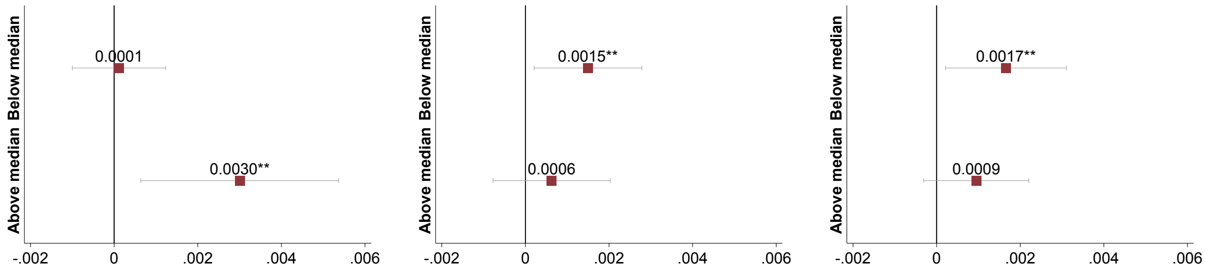
Notes: Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Each panel splits provinces using an indicator variable equal to one if the province is above the national median along a given dimension in the pre-reform period (2007–2013), and zero otherwise. In Panel (a), the dimension is the rate of intentional homicides per 100,000 inhabitants. In Panel (b), the dimension is the rate of domestic violence crimes per 100,000 inhabitants. Control variables are defined in Table 1.

7.2 Intentional Homicides and Domestic Violence

We next examine whether the effect of the reform varies with the pre-existing intensity of violent crime in the community. To this end, we rely on data on crimes related to domestic violence, obtained from ISTAT, which records crimes reported to the judicial authority by the State Police, the Carabinieri, and the Guardia di Finanza. These data are only available at the province-year level. First, we focus on the number of intentional homicides per 100,000 inhabitants. Second, we investigate heterogeneous effects with respect to the number of reports for domestic violence per 100,000 inhabitants. These exercises are motivated by the idea that the reform-induced shock to relationship dynamics may translate into lethal outcomes more easily in environments where conflict more often escalates into lethal forms of violence and where violent behavior in the household is relatively more common to start with. For each measure, we construct an indicator variable equal to one if the corresponding provincial crime intensity is above the national median.

We report the results in Figure 8 and in Table B.12 in the Appendix. Although the interaction terms are not individually significant, the effect of the reform is positive and statistically significant in provinces with above-median crime intensity, whether measured through intentional homicides or domestic violence, and is small and statistically indistinguishable from zero in provinces below the median. Taken together, these results document a pattern that is consistent with the effect of the reforms on femicides being concentrated in municipalities that already prior to the reform were more exposed to domestic violence and higher incidence of intentional homicides.

Figure 9. Heterogeneity: income, employment gap, and education gap



(a) Income (b) Employment rate gap (c) Education rate gap

Notes: Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Each panel splits municipalities using an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise. In Panel (a), the dimension is mean income at the municipal level in the pre-reform period (2007–2013). In Panel (b), the dimension is the employment rate gap, measured in the 2011 Census. In Panel (c), the dimension is the education rate gap, measured in the 2011 Census. Control variables are defined in Table 1.

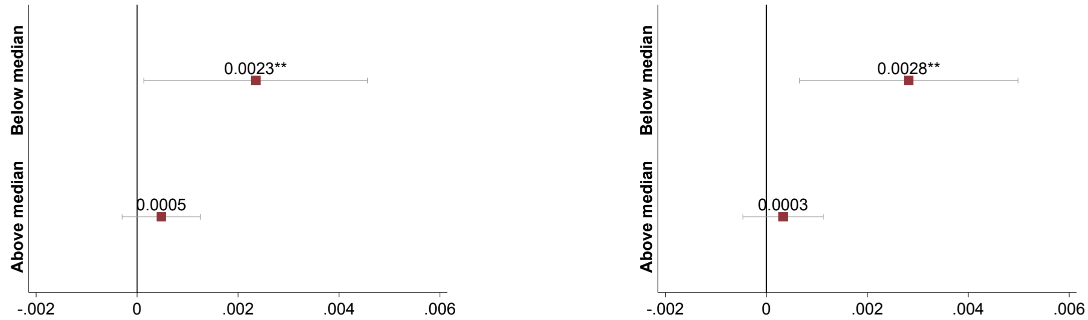
7.3 Average income and gender differences in employment and educational attainment

Finally, we explore heterogeneity across the average municipal income levels and gender gaps in the local community. These dimensions are motivated by a household bargaining perspective: the reforms can make the threat of separation more credible to abusive husbands who may eventually escalate to lethal violence if women have the economic means to take advantage of the reform. Conversely, where women are more economically dependent on their partner – due to weaker labor market attachment, lower education relative to men, or lower household income – leaving the relationship remains costly even after the legal barrier to divorce is removed, which may limit both the credibility of the threat and the extent to which it translates into conflict.

Income. We use data on the average municipality income in the pre-reform period using administrative data from tax records released by the Italian Ministry for the Economy and Finance. We then construct an indicator taking the value 1 if the municipality's average income exceeds that of the median municipality. The results, reported in Panel (a) of Figure 9 and in column (1) of Table B.13 in the Appendix, show a sharp pattern: the interaction term is positive and statistically significant, and the total effect in municipalities with above-median income is positive and statistically significant. Overall, this evidence is broadly consistent with a mechanism in which the reform increases women's bargaining power.

Employment and Education Gaps. We then investigate whether the effects are concentrated in contexts where women are more economically independent. To this end, we construct

Figure 10. Heterogeneity: religion and social capital



(a) Religious presence (b) Social capital

Notes: Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Each panel splits municipalities using an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise. In Panel (a), the dimension is religious presence, measured as the number of churches per 1,000 inhabitants. In Panel (b), the dimension is social capital, measured as the number of associations per 1,000 inhabitants. Control variables are defined in Table 1.

indicator variables equal to one if the municipal gender gap in employment, and correspondingly in education, measured in the 2011 Population Census, is above the national median. The results, reported in Panels (a) and (b) of Figure 9 and in columns (1) and (2) of Table B.13 in the Appendix, show that the effect of the reform is positive and statistically significant in municipalities with a below-median employment gap and a below-median education gap, while it is smaller and statistically indistinguishable from zero above the median for both measures. In neither case is the interaction term itself statistically significant. Therefore, while these patterns are consistent with our hypothesis that smaller gender gaps are associated with a stronger reform effect, they should be interpreted with some caution.[28]

7.4 Religious Presence and Social Capital

Religious Presence. We next investigate whether the local presence of religious institutions represents a relevant source of heterogeneity. We rely on geolocated data obtained from the Italian Episcopal Conference (CEI). We focus on a supply-side measure of religiosity, the density of Catholic churches per 1,000 inhabitants at the municipal level, which captures both the potential religiosity of the local community and the presence of an organized social institution that may contribute to mitigating family conflict. We then construct an indicator equal to one if the municipal density of churches is above the national median.

[28]One potential explanation for why we do not detect statistically significant effects is that we use gender gaps in employment rate and in educational attainment, which are rather indirect proxies for the gender gap in income levels. Unfortunately for our purposes, data on income do not display a break-up based on gender.

The results, reported in Panel (a) of Figure 10 and in column (1) of Table B.11, show that the estimated effect of the reform is positive and statistically significant in municipalities with a below-median density of churches, while it is small and statistically indistinguishable from zero above the median, with a negative and marginally significant interaction term.

Social Capital. We complement this analysis using a broader measure of social capital. We proxy for social capital using the number of associations registered under the Italian *5 per mille* tax scheme for the 2007-2013 period, published by the Italian Revenue Agency (*Agenzia delle Entrate*).[29] We then construct an indicator equal to one if the number of local nonprofit organizations per 1,000 inhabitants is above the national median.

The results, reported in Panel (b) of Figure 10 and in column (2) of Table B.11, document a pattern that is consistent and, if anything, is sharper than the one found for religious presence: the effect of the reform is positive and statistically significant in municipalities with a below-median density of associations, while it is close to zero and statistically insignificant above the median, with a negative and statistically significant interaction term.

Taken together, the patterns presented in Figure 10 and Table B.11 are consistent with local religious and civic institutions moderating the effect of the reform, although the evidence does not allow us to distinguish among the underlying channels.

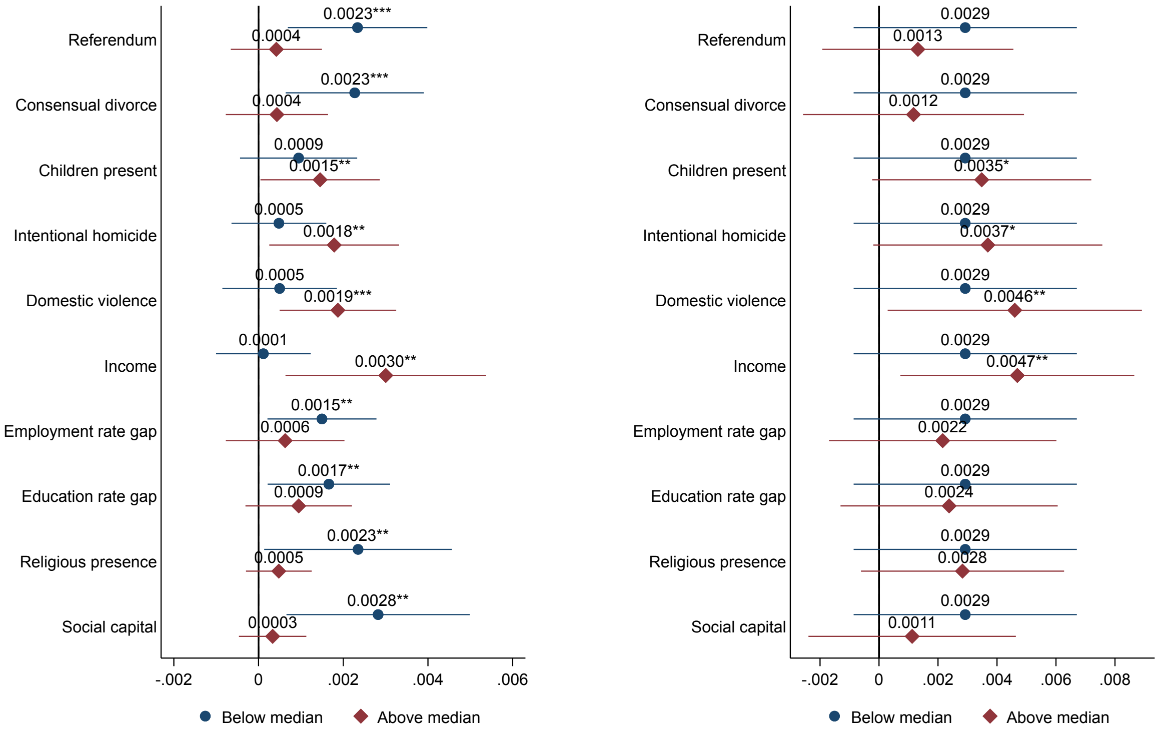
7.5 All Dimensions Jointly

The heterogeneity analyses presented in this section examine each dimension separately. Because several of these municipal characteristics may be correlated with each other – for instance, municipalities with lower income may also exhibit larger gender gaps or weaker social capital – it is useful to assess the explanatory power of each dimension of heterogeneity once all of them are accounted for simultaneously in the same regression.

Figure 11 reports the results. Panel (a) reproduces the single-dimension estimates discussed in previous sections, while the estimates from the saturated model are reported in Panel (b). Four dimensions retain statistically significant explanatory power: the municipal average income, the incidence of parents among couples that divorced before the reform, and the pre-reform incidences of intentional homicide and domestic violence, respectively. The remaining dimensions

[29] The number of associations registered under the 5 per mille scheme has been used as a proxy for social capital in related work, including Buonanno et al. (2024). This choice is also consistent with the findings of Durante et al. (2025), who show that social capital is a multidimensional construct encompassing distinct components – social participation, political participation, general trust, and institutional trust – that, while correlated, do not necessarily move together. In light of this evidence, relying on a single, well-defined measure rather than on a composite index avoids conflating conceptually distinct dimensions of social capital into a single number, and allows us to attribute any estimated effect to a specific, clearly identified channel – civic and associational participation. Moreover, our measure is observed annually and covers the universe of Italian municipalities, allowing us to exploit within-province variation over time.

Figure 11. Heterogeneity: all dimensions together



(a) Single dimensions

(b) All dimensions

Notes: Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Blue markers report the coefficient for municipalities below the national median along each dimension; red markers report the coefficient for municipalities above the median. Panel (a) estimates each heterogeneity dimension separately, replicating the specifications in Figure 7–Figure 9 one at a time. Panel (b) includes all dimensions jointly in a single regression. Dimensions are, from top to bottom: divorce attitudes (1974 referendum), the share of consensual divorces, the share of divorces with children, the rate of intentional homicides, the rate of domestic violence crimes, the mean income, the employment rate gap, the education rate gap, religious presence and social capital. Control variables are defined in Table 1.

– divorce attitudes, consensual divorces, religious presence, social capital, and the employment and education gender gaps – fail to clear conventional thresholds for statistical significance but remain comparable with the estimates obtained in the previous regressions.

8 Effect of Divorce Law Reforms on Non-Lethal Violence

Our main finding – that greater exposure to the divorce reforms increases femicide – stands in contrast with a number of results in the literature documenting that improvements in women's outside options tend to reduce intimate partner violence (e.g., Brassiolo, 2016; Stevenson and

Wolfers, 2006). This apparent tension is particularly striking because the reforms we study operate through the same broad mechanism emphasized in that literature: by lowering the costs of marital dissolution, they strengthen women's ability to exit the relationship. Yet there are at least two reasons why our findings need not be inconsistent with the existing evidence.[30]

First, much of the existing evidence on non-lethal violence relies on outcomes that are observed only when victims disclose abuse, contact the police, seek medical care, or turn to support services. These measures are therefore potentially affected by endogenous reporting. A reform that changes women's outside options and the likelihood of relationship dissolution may also change their incentives or opportunities to report violence, even in the absence of an equivalent change in underlying victimization. Femicide is much less exposed to this concern, since lethal events are considerably less likely to remain unobserved.

Second, the difference may reflect a conceptual distinction between lethal and non-lethal violence. Non-lethal domestic violence is often recurrent and may be used instrumentally to exert control, enforce compliance, or extract resources within an ongoing relationship (Aizer and Dal Bó, 2009). Femicide, instead, is disproportionately concentrated around relationship dissolution, when the perpetrator's control over the partner is threatened most acutely (Campbell et al., 2003; Wilson and Daly, 1993). Easier divorce may therefore reduce some forms of violence within continuing relationships while at the same time increasing the number of couples exposed to the particularly dangerous period surrounding separation.

These two explanations are not mutually exclusive. In what follows, we first ask whether the same reforms that increase femicide also reduce observed non-lethal violence in Italy, consistent with the existing literature. This within-setting comparison holds fixed the institutional reform and empirical design, while varying the type of violence considered. We then develop a conceptual framework that formalizes why lethal and non-lethal violence may respond in opposite directions to the same change in divorce costs.

As discussed in Section 7, administrative data on non-lethal forms of violence come directly from investigative reports but are not available at the municipality level. For this reason, the empirical analysis that follows is performed at the province-year level. Consequently, we recon-

[30]In our data, femicides do not co-move with more conventional measures of violence against women. Figure A.12 and Figure A.13 in the Appendix plot the correlation between femicide and a range of reported indicators of gender-based violence. Femicides are negatively correlated with several of these measures, including overall reports, mistreatment, stalking, and calls to anti-violence services, while the correlations with sexual assaults, judicial warnings, and restraining orders are small and statistically insignificant. Reported beatings are the main exception, displaying a positive correlation with femicide. Although these correlations are purely descriptive, they already suggest that lethal and non-lethal violence need not move together.

Table 4. Alternative measures of violence

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent variable =	Beatings	Threats	Offences	All DV crimes	Rapes	Intentional homicides	Femicides
Post-reform × Instrument (std.)	-1.5952	-10.8477*	-14.6714***	-29.5722***	-0.1100	0.0478***	0.0401**
	(1.0294)	(6.4315)	(4.2552)	(7.2797)	(0.2418)	(0.0132)	(0.0155)
Obs	1593	1593	1169	1169	1593	1605	1605
Provinces	107	107	107	107	107	107	107
Year and Province FE	✓	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓	✓	✓	✓	✓
Mean dep. var pre-reform	27.066	152.010	120.642	301.354	7.549	0.207	0.222
R-squared	0.672	0.857	0.914	0.722	0.580	0.167	0.126

Notes: OLS reduced-form estimates with province and year fixed effects. Standard errors (in parentheses) are clustered at the province level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. Columns (1)–(3) and (5) use as dependent variable the number of reported crimes per 100,000 inhabitants for, respectively, beatings, threats, and offenses, none of which are disaggregated by victim's sex at the province level. Column (4) combines these three offenses into a single domestic violence index (beatings, threats, and offenses summed together), while rapes are excluded from the index and used instead as a placebo outcome. Because offense data are only available through 2017, both column (3) and the domestic violence index in column (4) are restricted to this shorter sample period; all other outcomes span the full sample. Column (6) uses intentional homicides of women per 100,000 inhabitants, drawn from the ISTAT Cause of Death registry, and column (7) uses our hand-collected measure of femicides per 100,000 inhabitants. Control variables are described in Table 1 and measured at the province level.

struct our shift-share exposure variable for each province p as follows:

$$d_p = \sum_{a=1}^{A} \frac{Pop_{a,p,2011}}{Pop_{p,2011}} \times \frac{D_{a,2011}}{Pop_{a,2011}}, \tag{7}$$

and we estimate the reduced-form Difference-in-Differences model

$$F_{p,t} = \beta_{RF,p} * (d_p \times PostReform_t) + X_{p,2011} \sum_t \lambda_t + \eta_p + \theta_t + \varepsilon_{p,t}. \tag{8}$$

Table 4 reports the results from estimating Equation 8 on the incidence per 100,000 inhabitants at the province level of a number of crimes that are typically associated with violence against women – beatings, threats, offenses, and rapes – together with a composite index that aggregates reports of beatings, threats, and offenses reports.[31] We include the incidence of rape as a placebo outcome because it is a form of gendered violence that is less directly tied to the domestic setting affected by divorce law. To ensure that our headline results holds when the data are aggregated to the province level, we also estimate regressions in which the dependent variables measure femicides (from our manually-collected data) and intentional homicides of women (from administrative data).

The results reveal a clear divergence between non-lethal and lethal violence. The estimated effects on beatings, threats, and offenses are negative, with statistically significant declines in threats, offenses, and in the composite index of domestic violence. By contrast, the coefficient on rape is small and statistically indistinguishable from zero. Columns (6) and (7) show

[31]These province-level offenses are not disaggregated by victim gender and should therefore be interpreted as complementary evidence on non-lethal domestic violence rather than as a direct measure of IPV against women.

that the positive effect of the reforms on intentional homicides and femicides also survives aggregation to the province-year level, with both coefficients remaining positive and statistically significant.[32]

9 Conceptual framework

The empirical results above reveal a striking divergence: greater exposure to the divorce reforms is associated with less observed non-lethal domestic violence but more femicide. Reporting differences may account for part of this pattern, as discussed above. In this section, we develop a complementary explanation based on the idea that lethal and non-lethal violence arise at different stages of a relationship and can therefore respond differently to easier marital dissolution.

The intuition is rather simple. Easier separation changes behavior both within ongoing relationships and at the transition to separation. A more credible threat of exit can induce husbands to moderate violence to preserve the relationship – though the model also allows the opposite response, since sufficiently high violence can instead deter separation – and it draws some couples that would otherwise have stayed together into the separation process, where violence may rise or fall depending on husband type. Separation is itself a dangerous transition, however: a threatened loss of control over the relationship can trigger retaliatory escalation, which the model interprets as femicide. Easier divorce can therefore move non-lethal and lethal violence through distinct margins, allowing the two to diverge.

We formalize this with a model of instrumental marital violence and separation, solved by backward induction across three margins: (i) the husband chooses violence within the relationship, trading off greater control against its cost; (ii) the wife, having observed violence, decides whether to stay or attempt separation, with lower divorce costs making separation attractive over a wider range of violence; and (iii) once separation is attempted, the husband either accepts it or escalates – our formalization of femicide. Sorting couples by their equilibrium response to the reform yields sharp predictions, summarized in Propositions 1–3 below.

9.1 Timing and payoffs

There are two agents, a husband H and a wife W, generating relationship surplus $S > 0$. The husband exercises non-lethal violence $v \in [0, \overline{v}]$ while the relationship continues, consistent with an instrumental view of intimate partner violence as a tool for extracting resources or control

[32] The positive and statistically significant province-level estimate also helps rule out the possibility that our main result is driven by idiosyncrasies associated with measuring a rare outcome at a highly disaggregated geographical level.

(Aizer and Dal Bó, 2009; Bloch and Rao, 2002; Bobonis et al., 2013). The game has three stages: (1) the husband chooses v; (2) the wife observes v and decides whether to stay or attempt separation; (3) if she attempts separation, a shock $\varepsilon \sim U[-\overline{\varepsilon}, \overline{\varepsilon}]$ realizes and the husband, observing it, either accepts separation or escalates.[33]

If the wife stays, violence shifts the surplus division toward the husband, who receives share $\mu(v)$ ($\mu' > 0$, $\mu'' < 0$) at a type-specific cost $\theta c(v)$ (c increasing, convex, $c(0) = 0$). The type $\theta \in [\underline{\theta}, \overline{\theta}]$ is common knowledge and captures heterogeneity in the cost of sustaining violence as a control strategy – not in how the husband reacts once separation is already under way. Stay payoffs are

$$W^S(v) = (1 - \mu(v))S, \qquad H^S(v, \theta) = \mu(v)S - \theta c(v).$$

If the wife attempts separation and the husband accepts, only a fraction $\delta \in (0, 1)$ of S survives, divided by a default rule giving the husband share $\alpha \in (0, \frac{1}{2})$; both parties also bear a divorce cost κ – the primitive our reforms directly affect – and the husband additionally pays $\chi c(v)$, a cost tied to his history of violence that becomes salient once the relationship dissolves (legal, reputational, or economic):

$$W^A(\kappa) = (1 - \alpha)\delta S - \kappa, \qquad H^A(v, \kappa) = \alpha\delta S - \kappa - \chi c(v).$$

If instead the husband escalates – our formalization of femicide – the wife's payoff is normalized to $-M$, and the husband is detected with probability q: if undetected he keeps δS, while if detected he loses the surplus and pays a fixed sanction Ψ together with $\chi c(v)$. A shock ε, realized only after the wife's decision, captures idiosyncratic circumstances affecting the husband's relative propensity to escalate, so the wife evaluates separation through the *probability* it will be accepted rather than met with escalation. This margin formalizes the backlash channel triggered by the wife's separation attempt, consistent with evidence that shifts in bargaining power can provoke retaliation rather than only discipline violence (Bergvall, 2024; Bhalotra, Brito et al., 2021; Bhalotra, Kambhampati et al., 2021; González and Rodríguez-Planas, 2020; Guarnieri and Rainer, 2021; Heath, 2014).

A key asymmetry drives the model's central mechanism: the violence-related cost $\chi c(v)$ is borne with certainty if separation is accepted but only if the husband is detected should he escalate instead. This captures, in reduced form, an asymmetry in accountability for prior violence. If separation is accepted, the husband bears the consequences of his past violence with certainty. If he instead escalates lethally, those consequences are borne only if he is detected; when undetected, he avoids them altogether. A more violent history therefore makes accepting

[33] "Acceptance" is behavioral, not legal: it means the husband does not escalate in response to the separation attempt.

the wife's departure relatively less attractive and raises the risk of further escalation, linking violence during the relationship to lethal risk at the moment of separation.

The wife anticipates this and evaluates separation through

$$W^L(v,\kappa) = p(v,\kappa)W^A(\kappa) + (1-p(v,\kappa))W^E,$$

where $p(v,\kappa)$ is the equilibrium acceptance probability. She thus faces a trade-off: more violence makes staying less attractive but also makes leaving more dangerous, raising the husband's incentive to escalate. Divorce costs affect both sides at once – lowering κ makes separation directly less costly for the wife and makes acceptance relatively more attractive for the husband.

9.2 Equilibrium

Solving by backward induction, the probability of acceptance $p(v,\kappa)$ is strictly decreasing in both v and κ (Appendix C.3).

The wife's separation decision. Let $G(v,\kappa) := W^S(v) - W^L(v,\kappa)$; the wife stays when $G(v,\kappa) > 0$ and attempts separation when $G(v,\kappa) < 0$. This need not be monotone in v: staying dominates at low violence, but as violence rises the loss of relationship surplus can make separation preferable, while at sufficiently high violence the asymmetry above makes leaving increasingly dangerous, since the husband becomes relatively more likely to escalate than to accept. Together these forces make $G(v,\kappa)$ U-shaped, with $G(0,\kappa) > 0$.

Whenever a nondegenerate separation range is supported, there is a strictly positive lower threshold $v_1(\kappa)$. If staying is again optimal at sufficiently high violence, there is also an upper threshold $v_2(\kappa)$ and the separation range is

$$[v_1(\kappa), v_2(\kappa)].$$

Otherwise, the separation range extends to the upper boundary,

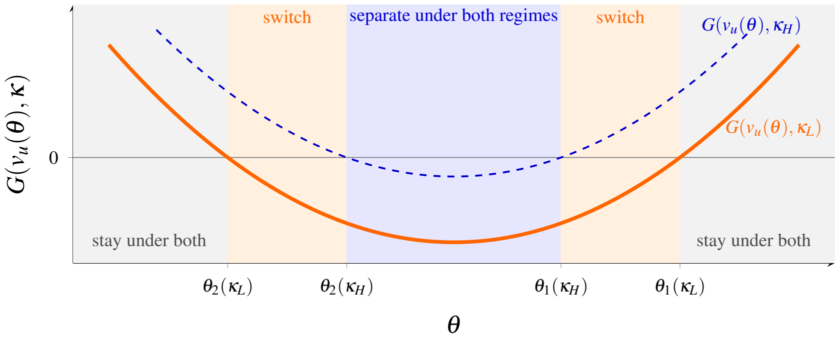
$$[v_1(\kappa), \bar{v}].$$

The Appendix provides the complete characterization, including the knife-edge cases (Appendix C).

The husband's violence choice. Let $v^*(\theta;\kappa)$ denote type θ's equilibrium choice. Absent the separation constraint, his preferred violence solves

$$S\mu'(v_u(\theta)) = \theta c'(v_u(\theta)),$$

Figure 12. Wife's separation decision in type space



Notes: The figure plots the wife's payoff difference evaluated at the husband's unconstrained violence choice, $G(v_u(\theta), \kappa)$. The two curves are drawn in parallel for simplicity, so as to illustrate only that lower divorce costs shift $G(v_u(\theta), \kappa)$ downward for every husband type. The thresholds $\theta_1(\kappa)$ and $\theta_2(\kappa)$ are defined implicitly by $v_u(\theta_1(\kappa)) = v_1(\kappa)$ and $v_u(\theta_2(\kappa)) = v_2(\kappa)$. Hence the blue region corresponds to husband types for which the wife's best response to the unconstrained violence choice is attempted separation under both regimes, while the orange regions correspond to types for which attempted separation is a best response only under κ_L. This is not yet an equilibrium separation figure: it shows where the husband's unconstrained violence choice is incompatible with staying, before he optimally adjusts violence or instead chooses to induce separation.

with $v_u(\theta)$ strictly decreasing in θ. Hence husband type maps one-to-one into an unconstrained violence choice. If $v_u(\theta)$ already induces the wife to stay, the husband can implement it directly. Otherwise, a husband who wishes to preserve the relationship must adjust violence, while other husbands instead induce separation. Whenever separation is induced, the husband chooses the lowest separation-inducing violence level, $v_1(\kappa)$, because his continuation payoff following a separation attempt is decreasing in prior violence (Appendix C.5).

9.3 The effect of the divorce reform

Proposition 1 (Effect of lower divorce costs on separation)**.** *Suppose divorce costs fall from κ_H to $\kappa_L < \kappa_H$. Then the range of violence levels at which attempted separation is a best response weakly expands, while the range at which staying is a best response weakly contracts. Equivalently, the range of husband types whose unconstrained violence choice would induce attempted separation weakly expands. In equilibrium, every husband type that induces separation under κ_H also induces separation under κ_L.*

Evaluating the wife's payoff difference at the husband's unconstrained choice provides a convenient representation in husband-type space. In the bounded case, define $\theta_1(\kappa)$ and $\theta_2(\kappa)$ by $v_u(\theta_1(\kappa)) = v_1(\kappa)$ and $v_u(\theta_2(\kappa)) = v_2(\kappa)$. Because $v_u(\theta)$ is strictly decreasing, $\theta_2(\kappa) < \theta_1(\kappa)$, and

$$G(v_u(\theta), \kappa) < 0 \quad \iff \quad \theta_2(\kappa) < \theta < \theta_1(\kappa).$$

Thus these thresholds identify the husband types whose unconstrained violence choice would induce the wife to attempt separation. Figure 12 represents the first two implications of Proposition 1 directly in husband-type space. The figure plots $G(v_u(\theta), \kappa)$, the wife's payoff difference evaluated at the violence level the husband would choose absent the separation constraint. In the bounded case illustrated, the two violence thresholds under each regime map one-to-one into two type-space thresholds. Lower divorce costs therefore expand the range of husband types whose unconstrained violence choice would induce attempted separation:

$$[\theta_2(\kappa_H), \theta_1(\kappa_H)] \subset [\theta_2(\kappa_L), \theta_1(\kappa_L)].$$

These four thresholds should not be confused with the equilibrium type cutoff introduced next. The thresholds $\theta_1(\kappa)$ and $\theta_2(\kappa)$ describe whether the husband's *unconstrained* violence choice is compatible with staying. The equilibrium cutoff instead accounts for the husband's ability to adjust violence when his unconstrained choice would trigger separation.

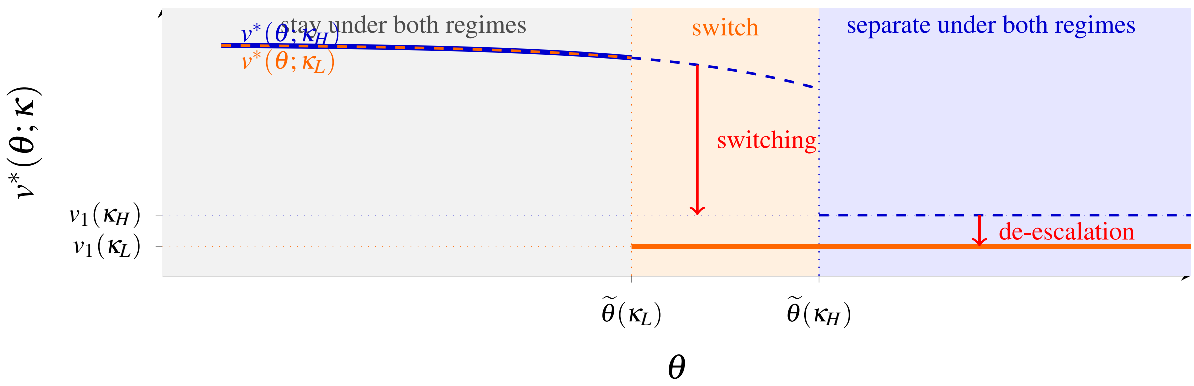
To characterize equilibrium relationship outcomes, compare the husband's best payoff from preserving the relationship with his payoff from inducing separation. Because higher-θ husbands find violence more costly while their continuation payoff following a separation attempt does not depend on θ, this comparison satisfies single crossing. Hence there is at most one interior cutoff $\tilde{\theta}(\kappa)$ such that types below the cutoff preserve the relationship and types above it induce attempted separation. When both regimes have interior cutoffs, lower divorce costs shift the equilibrium cutoff to the left, i.e., $\tilde{\theta}(\kappa_L) < \tilde{\theta}(\kappa_H)$. The type space is therefore divided into three equilibrium regions: types below $\tilde{\theta}(\kappa_L)$ stay under both regimes; types between the two cutoffs switch from staying under κ_H to attempted separation under κ_L; and types above $\tilde{\theta}(\kappa_H)$ induce attempted separation under both regimes.

Proposition 2 (Effect of lower divorce costs on equilibrium violence)**.** *Suppose divorce costs fall from κ_H to $\kappa_L < \kappa_H$.*

1. *Among couples that enter the separation process under both regimes, equilibrium violence strictly falls.*

2. *Among couples that remain in the relationship under both regimes, violence is unchanged if the husband's unconstrained choice remains compatible with staying under the low-cost regime. If the reform makes this choice incompatible with staying, violence strictly falls when no high-violence stay region exists under the low-cost regime; if such a region exists, a constrained husband may instead either reduce violence to preserve the relationship or increase it enough to deter separation.*

Figure 13. Equilibrium non-lethal violence and the effect of lower divorce costs



Notes: Configuration with a bounded separation range under both regimes and unconstrained choices compatible with staying. Gray: violence is unchanged. Orange (switch region, $\theta < \widehat{\theta}$ here): violence falls; the arrow isolates the switching component, from the high-cost stay choice to $v_1(\kappa_H)$, while the remaining decline from $v_1(\kappa_H)$ to $v_1(\kappa_L)$ is the de-escalation component. Blue: violence strictly falls from $v_1(\kappa_H)$ to $v_1(\kappa_L)$. Table C.14 in the Appendix covers the remaining configurations.

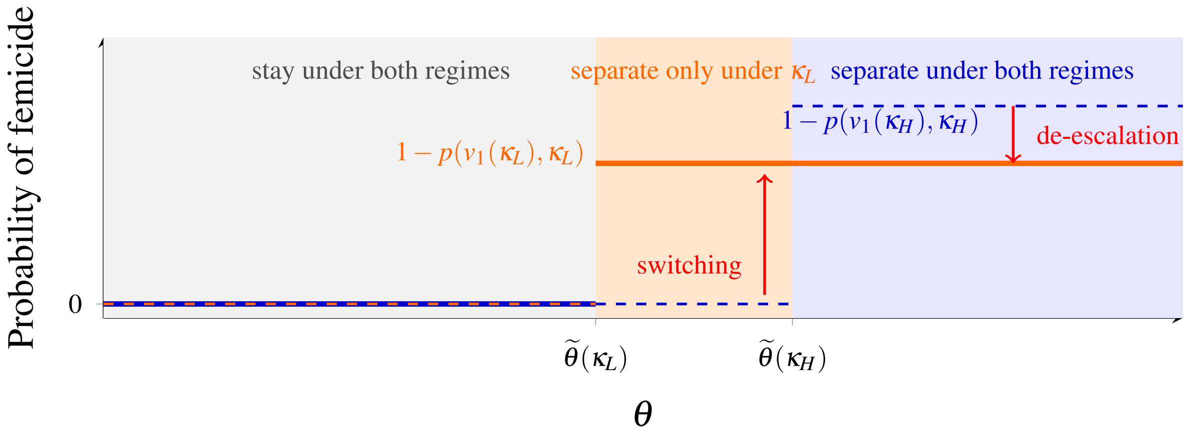
3. *Among couples that remain together under κ_H but enter the separation process under κ_L, equilibrium violence falls if $\theta < \widehat{\theta}$, is unchanged if $\theta = \widehat{\theta}$, and rises if $\theta > \widehat{\theta}$.*

Figure 13 illustrates Proposition 2 for a configuration in which all types in the switch region satisfy $\theta < \widehat{\theta}$: violence is unchanged among couples staying under both regimes, falls among couples induced to separate, and strictly falls among couples separating under both regimes.

Proposition 3 (Effect of lower divorce costs on femicide risk)**.** *Suppose divorce costs fall from κ_H to $\kappa_L < \kappa_H$. Couples that remain in the relationship under both regimes face zero equilibrium femicide risk under both regimes. Couples that remain in the relationship under κ_H but enter the separation process under κ_L move from zero to a strictly positive risk of femicide. Among couples that enter the separation process under both regimes, equilibrium femicide risk strictly falls.*

Figure 14 illustrates Proposition 3. The switching and de-escalation effects have opposite signs, so the reform's aggregate effect on femicide risk depends on the distribution of husband types across the two regions and on the magnitude of their opposing individual responses.

Figure 14. Equilibrium lethal violence and the effect of lower divorce costs



Notes: Gray: femicide risk is zero under both regimes. Orange (switching effect): risk rises from 0 to $1-p(v_1(\kappa_L),\kappa_L) > 0$. Blue (de-escalation effect): risk falls from $1-p(v_1(\kappa_H),\kappa_H)$ to $1-p(v_1(\kappa_L),\kappa_L)$. Unlike Figure 13, these signs do not depend on any additional type threshold.

10 Conclusions

This paper studies how easier access to divorce affects violence against women. We exploit two reforms introduced in Italy in 2014 and 2015 that substantially reduced the monetary, administrative, and time costs of marital dissolution. Combining administrative data on divorces with a newly assembled dataset on femicides, we exploit pre-existing differences in local exposure to the reforms arising from municipalities' age composition interacted with national age-specific divorce rates. We find that greater exposure to the reforms increases the probability and number of femicides. The effect is concentrated within the domestic sphere and is particularly pronounced for killings perpetrated by ex-partners and relatives. At the same time, we find that the reforms are associated with reductions in several measures of reported offenses commonly associated with non-lethal domestic violence in more exposed provinces, consistent with earlier results in the literature.

Taken together, these results show that lethal and non-lethal violence need not respond in the same way to an improvement in women's outside options. Differences in reporting may account for part of this divergence, since non-lethal violence is observed only when episodes come to the attention of institutions, and reporting behavior may be affected by the reform too. We formalize in a conceptual framework a complementary substantive explanation. Making separation easier can discipline violence within some continuing relationships and allow some women to leave abusive relationships in which they would otherwise have remained. Conditional on a separation attempt, lower divorce costs can also make peaceful dissolution more attractive and reduce the risk of lethal escalation. At the same time, however, the reform induces additional couples to enter the separation process, exposing women who would otherwise have remained in the relationship to a positive risk of lethal violence. The increase in femicide can therefore arise even if separation becomes less dangerous conditional on being attempted.

Our findings show that lowering barriers to marital dissolution can affect different margins of intimate partner violence in opposite directions. They do not imply that barriers to divorce should be preserved: reducing the cost of marital dissolution expands women's outside options and can reduce non-lethal violence. Rather, the results suggest that policies facilitating separation may be particularly effective when accompanied by measures that address the heightened risk surrounding relationship dissolution. More broadly, our findings emphasize that reforms affecting intra-household bargaining and exit options can generate different effects at different stages of a relationship, and that evaluating such reforms requires considering both violence within ongoing relationships and the risks associated with their dissolution.

References

Aizer, A. (2010). The gender wage gap and domestic violence. *American Economic Review 100*(4), 1847–1859.

Aizer, A. (2011). Poverty, Violence, and Health: The Impact of Domestic Violence During Pregnancy on Newborn Health. *The Journal of Human Resources 46*(3), 518–538.

Aizer, A. and P. Dal Bó (2009). Love, hate and murder: Commitment devices in violent relationships. *Journal of Public Economics 93*(3-4), 412–428.

Anderson, S. and G. Genicot (2015). Suicide and property rights in India. *Journal of Development Economics 114*, 64–78.

Angelucci, M. and R. Heath (2020). Women empowerment programs and intimate partner violence. *AEA Papers and Proceedings 110*, 610–614.

Asik, G. A. and E. N. Ozen (2021). It takes a curfew: The effect of covid-19 on female homicides. *Economics Letters 200*, 109761.

Bergvall, S. (2024). Women's economic empowerment and intimate partner violence. *Journal of Public Economics 239*, 105211.

Bhalotra, S., E. Brito, D. Clarke, P. Larroulet, F. Pino, et al. (2021). Dynamic impacts of lockdown on domestic violence: Evidence from multiple policy shifts in chile.

Bhalotra, S., U. Kambhampati, S. Rawlings, and Z. Siddique (2021). Intimate partner violence: The influence of job opportunities for men and women. *The World Bank Economic Review 35*(2), 461–479.

Bloch, F. and V. Rao (2002). Terror as a bargaining instrument: A case study of dowry violence in rural India. *American Economic Review 92*(4), 1029–1043.

Bobonis, G. J., M. González-Brenes, and R. Castro (2013). Public transfers and domestic violence: The roles of private information and spousal control. *American Economic Journal: Economic Policy 5*(1), 179–205.

Borusyak, K., P. Hull, and X. Jaravel (2022). Quasi-experimental shift-share research designs. *The Review of Economic Studies 89*(1), 181–213.

Borusyak, K., P. Hull, and X. Jaravel (2025). A practical guide to shift-share instruments. *Journal of Economic Perspectives 39*(1), 181–204.

Brassiolo, P. (2016). Domestic violence and divorce law: When divorce threats become credible. *Journal of Labor Economics 34*(2), 443–477.

Buonanno, P., I. Ferrari, and A. Saia (2024). All is not lost: Organized crime and social capital formation. *Journal of Public Economics 240*, 105257.

Buonanno, P., P. Giacobbe, and A. Mosca (2026). Voting after gendered shocks: The impact of femicides on electoral preferences.

Buonanno, P., A. Mosca, and P. Giacobbe (2025). The political economy of femicides: Local spending responses, electoral incentives and nationality bias.

Calabresi, E. (2024). Don't Leave Me Alone: Unilateral Divorce and Intimate Partner Violence. Technical report, Universitá degli Studi di Firenze.

Calabresi, E. and N. Rodríguez-Planas (2025). Female empowerment and intimate partner violence. *IZA Discussion Paper*.

Campbell, J. C. (2002). Health consequences of intimate partner violence. *The Lancet 359*(9314), 1331–1336.

Campbell, J. C., D. Webster, J. Koziol-McLain, C. Block, D. Campbell, M. A. Curry, et al. (2003). Risk factors for femicide in abusive relationships: Results from a multisite case control study. *American Journal of Public Health 93*(7), 1089–1097.

Carrell, S. E. and M. L. Hoekstra (2010). Externalities in the classroom: How children exposed to domestic violence affect everyone's kids. *American Economic Journal: Applied Economics 2*(1), 211–228.

Cerqua, A., C. Giannantoni, M. Letta, and G. Pinto (2026). Femicides, anti-violence centers, and policy targeting. *European Economic Review*, 105221.

Chen, J. and J. Roth (2024). Logs with zeros? some problems and solutions. *The Quarterly Journal of Economics 139*(2), 891–936.

Chiappori, P.-A., B. Fortin, and G. Lacroix (2002). Marriage market, divorce legislation, and household labor supply. *Journal of Political Economy 110*(1), 37–72.

Colagrossi, M., C. Deiana, D. Dragone, A. Geraci, L. Giua, and E. Iori (2023). Intimate partner violence and help-seeking: The role of femicide news. *Journal of Health Economics 87*, 102722.

Corradini, V. and G. Buccione (2023). Unilateral divorce rights, domestic violence and women's agency: Evidence from the Egyptian Khul reform. *Journal of Development Economics 160*, 102947.

Dee, T. S. (2003). Until death do you part: The effects of unilateral divorce on spousal homicides. *Economic Inquiry 41*(1), 163–182.

Durante, R., N. Mastrorocco, L. Minale, and J. M. Snyder (2025). Unpacking social capital. *The Economic Journal 135*(667), 773–807.

Duvvury, N., A. Vara-Horna, C. Brendel, and M. Chadha (2023). Productivity impacts of intimate partner violence: Evidence from Africa and South America. *SAGE Open 13*(4), 21582440231205524.

EIGE (2023). The costs of gender-based violence in the European Union.

Einiö, E., N. Metsä-Simola, M. Aaltonen, E. Hiltunen, and P. Martikainen (2023). Partner violence surrounding divorce: A record-linkage study of wives and their husbands. *Journal of Marriage and Family 85*(1), 33–54.

European Institute for Gender Equality (EIGE) (2024). Gender equality index 2024: Tackling violence against women, tackling gender inequalities. Technical report, European Institute for Gender Equality.

Fearon, J. and A. Hoeffler (2014). Benefits and costs of the conflict and violence targets for the post-2015 development agenda. *Conflict and violence assessment paper, Copenhagen Consensus Center*, 1–65.

Fernández-Kranz, D., N. Nollenberger, and J. Roff (2026). Bargaining Under Threats: The Effect of Joint Custody Laws on Intimate Partner Violence. *American Journal of Health Economics 12*(1), 118–163.

Friedberg, L. (1998). Did unilateral divorce raise divorce rates? Evidence from panel data.

Frisancho, V., E. Pappa, and C. Santantonio (2022). When women win: Can female representation decrease gender-based violence?

García-Ramos, A. (2021). Divorce laws and intimate partner violence: Evidence from Mexico. *Journal of Development Economics 150*, 102623.

González, L. and N. Rodríguez-Planas (2020). Gender norms and intimate partner violence. *Journal of Economic Behavior & Organization 178*, 223–248.

Gruber, J. (2004). Is making divorce easier bad for children? The long-run implications of unilateral divorce. *Journal of Labor Economics 22*(4), 799–833.

Guarnieri, E. and H. Rainer (2021). Colonialism and female empowerment: A two-sided legacy. *Journal of Development Economics 151*, 102666.

Gutierrez, I. A. and O. Molina (2021). Does domestic violence jeopardize the learning environment of peers within the school? Peer effects of exposure to domestic violence in urban Peru. *Economics of Education Review 84*.

Gutiérrez-Romero, R. (2024). Femicide laws, unilateral divorce, and abortion decriminalization fail to stop women from being killed in mexico. *arXiv preprint arXiv:2407.06722*.

Haushofer, J., C. Ringdal, J. P. Shapiro, and X. Y. Wang (2019). Income changes and intimate partner violence: Evidence from unconditional cash transfers in kenya. Technical report, National Bureau of Economic Research.

Heath, R. (2014). Women's access to labor market opportunities, control of household resources, and domestic violence: Evidence from Bangladesh. *World Development 57*, 32–46.

Hidrobo, M. and L. Fernald (2013). Cash transfers and domestic violence. *Journal of Health Economics 32*(1), 304–319.

Hoehn-Velasco, L. and A. Silverio-Murillo (2020). Do spouses negotiate in the shadow of the law? Evidence from unilateral divorce, suicides, and homicides in Mexico. *Economics Letters 187*, 108891.

ISTAT (2022). Matrimoni, unioni civili, separazioni e divorzi. Available at https://www.istat.it/it/archivio/292309, (accessed on 19 April 2024).

ISTAT (2025a). La violenza contro le donne, dentro e fuori la famiglia. https://www.istat.it/comunicato-stampa/la-violenza-contro-le-donne-dentro-e-fuori-la-famiglia-primi-risultati-anno-2025/, (accessed on 19 January 2026.

ISTAT (2025b). Le vittime di omicidio - Anno 2024. https://www.istat.it/comunicato-stampa/le-vittime-di-omicidio-anno-2024/, (accessed on 19 January 2026.

Lloyd, S. (1997). The effects of domestic violence on women's employment. *Law & Policy 19*(2), 139–167.

Molina, S., L. Wagner, and M. Kreyenfeld (2025). Women's economic independence and physical intimate partner violence (ipv) during separation. *Plos one 20*(6), e0326529.

Monckton Smith, J. (2020). Intimate partner femicide: Using Foucauldian analysis to track an eight stage progression to homicide. *Violence Against Women 26*(11), 1267–1285.

Perova, E., S. Reynolds, and I. Schmutte (2021). Does the Gender Wage Gap Influence Intimate Partner Violence in Brazil. *Evidence from Administrative Health Data. World Bank Policy Research Working Paper WPS9656*.

Peterson, C., M. C. Kearns, W. L. McIntosh, L. F. Estefan, C. Nicolaidis, K. E. McCollister, A. Gordon, and C. Florence (2018). Lifetime economic burden of intimate partner violence among us adults. *American Journal of Preventive Medicine 55*(4), 433–444.

Pisanelli, E. (2025). Divorce, domestic violence, and help seeking. *Italian Economic Journal 11*(3), 1091–1116.

Pollak, R. A. (2004). An intergenerational model of domestic violence. *Journal of Population Economics 17*(2), 311–329.

Shah, M. and L. Barski (2025). Intimate Partner Violence in Low- and Middle-Income Countries: Insights from Economic Research. NBER Working Paper 34337, National Bureau of Economic Research.

Silva, J. S. and S. Tenreyro (2006). The log of gravity. *The Review of Economics and Statistics*, 641–658.

Silverio-Murillo, A. (2019). Is it true that unilateral divorce improves the empowerment of women? Unpublished manuscript.

Spencer, C. M. and S. M. Stith (2020). Risk factors for male perpetration and female victimization of intimate partner homicide: A meta-analysis. *Trauma, Violence, & Abuse 21*(3), 527–540.

Stevenson, B. (2008). Divorce law and women's labor supply. *Journal of Empirical Legal Studies 5*(4), 853–873.

Stevenson, B. and J. Wolfers (2006). Bargaining in the shadow of the law: Divorce laws and family distress. *The Quarterly Journal of Economics 121*(1), 267–288.

Stock, J. H. and M. Yogo (2005). Testing for weak instruments in linear iv regression. In D. W. K. Andrews and J. H. Stock (Eds.), *Identification and Inference for Econometric Models: Essays in Honor of Thomas Rothenberg*, pp. 80–108. Cambridge University Press.

Tripi, D., A. Ghamlouch, F. Del Duca, M. Racciatti, P. Frati, S. Ferracuti, and A. Maiese (2026). Intimate partner femicide, understanding perpetrator behaviour: A cross-cultural analytical review of weapon choice, injury patterns, and cultural-criminological insights. *International Journal of Legal Medicine*.

UN Women and UNODC (2025). Femicides in 2024: Global estimates of intimate partner/family member femicides. Available at https://www.unwomen.org/sites/default/files/2025-11/femicides-in-2024-global-estimates-of-intimate-partner-family-member-femicides-en.pdf, (accessed on 11 August 2026).

United Nations (2012). Vienna declaration on femicide. Symposium on Femicide, Vienna.

UNODC and UN Women (2022). Statistical framework for measuring the gender-related killing of women and girls (also referred to as "femicide/feminicide"). Technical report, United Nations Office on Drugs and Crime and UN Women. Approved by the United Nations Statistical Commission, March 2022.

Voena, A. (2015). Yours, mine, and ours: Do divorce laws affect the intertemporal behavior of married couples? *American Economic Review 105*(8), 2295–2332.

Wilson, M. and M. Daly (1993). Spousal homicide risk and estrangement. *Violence and Victims 8*, 3–3.

Wilson, M. I. and M. Daly (1992). Who Kills Whom in Spouse Killings: On the Exceptional Sex Ratio of Spousal Homicides in the United States. *Criminology 30*(2), 189–215.

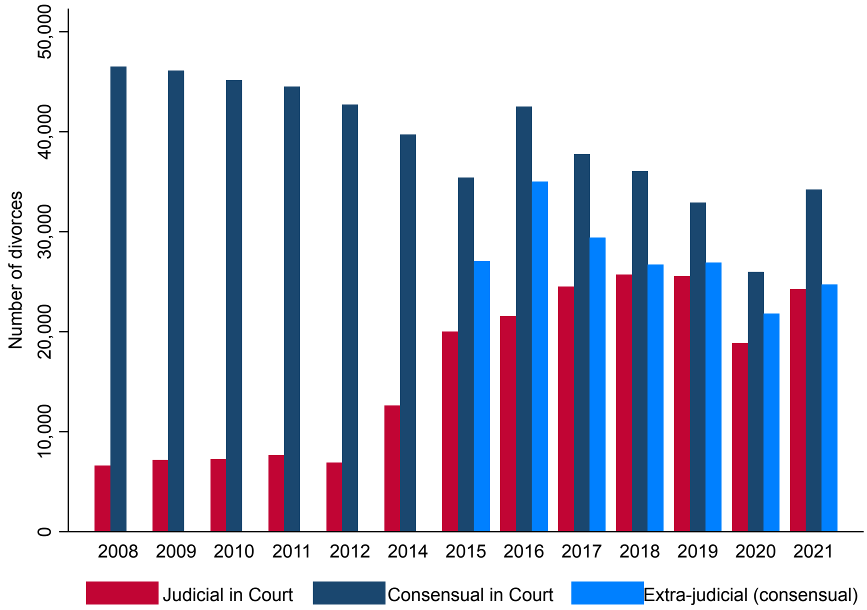
Wolfers, J. (2006). Did unilateral divorce laws raise divorce rates? A reconciliation and new results. *American Economic Review 96*(5), 1802–1820.

World Health Organization (2021). *Violence against women prevalence estimates, 2018: global, regional and national prevalence estimates for intimate partner violence against women and global and regional prevalence estimates for non-partner sexual violence against women*. World Health Organization.

Appendix

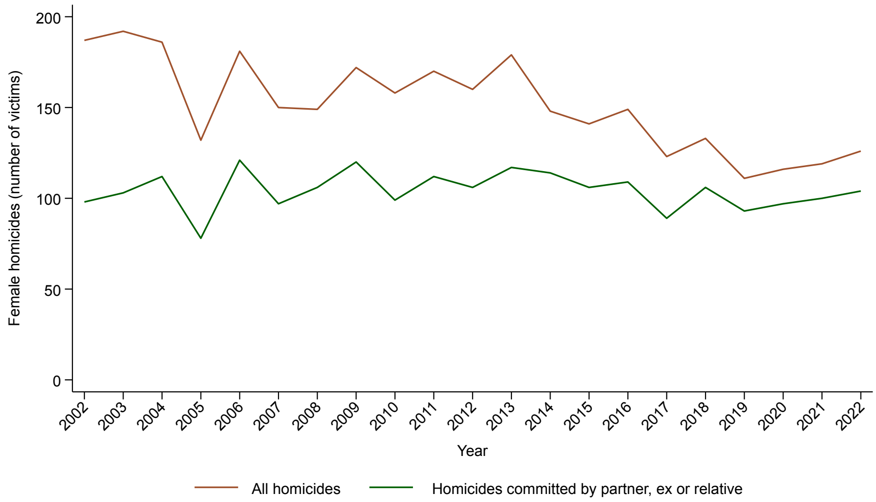
A Appendix Figures

Figure A.1. Consensual and judicial divorces



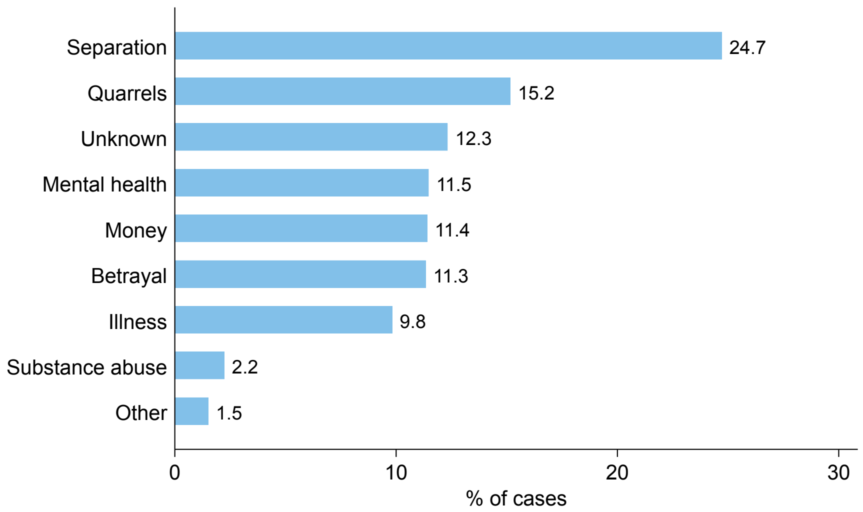
Notes: The figure plots the annual number of divorces by procedure. *Judicial in Court* refers to judicial divorces adjudicated by a court (red bars). *Consensual in Court* refers to mutually agreed divorces finalized before a court (dark blue bars). *Extra-judicial (consensual)* refers to mutually agreed divorces formalized outside the courtroom, either through lawyer-assisted negotiation or before a municipal civil registrar (light blue bars); this procedure was introduced by the December 2014 reform and is therefore observed only from 2015 onward. By construction, all divorces recorded through the municipal civil registrar are consensual, since judicial divorces are adjudicated exclusively by courts. Data for 2013 are unavailable.

Figure A.2. Female victims of homicide

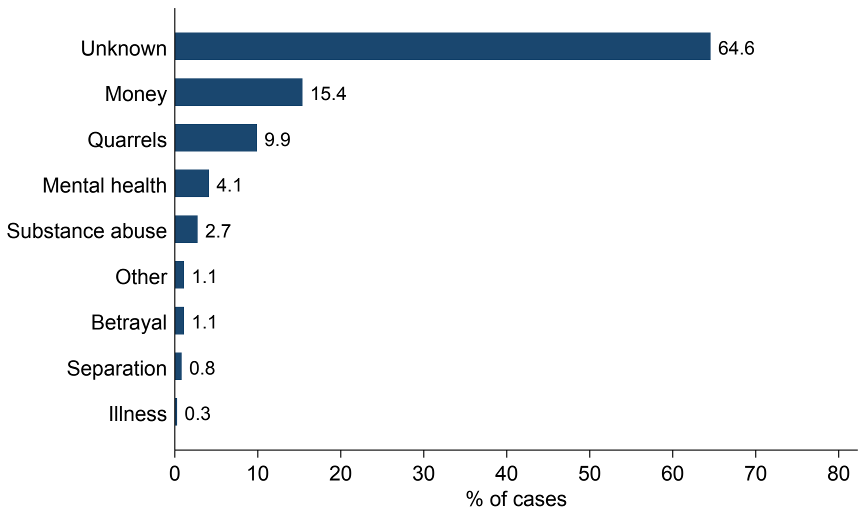


Notes: The figure illustrates the temporal trends of homicides in Italy over the period 2002-2022 according to the data of the Ministry of the Interior and of ISTAT. The orange line depicts the number of female victims reported by the Ministry of the Interior, while the green line represents the number of victims killed by a partner, an ex partner or a relative.

Figure A.3. Context of murder by type of killer



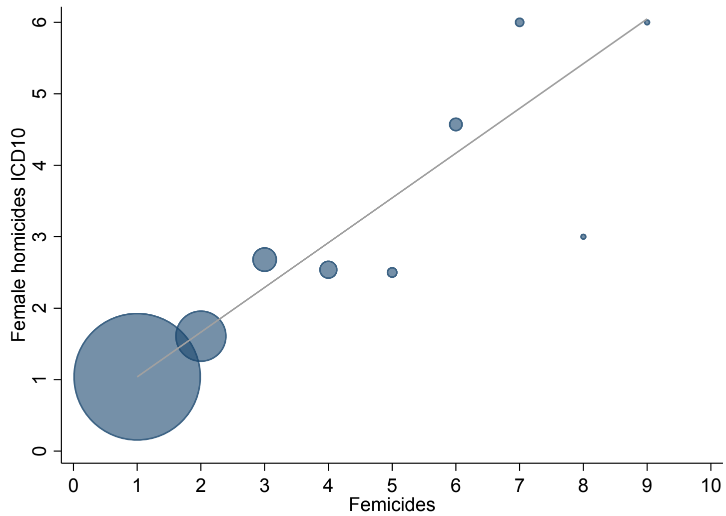
(a) Domestic cases



(b) Other cases

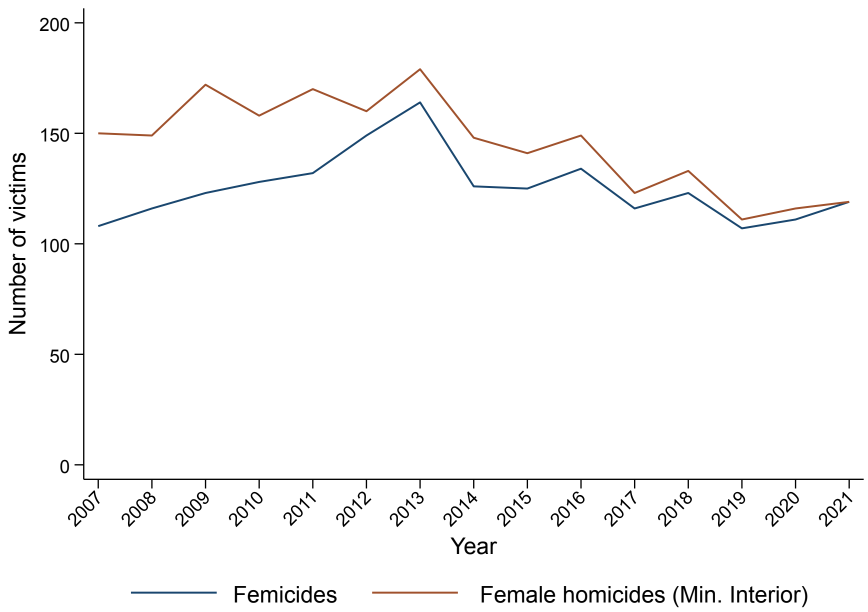
Notes: The figure plots the distribution of reported motives for femicide, as recorded in news coverage of each case. For each panel, bars show the percentage of cases falling into each motive category, computed separately within that panel's sample. Panel (a) restricts the sample to domestic cases (partner, ex-partner, or relative); Panel (b) restricts the sample to non-domestic cases (all other perpetrators). Categories are sorted in descending order of frequency within each panel.

Figure A.4. Correlation between femicides and ISTAT female homicides



Notes: The figure plots a binned scatter plot at the municipality-year level, restricted to observations with at least one femicide. The horizontal axis reports the number of femicides recorded in our dataset; the vertical axis reports the number of female homicides constructed using the ISTAT *Cause di morte* registry. Each marker shows the average value of the vertical-axis variable within a bin of the horizontal-axis variable; marker size is proportional to the number of observations in that bin. The solid line reports the linear fit across bin averages.

Figure A.5. Homicide and femicide victims



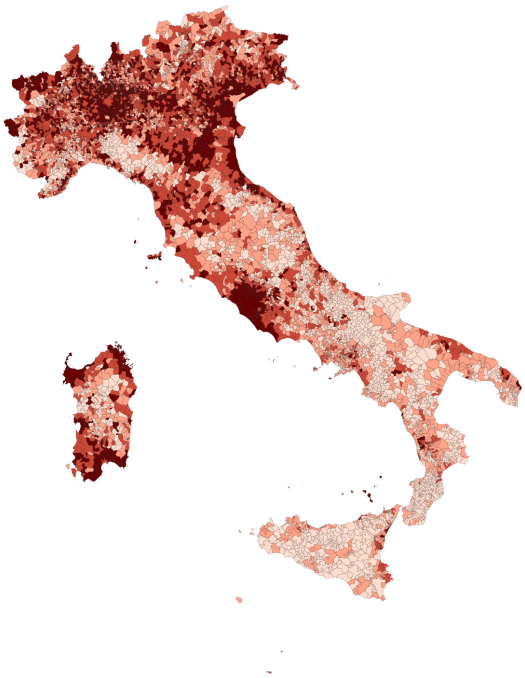
Notes: The figure plots the annual number of female homicide victims in Italy over the period 2007–2021, according to two sources. The blue line reports femicides recorded in our dataset. The orange line reports female homicide victims according to the Ministry of the Interior.

Figure A.6. Geographic distribution of the *shares*

(a) Age group 25-29 (b) Age group 45-49 (c) Age group 60-64

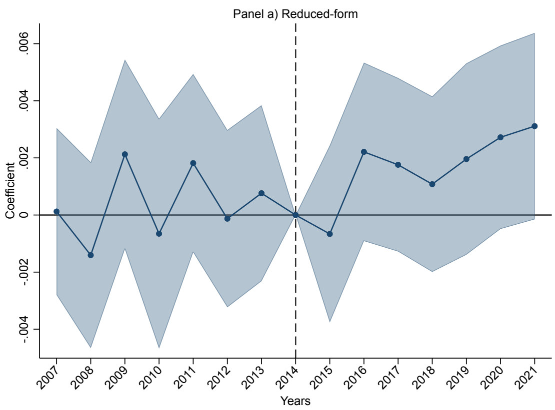
Notes: The figure plots the geographic distribution of the population share of individuals in age group 25–29 (Panel (a)), 45–49 (Panel (b)), and 60–64 (Panel (c)). The share is computed as the number of individuals in each age group divided by the total population of the municipality, as measured in the 2011 Census.

Figure A.7. Map of the instrument

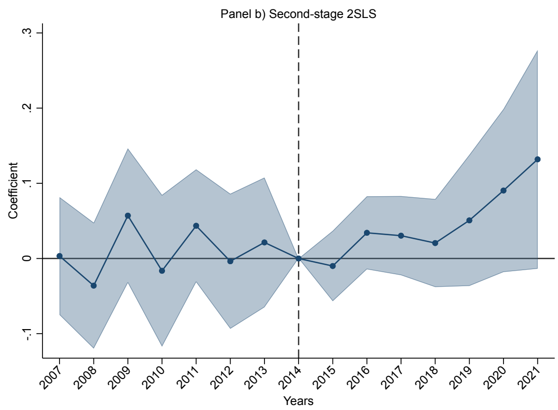


Notes: The map shows the distribution of the instrument as defined in Equation 1.

Figure A.8. Dynamic treatment effects - number of femicides



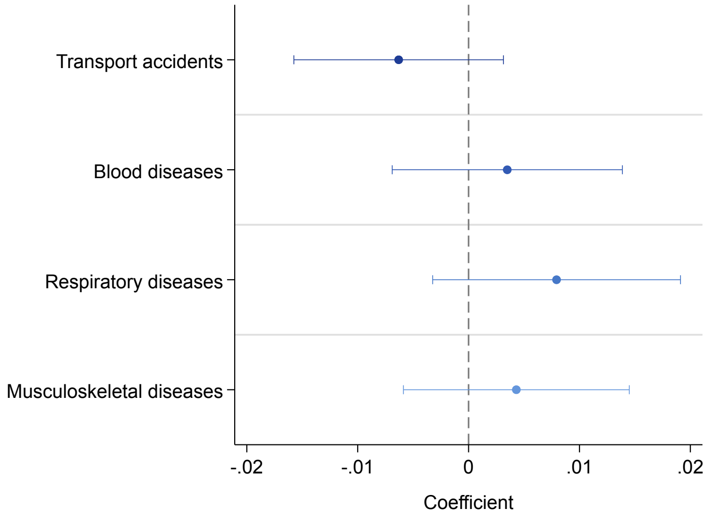
(a) Reduced-form



(b) Second-stage 2SLS

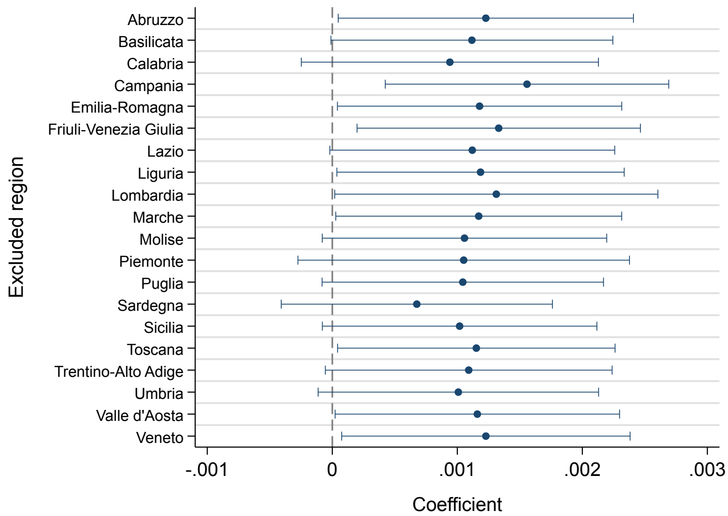
Notes: The figure plots the estimated coefficients from the dynamic specifications of Equation 2 (Panel (a)) and Equation 4 (Panel (b)). The dependent variable is the number of domestic femicides committed in municipality m in year t. The shaded area denotes 95% confidence intervals, based on standard errors clustered at the municipal level. The vertical dashed line marks the last pre-reform year, 2014.

Figure A.9. Transport accidents and other accidental deaths



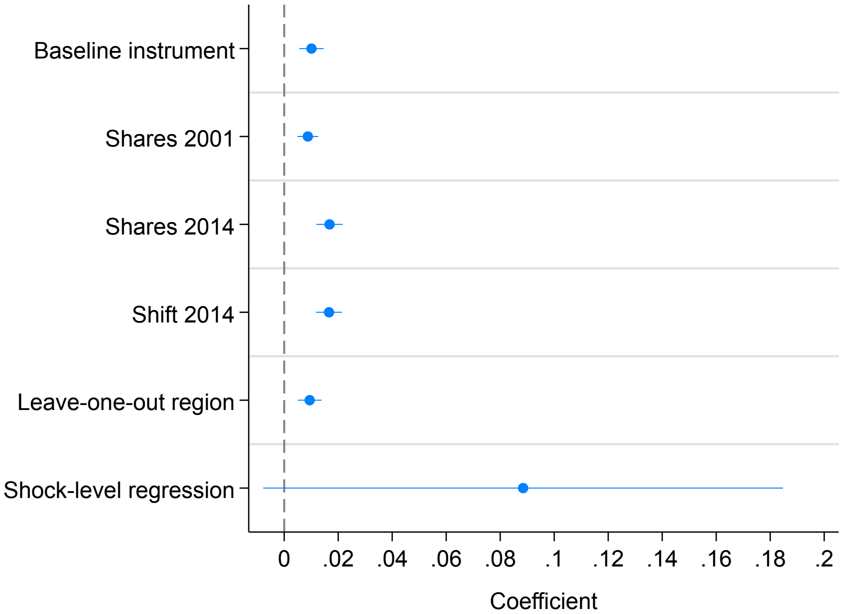
Notes: The figure plots the estimated coefficients from Equation 2 using four alternative placebo outcomes, capturing municipal mortality from causes unrelated to intentional violence. For each outcome, the dependent variable is an indicator equal to one if at least one death from that cause occurred in the municipality-year, standardized using its pre-reform mean and standard deviation. *Transport accidents* is identified using ICD-10 codes V01–V99, including accidents involving pedestrians, cyclists, motor vehicles, buses, trains, and water or air transport. *Blood diseases* corresponds to ICD-10 codes D50–D89, including conditions such as iron deficiency anemia, hemolytic anemias, coagulation defects (e.g., hemophilia), and immune disorders. *Respiratory diseases* corresponds to ICD-10 codes J00–J99, including acute respiratory infections, chronic obstructive pulmonary disease (COPD), pneumonia, and asthma. *Musculoskeletal diseases* corresponds to ICD-10 codes M00–M99, including rheumatoid arthritis, osteoarthritis, osteoporosis, and connective tissue disorders. Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Control variables are defined in Table 1.

Figure A.10. Probability of femicide, excluding regions



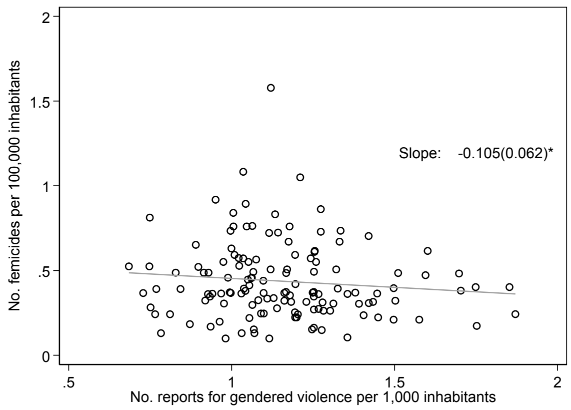
Notes: The figure plots the estimated coefficients from Equation 2, sequentially excluding one Italian region at a time. Each row reports the coefficient obtained when the region indicated on the vertical axis is dropped from the sample. Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Control variables are defined in Table 1.

Figure A.11. Alternative instruments

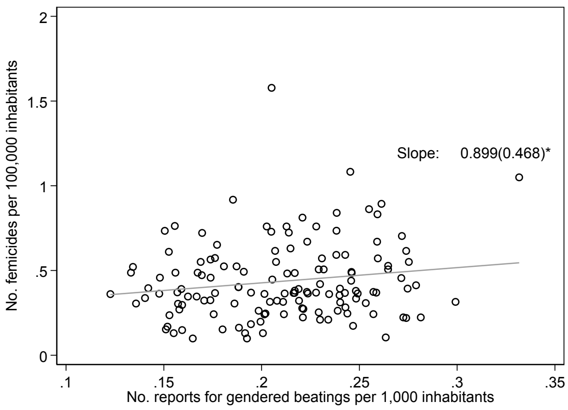


Notes: The figure plots the estimated coefficients from Equation 3 across alternative constructions of the instrument. *Baseline instrument* reports the coefficient using the instrument defined in Equation 1. *Shares 2001* uses population shares measured in the 2001 Census instead of 2011. *Shares 2014* uses population shares measured in 2014 (immediately before the reform) instead of 2011. *Shift 2014* uses both the 2014 population shares and the 2014 divorce rates to construct the shift. *Leave-one-out region* computes the shift using the divorce rate calculated after excluding the region of the municipality. *Shock-level regression* reports the coefficient from a regression run at the level of age groups, following the shock-level aggregation proposed by Borusyak et al. (2022). Horizontal bars denote 95% confidence intervals, based on standard errors clustered at the municipal level. Control variables are defined in Table 1.

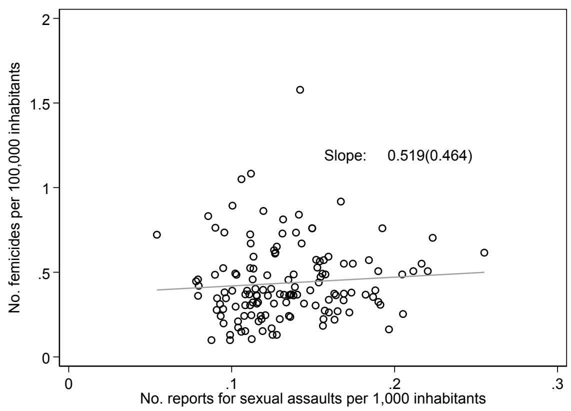
Figure A.12. Correlation between femicides and measures of reported gendered violence



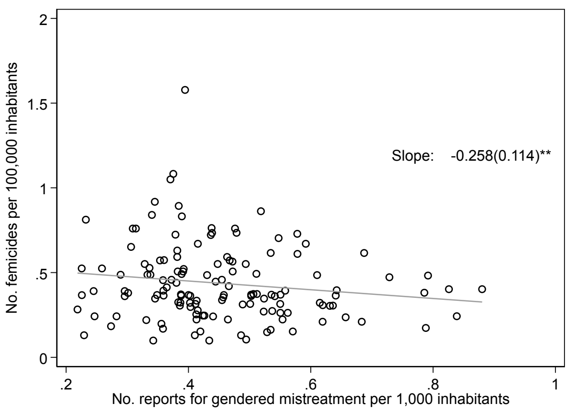
(a) Overall



(b) Beatings



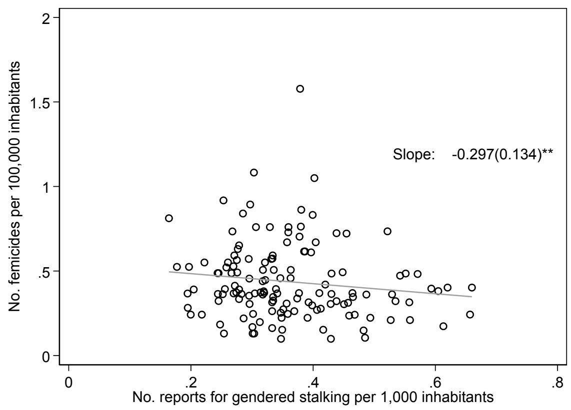
(c) Sexual assault



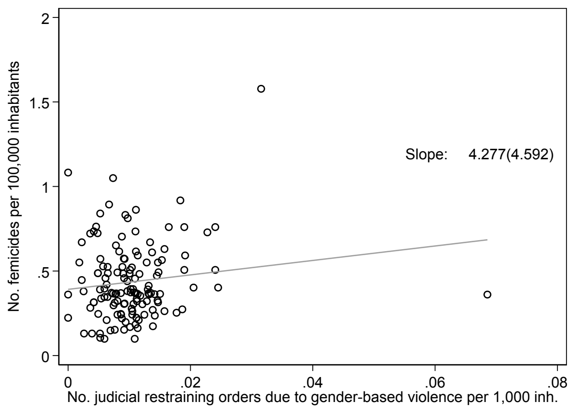
(d) Mistreatment

Notes: The figure plots the correlation between femicides and reported incidents of gender-based violence at the region-year level over the period 2014–2021: overall (Panel (a)), beatings (Panel (b)), sexual assaults (Panel (c)), and mistreatment (Panel (d)). Femicides are expressed per 100,000 female residents; reported incidents of gender-based violence are expressed per 1,000 female residents. In both cases, the female population is held fixed at its 2021 level throughout the sample period. Reported incidents refer to criminal offenses reported to the judicial authority by the State Police, the Carabinieri Corps, and the Financial Police, as recorded by ISTAT. Each panel reports the slope coefficient and standard error (in parentheses) from a linear fit of femicides on the corresponding measure of reported violence; *, **, *** denote significance at the 10%, 5%, and 1% level, respectively.

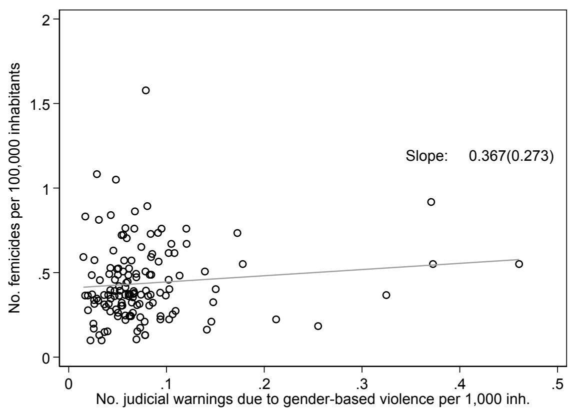
Figure A.13. Correlation between femicides and measures of reported gendered violence



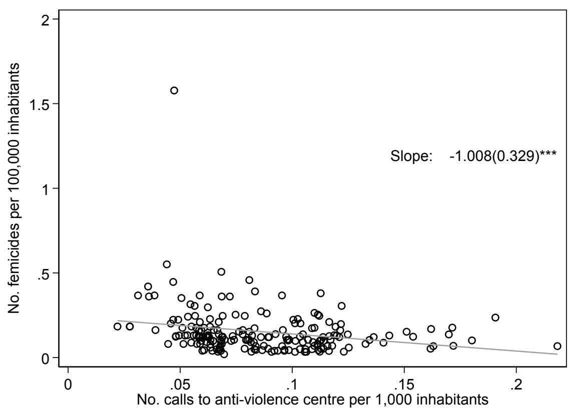
(a) Stalking



(b) Judicial restraining orders



(c) Judicial warnings



(d) Calls to AVC

Notes: The figure plots the correlation between femicides and reported incidents of gender-based violence at the region-year level over the period 2014–2021: stalking (Panel (a)), judicial restraining orders (Panel (b)), and judicial warnings (Panel (c)), each due to gender-based violence. Femicides are expressed per 100,000 female residents; reported incidents are expressed per 1,000 female residents. In both cases, the female population is held fixed at its 2021 level throughout the sample period. Data in Panels (a)–(c) refer to criminal offenses and judicial measures reported to the judicial authority by the State Police, the Carabinieri Corps, and the Financial Police, as recorded by ISTAT. Panel (d) plots the correlation between femicides and calls to anti-violence centers, at the region-quarter level over 2018–2021; call data are published at the provincial level by ISTAT and aggregated to the region level for consistency with the other panels. Each panel reports the slope coefficient and standard error (in parentheses) from a linear fit of femicides on the corresponding measure of reported violence; *, **, *** denote significance at the 10%, 5%, and 1% level, respectively.

B Appendix Tables

Table B.1. Descriptive statistics: femicides

	N.	Mean	Std. Dev.
Panel A: Victim's charactersitics			
Age	1868	49.677	20.434
Italian	1876	0.761	0.427
Single or widowed	1881	0.181	0.385
In a relationship	1881	0.503	0.500
Divorced/Separated	1881	0.216	0.412
Has children	1661	0.656	0.475
Panel B: Killer's characteristics			
Male	1807	0.972	0.164
Age	1731	48.677	17.346
Italian	1799	0.807	0.395
Single or widowed	1881	0.095	0.293
In a relationship	1881	0.475	0.500
Divorced/Separated	1881	0.238	0.426
Panel C: Victim-killer relationship			
Relationship: partner	1881	0.458	0.498
Relationship: ex-partner	1881	0.127	0.333
Relationship: relative	1881	0.223	0.416
Relationship: professional	1881	0.058	0.235
Relationship: stalker	1881	0.001	0.023
Relationship: other	1881	0.072	0.258
Relationship: unknown	1881	0.062	0.242
Reason for murder: Family issues	1881	0.603	0.489
Reason for murder: Monetary reasons	1881	0.161	0.367
Reason for murder: Mental health issues	1881	0.120	0.325
Reason for murder: Toxic dynamics	1881	0.233	0.423
Reason for murder: Addiction-related factors	1881	0.059	0.236
Killed in residence	1378	0.901	0.298

Notes: The table reports means and standard deviations for the main variables in our femicide dataset. Panel (a) reports victim characteristics, Panel (b) reports perpetrator characteristics, and Panel C reports the victim-perpetrator relationship, the reported motive for the murder, and whether the murder occurred in the victim's municipality of residence. Reasons for murder are not mutually exclusive, as a single case can be coded under more than one category. *Family issues* includes illness, abandonment, betrayal, quarrels, jealousy, honor-related conflicts, domestic abuse, and disputes involving children or other relatives. *Monetary reasons* includes debt, extortion, broader financial distress, inheritance disputes, and conflicts over separation payments or wages. *Mental health issues* includes depression, psychological distress, or other diagnosed mental health conditions. *Toxic dynamics* includes sexual assault, revenge, impulsive violent episodes, unrequited love, or possessive behavior. *Addiction-related factors* includes alcohol abuse, drug abuse, or gambling.

Table B.2. First stage

	(1)	(2)	(3)
Dependent variable =		New divorced (%)	
Post-reform × Instrument (std.)	0.0081***	0.0082***	0.0101***
	(0.0018)	(0.0018)	(0.0023)
Obs	118410	118410	118410
Municipalities	7894	7894	7894
Year and Mun. FE	✓	✓	✓
Population × Year		✓	✓
Pre-determined covs × Year			✓
Mean dep. var pre-reform	0.137	0.137	0.137
R-squared	0.254	0.254	0.257
First-stage F	20.327	20.615	19.375

Notes: OLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *Instrument (std.)* is the standardized measure defined in Equation 1. *Post-reform* is a dummy equal to 1 from 2015 onward and 0 otherwise. *Population* is the municipal population, measured in the 2011 Census and interacted with year fixed effects. Pre-determined characteristics include the shares of women, children aged 0–4, individuals with tertiary education, and foreign residents, all measured in the 2011 Census and interacted with year fixed effects.

Table B.3. First stage with divorce applications

	(1)	(2)	(3)
Dependent variable =	New divorced (applications) (%)		
Post-reform × Instrument (std.)	0.0073***	0.0073***	0.0097***
	(0.0018)	(0.0018)	(0.0023)
Obs	118410	118410	118410
Municipalities	7894	7894	7894
Year and Mun. FE	✓	✓	✓
Yearly population		✓	✓
Pre-determined covs × Year			✓
Mean dep. var pre-reform	0.140	0.140	0.140
R-squared	0.260	0.260	0.262
First-stage F	16.345	16.628	18.347

Notes: New divorced (applications) (%) is defined using divorce applications rather than finalized divorces, unlike *New divorced (%)* used in the main tables. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *Instrument (std.)* is the standardized measure defined in Equation 1. *Post-reform* is a dummy equal to 1 from 2015 onward and 0 otherwise. Control variables are described in Table 1.

Table B.4. Change in divorce characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable (Type of divorce) =	Consensual (%)	Judicial (%)	Common property regime (%)	Children present (%)	Joint custody (%)	Custody to mother (%)	Custody to father (%)	Transfers (%)
Post-reform × Instrument (std.)	-1.9784***	1.7317***	1.4973***	0.5693	-1.7596***	1.6899***	0.1126	-0.6104**
	(0.3290)	(0.3257)	(0.4247)	(0.3592)	(0.4773)	(0.4677)	(0.1876)	(0.2412)
Obs	15368	15368	15368	15368	14204	14204	14204	15368
Municipalities	7684	7684	7684	7684	7102	7102	7102	7684
Mun. FE	✓	✓	✓	✓	✓	✓	✓	✓
Pre-det. covs × Post-reform	✓	✓	✓	✓	✓	✓	✓	✓
Mean dep. var pre-reform	84.817	14.409	55.877	36.787	44.382	50.605	4.343	88.211
R-squared	0.587	0.585	0.671	0.543	0.817	0.803	0.533	0.549

Notes: OLS estimates comparing two periods, 2013 (pre-reform) and 2015 (post-reform), with municipality fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *Instrument (std.)* is the standardized measure of the instrument defined in Equation 1, interacted with the post-reform period. *Post-reform* is a dummy equal to 1 in 2015 and 0 in 2013. Each dependent variable is the share (%) of divorces of a given type out of the total number of divorces for which that information is available: type of completion procedure (columns (1)–(2)), matrimonial property regime (column (3)), presence of children (column (4)), and custody arrangements and transfers (columns (5)–(8)). The sample size is smaller in columns (5)–(7) because custody-related information is available only for divorces involving children. Data are obtained from custom tabulations produced by ISTAT. Pre-determined characteristics – municipal population and the shares of women, children aged 0–4, individuals with tertiary education, and foreign residents, all measured in the 2011 Census – are interacted with the post-reform dummy.

Table B.5. Probability of femicide - second-stage 2SLS

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable =	At least one femicide (any killer)		At least one domestic femicide		At least one femicide committed by other killers	
New divorced (%)	0.1395**	0.1120*	0.1554**	0.1167**	-0.0011	0.0038
	(0.0658)	(0.0621)	(0.0643)	(0.0589)	(0.0238)	(0.0250)
Obs	118410	118410	118410	118410	118410	118410
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.013	0.013	0.011	0.011	0.003	0.003

Notes: 2SLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *New divorced (%)* is the endogenous regressor, instrumented as described in Equation 1. The dependent variable is an indicator equal to one if at least one femicide occurred in the municipality, committed by any killer in columns (1)–(2), by a partner, ex-partner, or relative (domestic femicide) in columns (3)–(4), and by another type of perpetrator (e.g., professional relation, acquaintance, stalker, or unknown) in columns (5)–(6). Control variables are described in Table 1.

Table B.6. Number of femicides - Second-stage 2SLS

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable =	Total number of femicides		Total number of domestic femicides		Total number of femicides (other killers)	
New divorced (%)	0.2001**	0.1540**	0.1971**	0.1395**	0.0031	0.0145
	(0.0815)	(0.0758)	(0.0771)	(0.0689)	(0.0258)	(0.0269)
Obs	118410	118410	118410	118410	118410	118410
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.017	0.017	0.013	0.013	0.004	0.004

Notes: 2SLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *New divorced (%)* is the endogenous regressor, instrumented as described in Equation 1. The dependent variable is the total number of femicides in the municipality in columns (1)–(2), the number of domestic femicides (committed by a partner, ex-partner, or relative) in columns (3)–(4) and the number of femicides committed by other killers (e.g., professional relation, acquaintance, stalker, or unknown) in columns (5)-(6). Control variables are described in Table 1.

Table B.7. Number of femicides - alternative specifications

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Reduced-form						
Dependent variable =		Total number of femicides			Total number of domestic femicides	
Post-reform × Instrument (std.)	0.0012**	0.2433***	0.2222***	0.0012**	0.2344***	0.2132**
	(0.0006)	(0.0830)	(0.0832)	(0.0005)	(0.0902)	(0.0904)
Estimator	IHS	PPML	PPML - rate	IHS	PPML	PPML - rate
Obs	118410	14970	14970	118410	12975	12975
Municipalities	7894	998	998	7894	865	865
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓	✓	✓	✓
Mean dep. var pre-reform	0.013	0.131	0.131	0.011	0.118	0.118
R-squared	0.268			0.232		
Pseudo R-squared		0.173	0.173		0.144	0.145
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B: Second-stage 2SLS						
Dependent variable =		Total number of femicides			Total number of domestic femicides	
New divorced (%)	0.1215**			0.1157**		
	(0.0612)			(0.0569)		
Linear prediction		24.0714***	21.9835***		23.1913***	21.0886**
		(8.2085)	(8.2291)		(8.9194)	(8.9474)
Estimator	IHS	PPML	PPML - rate	IHS	PPML	PPML - rate
Obs	118410	14970	14970	118410	12975	12975
Municipalities	7894	998	998	7894	865	865
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓	✓	✓	✓
Mean dep. var pre-reform	0.013	0.131	0.131	0.013	0.118	0.118
Pseudo R-squared		0.173	0.173		0.144	0.145

Notes: The table reports estimates using alternative estimators for the femicide count, with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. Panel (a) reports reduced-form estimates; Panel (b) reports second-stage 2SLS estimates, where *New divorced (%)* is the endogenous regressor, instrumented as described in Equation 1. The dependent variable is the total number of femicides in the municipality in columns (1)–(3), and the number of domestic femicides (committed by a partner, ex-partner, or relative) in columns (4)–(6). For each dependent variable, we report three estimators, indicated in the *Estimator* row. The first column applies the inverse hyperbolic sine (IHS) transformation to the outcome; the second and third columns use the Poisson pseudo-maximum likelihood (PPML) estimator, with the third additionally using yearly municipal population as the exposure variable, so that the coefficient can be interpreted as an effect on the femicide rate. In the PPML columns of Panel (b), *Linear prediction* reports the coefficient on the first-stage predicted value of *New divorced (%)*, following a control-function approach, since standard 2SLS is not valid under the nonlinear PPML model. The IHS transformation is similar to the natural logarithm, reducing the influence of extreme values while retaining zero observations. The number of observations is lower in the PPML columns than in the main tables because PPML drops observations for which the outcome is perfectly predicted. Control variables are described in Table 1.

Table B.8. Probability of female homicide - second-stage 2SLS

	(1)	(2)	(3)	(4)	(5)
Dependent variable =	At least one femicide (any killer)	At least one female homicide	At least one female homicide (balanced sample)	At least one female homicide (selected weapons)	At least one male homicide (selected weapons)
New divorced (%)	0.1120*	0.1353**	0.1750**	0.0830*	0.0114
	(0.0621)	(0.0632)	(0.0709)	(0.0477)	(0.0476)
Obs	118410	117859	80925	117859	117859
Municipalities	7894	7894	5395	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓	✓	✓
Mean dep. var pre-reform	0.013	0.012	0.017	0.006	0.007

Notes: 2SLS estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. *New divorced (%)* is the endogenous regressor, instrumented as described in Equation 1. The dependent variable in column (1) is the main dependent variable described in Table 1. In column (2), the dependent variable is an indicator equal to one if at least one female homicide occurred in the municipality, constructed using the ISTAT *Cause di morte* registry, which classifies deaths by underlying cause; homicides are identified using ICD codes X85–Y09 and Y87.1. Column (3) restricts the sample to municipalities in which the exact count of female homicide victims is never suppressed for privacy reasons. Column (4) further restricts female homicides to those committed with a selected set of weapons; column (5) applies the same weapon restriction to male homicide victims as a placebo. Control variables are described in Table 1.

Table B.9. Probability of femicide, excluding 2021

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Reduced-form						
Dependent variable =	At least one femicide (any killer)		At least one domestic femicide		At least one femicide committed by other killers	
Post-reform × Instrument (std.)	0.0010**	0.0011*	0.0011**	0.0011**	0.0000	0.0001
	(0.0005)	(0.0006)	(0.0004)	(0.0005)	(0.0002)	(0.0003)
Obs	110516	110516	110516	110516	110516	110516
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.013	0.013	0.011	0.011	0.003	0.003
R-squared	0.187	0.187	0.172	0.173	0.175	0.184
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B: Second-stage 2SLS						
Dependent variable =	At least one femicide (any killer)		At least one domestic femicide		At least one femicide committed by other killers	
New divorced (%)	0.0800*	0.0773*	0.0850**	0.0775*	0.0040	0.0039
	(0.0410)	(0.0437)	(0.0386)	(0.0409)	(0.0163)	(0.0187)
Obs	110516	110516	110516	110516	110516	110516
Municipalities	7894	7894	7894	7894	7894	7894
Year and Mun. FE	✓	✓	✓	✓	✓	✓
Pre-det. covs × Year		✓		✓		✓
Mean dep. var pre-reform	0.013	0.013	0.011	0.011	0.003	0.003

Notes: The table replicates the main specification excluding 2021 from the sample. Panel (a) reports reduced-form estimates; Panel (b) reports second-stage 2SLS estimates, where *New divorced (%)* is the endogenous regressor, instrumented as described in Equation 1. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is an indicator equal to one if at least one femicide occurred in the municipality, committed by any killer in columns (1)–(2), by a partner, ex-partner, or relative (domestic femicide) in columns (3)–(4), and by another type of perpetrator (e.g., professional relation, acquaintance, stalker, or unknown) in columns (5)–(6). Control variables are described in Table 1.

Table B.10. Heterogeneity by divorce attitudes and characteristics

	(1)	(2)	(3)
Dependent variable =	At least one domestic femicide		
Post-reform × Instrument (std.)	0.0023***	0.0023***	0.0009
	(0.0008)	(0.0008)	(0.0007)
Post × Instrument × Share of voters against divorce (above median)	-0.0019**		
	(0.0009)		
Post × Instrument × Share consensual divorces (above median)		-0.0018*	
		(0.0009)	
Post × Instrument × Share divorces with children (above median)			0.0005
			(0.0009)
Obs	117675	117240	117240
Municipalities	7845	7816	7816
Year and Mun. FE	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓
P-value ($\beta_1 + \beta_2$)	0.444	0.482	0.043
Mean dep. var pre-reform	0.011	0.011	0.011
R-squared	0.169	0.169	0.169

Notes: OLS reduced-form estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is an indicator equal to one if at least one domestic femicide (committed by a partner, ex-partner, or relative) occurred in the municipality. Each column adds an interaction between *Post-reform × Instrument (std.)* and an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise: divorce attitudes, measured as the share of voters who voted "Yes" in the 1974 divorce referendum (column (1)); the pre-reform share of consensual divorces out of total divorces (column (2)); and the pre-reform share of divorces involving children out of total divorces (column (3)). The coefficient on *Post-reform × Instrument (std.)* (β_1) captures the effect for municipalities below the median; the coefficient on the triple interaction (β_2) captures the difference in effect for municipalities above the median. *P-value* ($\beta_1 + \beta_2$) reports the p-value from a test of the null hypothesis that the total effect for municipalities above the median, $\beta_1 + \beta_2$, is equal to zero. Control variables are described in Table 1.

Table B.11. Heterogeneity by religious presence and social capital

	(1)	(2)
Dependent variable =	At least one domestic femicide	
Post-reform × Instrument (std.)	0.0023**	0.0028**
	(0.0011)	(0.0011)
Post × Instrument × N. churches per 1000 inh. (above median)	-0.0019*	
	(0.0011)	
Post × Instrument × N. associations per 1000 inh. (above median)		-0.0025**
		(0.0011)
Obs	118005	118155
Municipalities	7867	7877
Year and Mun. FE	✓	✓
Pre-det. covs × Year	✓	✓
P-value ($\beta_1 + \beta_2$)	0.226	0.413
Mean dep. var pre-reform	0.011	0.011
R-squared	0.168	0.168

Notes: OLS reduced-form estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is an indicator equal to one if at least one domestic femicide (committed by a partner, ex-partner, or relative) occurred in the municipality. Each column adds an interaction between *Post-reform* × *Instrument (std.)* and an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise: religious presence, measured as the number of churches per 1,000 inhabitants (column (1)); and social capital, measured as the number of associations per 1,000 inhabitants (column (2)). The coefficient on *Post-reform* × *Instrument (std.)* (β_1) captures the effect for municipalities below the median; the coefficient on the triple interaction (β_2) captures the difference in effect for municipalities above the median. *P-value* ($\beta_1 + \beta_2$) reports the p-value from a test of the null hypothesis that the total effect for municipalities above the median, $\beta_1 + \beta_2$, is equal to zero. Control variables are described in Table 1.

Table B.12. Heterogeneity by intentional homicide and domestic violence

	(1)	(2)
Dependent variable =	At least one domestic femicide	
Post-reform × Instrument (std.)	0.0005	0.0005
	(0.0006)	(0.0007)
Post × Instrument × N. homicides per 100000 inh. (above median)	0.0013	
	(0.0009)	
Post × Instrument × N. DV crimes per 100000 inh. (above median)		0.0014
		(0.0009)
Obs	118410	118410
Municipalities	7894	7894
Year and Mun. FE	✓	✓
Pre-det. covs × Year	✓	✓
P-value ($\beta_1 + \beta_2$)	0.022	0.008
Mean dep. var pre-reform	0.011	0.011
R-squared	0.169	0.168

Notes: OLS reduced-form estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is an indicator equal to one if at least one domestic femicide (committed by a partner, ex-partner, or relative) occurred in the municipality. Each column adds an interaction between *Post-reform × Instrument (std.)* and an indicator variable equal to one if the province in which the municipality is located is above the national median along a given dimension in the pre-reform period (2007–2013), and zero otherwise: the rate of intentional homicides per 100,000 inhabitants (column (1)); and the rate of domestic violence crimes per 100,000 inhabitants (column (2)). The coefficient on *Post-reform × Instrument (std.)* (β_1) captures the effect for municipalities in below-median provinces; the coefficient on the triple interaction (β_2) captures the difference in effect for municipalities in above-median provinces. *P-value ($\beta_1 + \beta_2$)* reports the p-value from a test of the null hypothesis that the total effect for municipalities in above-median provinces, $\beta_1 + \beta_2$, is equal to zero. Control variables are described in Table 1.

Table B.13. Heterogeneity by income, employment gap and education gap

	(1)	(2)	(3)
Dependent variable =		At least one domestic femicide	
Post-reform × Instrument (std.)	0.0001	0.0015**	0.0017**
	(0.0006)	(0.0007)	(0.0007)
Post × Instrument × Mean income (above median)	0.0029**		
	(0.0013)		
Post × Instrument × Employment rate gap (above median)		-0.0009	
		(0.0009)	
Post × Instrument × Education rate gap (above median)			-0.0007
			(0.0009)
Obs	118410	118410	118410
Municipalities	7894	7894	7894
Year and Mun. FE	✓	✓	✓
Pre-det. covs × Year	✓	✓	✓
P-value ($\beta_1 + \beta_2$)	0.013	0.380	0.139
Mean dep. var pre-reform	0.011	0.011	0.011
R-squared	0.168	0.169	0.168

Notes: OLS reduced-form estimates with municipality and year fixed effects. Standard errors (in parentheses) are clustered at the municipal level. *, **, *** denote significance at the 10%, 5%, and 1% level, respectively. The dependent variable is an indicator equal to one if at least one domestic femicide (committed by a partner, ex-partner, or relative) occurred in the municipality. Each column adds an interaction between *Post-reform* × *Instrument (std.)* and an indicator variable equal to one if the municipality is above the national median along a given dimension, and zero otherwise: the mean income at the municipal level in the pre-reform period (2007–2013) (column (1)); the employment rate gap, measured in the 2011 Census (column (2)); and the education rate gap, measured in the 2011 Census (column (3)). The coefficient on *Post-reform* × *Instrument (std.)* (β_1) captures the effect for municipalities below the median; the coefficient on the triple interaction (β_2) captures the difference in effect for municipalities above the median. *P-value* ($\beta_1 + \beta_2$) reports the p-value from a test of the null hypothesis that the total effect for municipalities above the median, $\beta_1 + \beta_2$, is equal to zero. Control variables are described in Table 1.

C Theoretical Appendix: Assumptions, Lemmas, and Proofs

This appendix provides the formal characterization of the model in the main text. Table C.14 summarizes the characterization of the predicted effect of reducing the cost of divorce. Rather than naming abstract sets of violence levels or husband types, we describe every choice directly as a comparison against a small number of thresholds, each introduced at the point where it first arises.

C.1 Model setup and payoffs

There are two agents, a husband H and a wife W. The couple generates relationship surplus $S > 0$. The husband chooses non-lethal violence $v \in [0, \overline{v}]$ and has type $\theta \in [\underline{\theta}, \overline{\theta}]$, observed by both spouses.

The timing is as follows. First, the husband chooses v. Second, after observing v, the wife chooses whether to stay or attempt separation. Third, if separation is attempted, a shock ε is realized and observed by the husband, who then chooses whether to accept separation or escalate.

If the wife stays, payoffs are

$$W^S(v) = (1 - \mu(v))S, \qquad H^S(v, \theta) = \mu(v)S - \theta c(v).$$

If the wife attempts separation and the husband accepts, payoffs are

$$W^A(\kappa) = (1 - \alpha)\delta S - \kappa, \qquad H^A(v, \kappa) = \alpha\delta S - \kappa - \chi c(v).$$

If instead the husband escalates, payoffs are

$$W^E = -M, \qquad H^E(v, \varepsilon) = (1 - q)\delta S - q[\Psi + \chi c(v)] + \varepsilon.$$

Let $p(v, \kappa)$ denote the equilibrium probability that the husband accepts separation following a separation attempt. The wife's expected payoff from attempting separation is

$$W^L(v, \kappa) = p(v, \kappa)W^A(\kappa) + \big(1 - p(v, \kappa)\big)W^E,$$

and define

$$G(v, \kappa) := W^S(v) - W^L(v, \kappa).$$

C.2 Assumptions

Assumption C.1 (Primitive regularity and bounded domains). *The functions μ and c satisfy the following conditions:*

1. *μ and c are continuous on $[0,\overline{v}]$ and twice continuously differentiable on $(0,\overline{v})$.*

2. *$0 \le \mu(v) < 1$ for all $v \in [0,\overline{v}]$, $\mu'(v) > 0$ and $\mu''(v) < 0$ for all $v \in (0,\overline{v})$, and the one-sided derivative $\mu'(0^+) := \lim_{v\downarrow 0} \mu'(v)$ exists and is strictly positive.*

3. *$c(0) = 0$, $c'(v) > 0$, and $c''(v) > 0$ for all $v \in (0,\overline{v})$.*

4. *The relevant one-sided endpoint conditions are $\lim_{v\downarrow 0} c'(v) = 0$ and $\lim_{v\uparrow\overline{v}} \mu'(v) = 0$.*

Lemma C.1 (Monotonicity and limit behavior of the ratio μ'/c'). *Under Assumption C.1, the ratio $\mu'(v)/c'(v)$ is strictly decreasing on $(0,\overline{v})$ and satisfies*

$$\lim_{v\downarrow 0} \frac{\mu'(v)}{c'(v)} = +\infty, \qquad \lim_{v\uparrow\overline{v}} \frac{\mu'(v)}{c'(v)} = 0.$$

Proof. For every $v \in (0,\overline{v})$,

$$\frac{d}{dv}\left(\frac{\mu'(v)}{c'(v)}\right) = \frac{\mu''(v)c'(v) - \mu'(v)c''(v)}{[c'(v)]^2} < 0,$$

because $\mu''(v) < 0$, $c'(v) > 0$, $\mu'(v) > 0$, and $c''(v) > 0$. Hence the ratio is strictly decreasing. At zero, $\mu'(0^+) > 0$ and $c'(v) \to 0$ by Assumption C.1, so the ratio diverges to $+\infty$. Near $\overline{v}$, strict convexity implies that $c'(v)$ is bounded away from zero on every interval $[v_0,\overline{v})$ with $v_0 > 0$, while $\mu'(v) \to 0$, so the ratio converges to zero. □

Assumption C.2 (Payoff parameters, divorce-cost regimes, and the zero-violence endpoint). *The payoff parameters satisfy*

$$S > 0, \qquad \delta \in (0,1), \qquad \alpha \in \left(0, \tfrac{1}{2}\right), \qquad q \in (0,1), \qquad \chi > 0, \qquad M > 0, \qquad \Psi > 0.$$

The husband type satisfies $0 < \underline{\theta} < \overline{\theta} < \infty$ and the two relevant divorce-cost regimes satisfy $0 \le \kappa_L < \kappa_H < (1-\alpha)\delta S + M$.

Finally,

$$(1-\alpha)\delta \le 1 - \mu(0). \tag{C.1}$$

Thus, even at zero divorce cost, an amicably accepted separation does not give the wife a higher payoff than remaining in a non-violent relationship. Because attempted separation also carries *a strictly positive probability of escalation, condition* (C.1) *implies that the wife's expected payoff from attempting separation is strictly below her stay payoff at* $v = 0$*.*

Assumption C.3 (Leave-stage uncertainty and interior acceptance)**.** *The leave-stage shock is drawn from*

$$\varepsilon \sim U[-\overline{\varepsilon}, \overline{\varepsilon}].$$

Let

$$\Delta(v,\kappa) = \alpha\delta S - \kappa - (1-q)\delta S + q\Psi - (1-q)\chi c(v).$$

Throughout the appendix, we maintain interiority of the acceptance cutoff for all relevant $(v,\kappa) \in [0,\overline{v}] \times [\kappa_L, \kappa_H]$*:*

$$-\overline{\varepsilon} < \Delta(v,\kappa) < \overline{\varepsilon}.$$

Assumption C.4 (Indifference and knife-edge treatment)**.** *At the accept-versus-escalate stage, equality is assigned to acceptance; because the shock is continuous, this tie-breaking rule does not affect any probability. At the wife's stay-or-leave stage, staying is strictly preferred when* $G(v,\kappa) > 0$*, attempted separation is strictly preferred when* $G(v,\kappa) < 0$*, and both are best responses when* $G(v,\kappa) = 0$*. No tie-breaking rule is imposed on the husband's stage-1 violence choice: whenever two choices yield the same maximal equilibrium payoff, both are retained.*

Two wife-side knife edges recur below:

$$G(v^m(\kappa),\kappa) = 0 \qquad \textit{and} \qquad G(\overline{v},\kappa) = 0.$$

We characterize these cases explicitly in Proposition C.1 and exclude them from the generic comparative-static characterization. Stage-1 indifference is retained as a set-valued equilibrium outcome, while the final type-by-type comparisons focus on husband types whose stay-versus-separation choice is strict under both regimes.

C.3 The accept-versus-escalate decision

After the wife attempts separation and the shock ε is realized, the husband accepts separation if and only if $H^A(v,\kappa) \geq H^E(v,\varepsilon)$ or equivalently $\varepsilon \leq \Delta(v,\kappa)$. Under Assumption C.3, the probability of acceptance is therefore

$$p(v,\kappa) = \Pr\{\varepsilon \leq \Delta(v,\kappa)\} = \frac{\Delta(v,\kappa) + \overline{\varepsilon}}{2\overline{\varepsilon}}. \tag{C.2}$$

Differentiating (C.2) gives

$$\frac{\partial p(v,\kappa)}{\partial v} = -\frac{(1-q)\chi}{2\bar{\varepsilon}} c'(v) < 0, \qquad \frac{\partial p(v,\kappa)}{\partial \kappa} = -\frac{1}{2\bar{\varepsilon}} < 0.$$

Thus prior violence makes escalation more likely, while lowering the divorce cost makes acceptance more likely.

C.4 The wife's stay-or-leave decision

Using (C.2), the wife's expected payoff from attempting separation is

$$W^L(v,\kappa) = p(v,\kappa)\big((1-\alpha)\delta S - \kappa + M\big) - M. \tag{C.3}$$

Lemma C.2 (Shape of the wife's payoff difference)**.** *Under Assumptions C.1–C.3,*

$$G_v(v,\kappa) = -S\mu'(v) + B(\kappa)c'(v),$$

where

$$B(\kappa) = \frac{(1-q)\chi}{2\bar{\varepsilon}}\big((1-\alpha)\delta S - \kappa + M\big) > 0.$$

For each relevant κ, $G(\cdot,\kappa)$ therefore has a unique interior minimizer $v^m(\kappa) \in (0,\bar{v})$.

Proof. From (C.3) and (C.2),

$$W^L_v(v,\kappa) = -\frac{(1-q)\chi}{2\bar{\varepsilon}}\big((1-\alpha)\delta S - \kappa + M\big)c'(v) = -B(\kappa)c'(v),$$

which gives the stated expression for G_v. Positivity of $B(\kappa)$ follows from Assumption C.2. By Lemma C.1, $\mu'(v)/c'(v)$ is continuous and strictly decreasing from $+\infty$ to 0. Since $B(\kappa)/S > 0$, there is a unique $v^m(\kappa)$ satisfying $G_v(v^m,\kappa) = 0$, with $G_v < 0$ before that point and $G_v > 0$ after it. □

Lemma C.3 (Effect of lower divorce costs on the wife's payoff difference)**.** *For every relevant (v,κ),*

$$G_\kappa(v,\kappa) > 0.$$

Hence, if $\kappa_L < \kappa_H$,

$$G(x,\kappa_L) < G(x,\kappa_H) \qquad \textit{for every } x \in [0,\bar{v}].$$

Consequently: (a) if the wife strictly stays at x under κ_L, she also strictly stays at x under κ_H; and (b) if x is a stay-separation boundary under κ_H, she strictly prefers separation at x under

κ_L.

Proof. Differentiating (C.3) gives

$$W_\kappa^L(v,\kappa) = -\frac{(1-\alpha)\delta S - \kappa + M}{2\bar{\varepsilon}} - p(v,\kappa) < 0,$$

where the first term is strictly negative by Assumption C.2. Since $W^S(v)$ does not depend on κ, $G_\kappa(v,\kappa) = -W_\kappa^L(v,\kappa) > 0$. The remaining statements follow immediately. □

This pointwise shift is used repeatedly below.

Proposition C.1 (Wife's separation thresholds)**.** *Fix a divorce cost κ.*

1. ***No separation.*** *If $G(v^m(\kappa),\kappa) > 0$, the wife stays at every violence level.*

2. ***Bounded separation range.*** *If $G(v^m(\kappa),\kappa) < 0$ and $G(\bar{v},\kappa) > 0$, there exist thresholds*

$$0 < v_1(\kappa) < v^m(\kappa) < v_2(\kappa) < \bar{v}$$

such that attempted separation is a best response exactly for $v \in [v_1(\kappa), v_2(\kappa)]$, is strictly preferred in the interior, and staying is strictly preferred outside the closed interval. Violence levels above $v_2(\kappa)$ form a high-violence stay region, which makes coercive retention feasible.

3. ***Upper-reaching separation range.*** *If $G(v^m(\kappa),\kappa) < 0$ and $G(\bar{v},\kappa) < 0$, there exists a single threshold $0 < v_1(\kappa) < v^m(\kappa)$ such that attempted separation is a best response for $v \in [v_1(\kappa), \bar{v}]$, is strictly preferred for $v \in (v_1(\kappa), \bar{v}]$, and staying is strictly preferred for $v < v_1(\kappa)$. No high-violence stay region exists in this case.*

At the knife edge $G(v^m(\kappa),\kappa) = 0$, both staying and attempted separation are best responses at $v^m(\kappa)$, while staying is strictly preferred everywhere else. At the knife edge $G(\bar{v},\kappa) = 0$ (with $G(v^m(\kappa),\kappa) < 0$), both actions are best responses at $v_1(\kappa)$ and $\bar{v}$, while attempted separation is strictly preferred for $v \in (v_1(\kappa), \bar{v})$. Per C.4, both knife edges are excluded from the generic characterizations below.

Proof. We first establish the sign at zero from primitives. By (C.1),

$$W^A(\kappa = 0) = (1-\alpha)\delta S \leq (1-\mu(0))S = W^S(0),$$

and therefore $W^A(\kappa) \leq W^S(0)$ for every relevant κ. Moreover, $W^E = -M < W^S(0)$, and Assumption C.3 implies $p(0,\kappa) \in (0,1)$. Hence $W^L(0,\kappa) < W^S(0)$, i.e. $G(0,\kappa) > 0$.

By Lemma C.2, $G(\cdot,\kappa)$ is continuous, strictly decreasing before $v^m(\kappa)$, and strictly increasing after it. If the minimum is positive, G is positive everywhere. If the minimum is negative and $G(\bar{v},\kappa)>0$, continuity and the U-shape give exactly two roots, with $G<0$ strictly between them. If instead the minimum and $G(\bar{v},\kappa)$ are both negative, there is exactly one root on the decreasing side and G remains negative through the upper endpoint. The two knife-edge cases follow immediately from the same geometry. □

C.5 The husband's stage-1 problem

Let $v^*(\theta;\kappa)$ denote the husband's equilibrium stage-1 violence choice for type θ under divorce cost κ.

If the wife stays, the husband's payoff is $\Pi^S(v,\theta)=S\mu(v)-\theta c(v)$. If the wife attempts separation, the husband's continuation value averages his stage payoffs H^A, H^E over the shock:

$$\Pi^L(v,\kappa)=\mathbb{E}\big[\max\{H^A(v,\kappa),H^E(v,\varepsilon)\}\big].$$

Below, V^S and V^L optimize Π^S and Π^L over the violence levels available to the husband under each of the wife's responses – so the roadmap is stage payoffs (H) $\to$ continuation values that integrate over ε (Π) $\to$ values optimized over feasible violence (V).

Using the acceptance cutoff $\Delta(v,\kappa)$ and the uniform distribution of ε,

$$\Pi^L(v,\kappa)=(1-q)\delta S-q\Psi-q\chi c(v)+\frac{\Delta(v,\kappa)}{2}+\frac{\Delta(v,\kappa)^2}{4\bar{\varepsilon}}+\frac{\bar{\varepsilon}}{4}. \tag{C.4}$$

Since $\Delta_v(v,\kappa)=-(1-q)\chi c'(v)$, differentiating (C.4) gives

$$\frac{\partial \Pi^L(v,\kappa)}{\partial v}=-\chi c'(v)\left[q+\frac{1-q}{2}+\frac{1-q}{2\bar{\varepsilon}}\Delta(v,\kappa)\right]<0,$$

where the inequality follows from Assumption C.3. Thus, whenever the husband induces separation, he chooses the lowest violence level that supports it.

Lemma C.4 (Unconstrained stay optimum)**.** *For every $\theta\in[\underline{\theta},\overline{\theta}]$, the unconstrained stay problem has a unique interior solution $v_u(\theta)\in(0,\bar{v})$ satisfying $S\mu'(v_u(\theta))=\theta c'(v_u(\theta))$. Moreover, $v_u(\theta)$ is strictly decreasing in θ.*

Proof. By Lemma C.1, $\Pi^S_v(v,\theta)=S\mu'(v)-\theta c'(v)$ is positive near zero and negative near $\bar{v}$. Since $\Pi^S_{vv}=S\mu''(v)-\theta c''(v)<0$, the payoff is strictly concave and has a unique maximizer. Monotonicity follows from the fact that μ'/c' is strictly decreasing. □

When a bounded separation range exists under divorce cost κ, define the type-space counterparts of the wife's two violence thresholds by

$$v_u(\theta_1(\kappa)) = v_1(\kappa), \qquad v_u(\theta_2(\kappa)) = v_2(\kappa).$$

Since $v_u(\theta)$ is strictly decreasing,

$$\theta_2(\kappa) < \theta_1(\kappa),$$

and

$$v_1(\kappa) < v_u(\theta) < v_2(\kappa) \quad \Longleftrightarrow \quad \theta_2(\kappa) < \theta < \theta_1(\kappa).$$

These thresholds describe whether the husband's unconstrained violence choice is compatible with staying; they are distinct from the equilibrium stay-versus-separation cutoff $\tilde{\theta}(\kappa)$ introduced below.

Fix κ such that separation is supported for a nondegenerate range of violence levels (Proposition C.1, case 2 or 3), so that $v_1(\kappa)$ is defined. The best separation-inducing violence level is therefore $v_1(\kappa)$; define

$$V^L(\kappa) := \Pi^L(v_1(\kappa), \kappa).$$

Say that the husband's unconstrained choice $v_u(\theta)$ is *compatible with staying under* κ if $v_u(\theta) \leq v_1(\kappa)$, or if $v_2(\kappa)$ exists and $v_u(\theta) \geq v_2(\kappa)$. Define his best stay payoff as

$$V^S(\theta;\kappa) := \begin{cases} \Pi^S(v_u(\theta), \theta), & \text{if } v_u(\theta) \text{ is compatible with staying under } \kappa, \\ \max\{\Pi^S(v_1(\kappa), \theta), \Pi^S(v_2(\kappa), \theta)\}, & \text{if } v_2(\kappa) \text{ exists and } v_1(\kappa) < v_u(\theta) < v_2(\kappa), \\ \Pi^S(v_1(\kappa), \theta), & \text{if } v_2(\kappa) \text{ does not exist and } v_u(\theta) > v_1(\kappa). \end{cases}$$

This is well defined because $\Pi^S(\cdot, \theta)$ is strictly concave with unique maximizer $v_u(\theta)$ (Lemma C.4). Whenever $v_u(\theta)$ is not compatible with staying, the best choice on each connected component of the stay region is the boundary closest to $v_u(\theta)$ on that component; the husband then compares the payoffs at those boundaries.

Lemma C.5 (Single crossing in husband type)**.** *Fix κ such that separation is supported for a nondegenerate range of violence levels. Then $V^S(\theta;\kappa)$ is continuous and strictly decreasing in θ, so*

$$D(\theta;\kappa) := V^S(\theta;\kappa) - V^L(\kappa)$$

is continuous and strictly decreasing in θ. Hence there is at most one cutoff $\widetilde{\theta}(\kappa) \in [\underline{\theta}, \overline{\theta}]$ satisfying $D(\widetilde{\theta}(\kappa);\kappa) = 0$. If an interior cutoff exists, types below it strictly prefer staying and *types above it strictly prefer attempted separation. If no cutoff exists, all husband types strictly prefer the same branch.*

Proof. By construction, $V^S(\theta;\kappa)$ evaluates the continuous function $\Pi^S(\cdot,\theta)$ at one or two fixed points ($v_1(\kappa)$ and, when it exists, $v_2(\kappa)$) or at $v_u(\theta)$, and $v_u(\theta)$ is itself continuous; hence $V^S(\theta;\kappa)$ is continuous. For strict monotonicity, let $\theta_2 > \theta_1$ and let $\widehat{v}_2 \in \{v_u(\theta_2), v_1(\kappa), v_2(\kappa)\}$ be the point attaining $V^S(\theta_2;\kappa)$. In every case $\widehat{v}_2 > 0$ (by Lemma C.4 or because $v_1(\kappa) > 0$), so $c(\widehat{v}_2) > 0$. Using $\widehat{v}_2$ as a feasible point for type θ_1,

$$V^S(\theta_1;\kappa) \geq V^S(\theta_2;\kappa) + (\theta_2 - \theta_1)c(\widehat{v}_2) > V^S(\theta_2;\kappa).$$

Because $V^L(\kappa)$ does not depend on θ, the remaining claims follow. □

Proposition C.2 (Equilibrium violence choice)**.** *Fix a divorce cost κ.*

1. *If separation is never supported (Proposition C.1, case 1), the husband chooses $v^*(\theta;\kappa) = v_u(\theta)$ for every θ.*

2. *Otherwise, $D(\theta;\kappa) > 0$ implies a stay outcome, while $D(\theta;\kappa) < 0$ implies attempted separation with $v^*(\theta;\kappa) = v_1(\kappa)$. At $D(\theta;\kappa) = 0$, both outcomes are equilibrium best responses. When an interior cutoff $\widetilde{\theta}(\kappa)$ exists, these cases correspond respectively to $\theta < \widetilde{\theta}(\kappa)$, $\theta > \widetilde{\theta}(\kappa)$, and $\theta = \widetilde{\theta}(\kappa)$.*

3. *Conditional on a stay outcome, if $v_u(\theta)$ is compatible with staying under κ, then $v^*(\theta;\kappa) = v_u(\theta)$. If instead separation reaches the upper endpoint (Proposition C.1, case 3) and $v_u(\theta) > v_1(\kappa)$, every such constrained stayer chooses $v^*(\theta;\kappa) = v_1(\kappa)$; there is no high-violence alternative.*

4. *If a high-violence stay region exists (Proposition C.1, case 2) and $v_1(\kappa) < v_u(\theta) < v_2(\kappa)$, a constrained stayer chooses*

$$v^{S*}(\theta;\kappa) = \begin{cases} v_2(\kappa), & \theta < \theta^\dagger(\kappa), \\ \{v_1(\kappa), v_2(\kappa)\}, & \theta = \theta^\dagger(\kappa), \\ v_1(\kappa), & \theta > \theta^\dagger(\kappa), \end{cases}$$

where

$$\theta^\dagger(\kappa) = \frac{S[\mu(v_2(\kappa)) - \mu(v_1(\kappa))]}{c(v_2(\kappa)) - c(v_1(\kappa))}.$$

Per C.4, the wife-side knife edge $G(\overline{v},\kappa) = 0$—at which $\overline{v}$ is an isolated stay point in addition to the lower stay branch—is excluded from the generic characterization above.

Proof. The candidate values $V^S(\theta;\kappa)$ and $V^L(\kappa)$ are attainable equilibrium values. A maximizer of the stay problem cannot profitably deviate to another stay-compatible violence level; if $V^S > V^L$, deviations inducing separation are also unprofitable. Conversely, if $V^L > V^S$, the lowest separation-inducing level $v_1(\kappa)$ is optimal because Π^L is decreasing in violence. At equality, both outcomes can be supported because the wife's best response is set-valued at the boundary.

If $v_u(\theta)$ is compatible with staying, strict concavity makes it the unique best stay choice. If separation reaches the upper endpoint, the only stay-compatible region is $[0, v_1(\kappa)]$; for $v_u(\theta) > v_1(\kappa)$, $\Pi^S(\cdot,\theta)$ is increasing throughout this interval, so the constrained optimum is $v_1(\kappa)$.

In the bounded case, when $v_1(\kappa) < v_u(\theta) < v_2(\kappa)$, strict concavity implies that the best choice on the lower stay region is $v_1(\kappa)$ and on the upper stay region is $v_2(\kappa)$. Their payoff difference is

$$S[\mu(v_2(\kappa)) - \mu(v_1(\kappa))] - \theta[c(v_2(\kappa)) - c(v_1(\kappa))],$$

which is zero exactly at $\theta^\dagger(\kappa)$ and has the stated sign on either side. □

Whenever a nondegenerate separation range is supported under the low-cost regime, define the type-space threshold

$$\widehat{\theta} := \frac{S\mu'(v_1(\kappa_L))}{c'(v_1(\kappa_L))}.$$

By the first-order condition characterizing $v_u(\theta)$ and the strict monotonicity of $\mu'(v)/c'(v)$,

$$\theta < \widehat{\theta} \iff v_u(\theta) > v_1(\kappa_L), \qquad \theta > \widehat{\theta} \iff v_u(\theta) < v_1(\kappa_L).$$

Thus $\widehat{\theta}$ is the type threshold at which the husband's unconstrained violence choice crosses the low-cost lower separation threshold. It determines the sign of the violence response for couples induced by the reform to separate.

C.6 Type cutoffs and the divorce-cost reform

Proposition C.3 (Effect of lower divorce costs on separation)**.** *Suppose $\kappa_L < \kappa_H$ and exclude the wife-side knife edges in C.4.*

1. ***Violence thresholds.*** *If a nondegenerate separation range is supported under κ_H, then*

$$v_1(\kappa_L) < v_1(\kappa_H).$$

If both regimes have a bounded separation range, then in addition

$$v_1(\kappa_L) < v_1(\kappa_H) < v_2(\kappa_H) < v_2(\kappa_L).$$

2. ***Husband types.*** *Every husband type that strictly prefers attempted separation under κ_H also strictly prefers it under κ_L. Thus the set of types whose equilibrium involves attempted separation weakly expands. When both regimes have interior type cutoffs,*

$$\widetilde{\theta}(\kappa_L) < \widetilde{\theta}(\kappa_H).$$

3. ***Support for staying at high violence.*** *Lower divorce costs cannot make staying optimal at $\bar{v}$ if it was not already optimal there under the high-cost regime. Generically, exactly one of three endpoint configurations obtains:*

   (a) $G(\bar{v}, \kappa_L) > 0$: staying is strictly optimal at $\bar{v}$ under both regimes;

   (b) $G(\bar{v}, \kappa_L) < 0 < G(\bar{v}, \kappa_H)$: the reform eliminates staying as a best response at $\bar{v}$;

   (c) $G(\bar{v}, \kappa_H) < 0$: attempted separation is strictly optimal at $\bar{v}$ under both regimes.

   Whenever the low-cost regime has a nondegenerate separation range, the first configuration is the one in which a high-violence stay region survives under κ_L; in the second and third configurations, no such region exists under κ_L. The reverse transition—from unavailable under κ_H to available under κ_L—is impossible. This concerns the wife's best response, not whether a particular husband chooses coercive retention.

Proof. Violence thresholds. By Lemma C.3, a high-cost boundary becomes a strict separation point under the low-cost regime. Applying this at $v_1(\kappa_H)$ gives $v_1(\kappa_L) < v_1(\kappa_H)$. If both regimes are bounded, applying the same argument at $v_2(\kappa_H)$ gives $v_2(\kappa_H) < v_2(\kappa_L)$.

Husband types. If separation is never supported under κ_H, the claim is vacuous. Otherwise, every violence level compatible with staying under κ_L is also compatible with staying under κ_H by Lemma C.3, so

$$V^S(\theta;\kappa_L) \leq V^S(\theta;\kappa_H).$$

Moreover, Π^L is strictly decreasing in v and $v_1(\kappa_L) < v_1(\kappa_H)$, while differentiating (C.4) with respect to κ gives

$$\Pi^L_\kappa(v,\kappa) = -\frac{1}{2} - \frac{\Delta(v,\kappa)}{2\bar{\varepsilon}} = -p(v,\kappa) < 0.$$

Hence, since Π^L falls in both v and κ,

$$V^L(\kappa_L) = \Pi^L(v_1(\kappa_L),\kappa_L) > \Pi^L(v_1(\kappa_H),\kappa_H) = V^L(\kappa_H).$$

Therefore

$$D(\theta;\kappa_L) < D(\theta;\kappa_H)$$

for every θ. Any type strictly preferring separation under κ_H therefore also strictly prefers it under κ_L. If both cutoffs are interior, strict single crossing (Lemma C.5) implies $\widetilde{\theta}(\kappa_L) < \widetilde{\theta}(\kappa_H)$.

Support for staying at high violence. By Lemma C.3,

$$G(\bar{v}, \kappa_L) < G(\bar{v}, \kappa_H),$$

which yields the three generic endpoint configurations and rules out the reverse transition. $\square$

C.7 Violence and probability of femicide

For compactness, write $v_j^r := v_j(\kappa_r)$ for $r \in \{L, H\}$ whenever the relevant threshold is defined. Call a husband type non-marginal if its stay-versus-separation choice is strict under both regimes. For such a type, define equilibrium femicide risk as

$$F^*(\theta;\kappa) = \begin{cases} 0, & \text{if equilibrium preserves the relationship,} \\ 1 - p(v_1(\kappa), \kappa), & \text{if equilibrium involves attempted separation.} \end{cases}$$

By Proposition C.3, no type can move from attempted separation under κ_H to staying under κ_L. Every non-marginal type therefore belongs to exactly one of three groups: stay under both regimes, switch from staying under κ_H to attempted separation under κ_L, or attempt separation under both regimes. When both type cutoffs are interior, these groups correspond respectively to

$$\theta < \widetilde{\theta}(\kappa_L), \qquad \widetilde{\theta}(\kappa_L) < \theta < \widetilde{\theta}(\kappa_H), \qquad \theta > \widetilde{\theta}(\kappa_H).$$

Proposition C.4 (Full characterization of the effect on equilibrium violence)**.** *Suppose divorce costs fall from κ_H to $\kappa_L < \kappa_H$, and consider non-marginal husband types.*

1. ***Stay under both regimes.***

   (a) *If $v_u(\theta)$ is compatible with staying under κ_L, then*

   $$v^*(\theta;\kappa_L) = v^*(\theta;\kappa_H) = v_u(\theta),$$

   so the effect is zero.

   (b) *If $v_u(\theta)$ is not compatible with staying under κ_L and no high-violence stay region survives under κ_L, the constrained choice is v_1^L, and equilibrium violence strictly falls:*

   $$v^*(\theta;\kappa_L) - v^*(\theta;\kappa_H) < 0.$$

(c) *If $v_u(\theta)$ is not compatible with staying under κ_L but a high-violence stay region survives under κ_L (so $v_1^L < v_u(\theta) < v_2^L$), then*

$$\theta > \theta^{\dagger}(\kappa_L) \implies v^*(\theta;\kappa_L) - v^*(\theta;\kappa_H) < 0,$$
$$\theta < \theta^{\dagger}(\kappa_L) \implies v^*(\theta;\kappa_L) - v^*(\theta;\kappa_H) > 0.$$

At $\theta = \theta^{\dagger}(\kappa_L)$, both low-cost boundary choices are equilibrium best responses, yielding one strictly negative and one strictly positive change.

2. ***Switch from staying to attempted separation.*** *Under the low-cost regime, $v^*(\theta;\kappa_L) = v_1(\kappa_L)$, and writing $\Delta v := v^*(\theta;\kappa_L) - v^*(\theta;\kappa_H)$, the sign of the change is completely characterized by $\widehat{\theta}$:*

$$\theta < \widehat{\theta} \implies \Delta v < 0, \qquad \theta = \widehat{\theta} \implies \Delta v = 0, \qquad \theta > \widehat{\theta} \implies \Delta v > 0.$$

If a nondegenerate separation range is supported under κ_H, the change admits the decomposition

$$\Delta v = \underbrace{\left[v_1(\kappa_H) - v^*(\theta;\kappa_H)\right]}_{\text{switching component}} + \underbrace{\left[v_1(\kappa_L) - v_1(\kappa_H)\right]}_{\text{de-escalation component}<0}.$$

Thus one component of the change is always strictly negative even when the total change is positive.

3. ***Attempted separation under both regimes.*** *In each regime the husband chooses the lowest separation-inducing violence level, so by Proposition C.3,*

$$v^*(\theta;\kappa_L) = v_1(\kappa_L) < v_1(\kappa_H) = v^*(\theta;\kappa_H),$$

and non-lethal violence strictly falls.

Proof. Stay under both regimes. If $v_u(\theta)$ is compatible with staying under κ_L, it is also compatible under κ_H by Lemma C.3, and strict concavity makes it the unique best choice under both regimes.

Suppose instead $v_u(\theta)$ is not compatible with staying under κ_L. If no high-violence stay region survives under κ_L, Proposition C.2 gives $v^*(\theta;\kappa_L) = v_1^L$. Since v_1^L is exactly a stay-separation boundary under κ_L (i.e. $G(v_1^L,\kappa_L) = 0$), Lemma C.3's core inequality $G(v_1^L,\kappa_L) < G(v_1^L,\kappa_H)$ gives $G(v_1^L,\kappa_H) > 0$, so v_1^L is a *strict* stay point under κ_H, and a right-neighborhood of it is compatible with staying there too. Because $v_u(\theta) > v_1^L$, the stay payoff is increasing at v_1^L, so $v^*(\theta;\kappa_H) > v_1^L$, proving a strict decline.

If instead a high-violence stay region survives under κ_L, then $v_1^L < v_u(\theta) < v_2^L$, and the same argument applied at both v_1^L and v_2^L gives $G(v_j^L, \kappa_H) > 0$ for $j = 1,2$: both are strict stay points under κ_H, with compatible neighborhoods on the appropriate sides. Since the stay payoff is strictly increasing below $v_u(\theta)$ and strictly decreasing above it, every optimal choice under κ_H satisfies $v_1^L < v^*(\theta; \kappa_H) < v_2^L$ – regardless of whether a high-violence stay region itself survives under κ_H. The low-cost constrained choice is v_1^L for $\theta > \theta^\dagger(\kappa_L)$ and v_2^L for $\theta < \theta^\dagger(\kappa_L)$, so the former strictly reduces violence and the latter strictly increases it.

Switchers. Every switcher attempts separation under κ_L at $v^*(\theta; \kappa_L) = v_1(\kappa_L)$. If $\theta > \widehat{\theta}$, then $v_u(\theta) < v_1(\kappa_L)$, so $v_u(\theta)$ is compatible with staying under κ_L and hence under κ_H by Lemma C.3; thus $v^*(\theta; \kappa_H) = v_u(\theta) < v_1(\kappa_L)$, and violence rises. If $\theta = \widehat{\theta}$, then $v_u(\theta) = v_1(\kappa_L)$, and $G(v_1(\kappa_L), \kappa_H) > G(v_1(\kappa_L), \kappa_L) = 0$ by Lemma C.3, so $v^*(\theta; \kappa_H) = v_u(\theta) = v_1(\kappa_L)$ and the change is zero. If $\theta < \widehat{\theta}$, then $v_u(\theta) > v_1(\kappa_L)$; since $v_1(\kappa_L)$ is again a strict stay point under κ_H with a compatible right-neighborhood, and the stay payoff is increasing there, $v^*(\theta; \kappa_H) > v_1(\kappa_L)$, and violence falls. When a nondegenerate separation range is supported under κ_H, adding and subtracting $v_1(\kappa_H)$ yields the stated decomposition, with $v_1(\kappa_L) - v_1(\kappa_H) < 0$ by Proposition C.3.

Attempted separation under both regimes. Immediate from Proposition C.2 and the strict inequality $v_1(\kappa_L) < v_1(\kappa_H)$ (Proposition C.3). □

Proposition C.5 (Full characterization of the effect on femicide risk)**.** *Suppose divorce costs fall from κ_H to $\kappa_L < \kappa_H$, and consider non-marginal husband types.*

1. ***Stay under both regimes.*** *Femicide risk is zero under both regimes.*

2. ***Switch from staying to attempted separation.*** *Femicide risk rises strictly for every switcher:*

$$F^*(\theta; \kappa_H) = 0 < F^*(\theta; \kappa_L) = 1 - p(v_1(\kappa_L), \kappa_L).$$

3. ***Attempted separation under both regimes.*** *Femicide risk strictly falls:*

$$1 - p(v_1(\kappa_L), \kappa_L) < 1 - p(v_1(\kappa_H), \kappa_H).$$

Proof. Stay under both regimes. By Proposition C.2, the outcome is a stay outcome under both regimes, so $F^*(\theta; \kappa_L) = F^*(\theta; \kappa_H) = 0$ by definition.

Switchers. By definition of this group, the outcome is stay under κ_H (so $F^*(\theta; \kappa_H) = 0$) and attempted separation at $v_1(\kappa_L)$ under κ_L, so $F^*(\theta; \kappa_L) = 1 - p(v_1(\kappa_L), \kappa_L)$, strictly positive by the maintained interiority of p (Assumption C.3).

Attempted separation under both regimes. By Proposition C.2, $F^*(\theta;\kappa_r) = 1 - p(v_1(\kappa_r),\kappa_r)$. Since $v_1(\kappa_L) < v_1(\kappa_H)$ (Proposition C.3) and $p(v,\kappa)$ is strictly decreasing in both arguments,

$$p(v_1(\kappa_L),\kappa_L) > p(v_1(\kappa_H),\kappa_H),$$

because both the lower violence level and the lower divorce cost raise acceptance. Hence femicide risk strictly falls. $\square$

Table C.14. Complete individual-level comparative statics

Equilibrium region	**Additional condition**	**Separation**	Δv	ΔF^*
Stay under both	$v_u(\theta)$ compatible with staying under κ_L	stay $\rightarrow$ stay	0	0
Stay under both	constrained; no high-violence region under κ_L	stay $\rightarrow$ stay	< 0	0
Stay under both	constrained; high-violence region survives; $\theta > \theta^\dagger(\kappa_L)$	stay $\rightarrow$ stay	< 0	0
Stay under both	constrained; high-violence region survives; $\theta < \theta^\dagger(\kappa_L)$	stay $\rightarrow$ stay	> 0	0
Switch	$\theta < \widehat{\theta}$	stay $\rightarrow$ separate	< 0	> 0
Switch	$\theta = \widehat{\theta}$	stay $\rightarrow$ separate	0	> 0
Switch	$\theta > \widehat{\theta}$	stay $\rightarrow$ separate	> 0	> 0
Separate under both	–	separate $\rightarrow$ separate	< 0	< 0

Notes: $\Delta v := v^*(\theta;\kappa_L) - v^*(\theta;\kappa_H)$ and $\Delta F^* := F^*(\theta;\kappa_L) - F^*(\theta;\kappa_H)$. The table reports non-marginal cases. At $\theta = \theta^\dagger(\kappa_L)$, when a high-violence region survives, both low-cost boundaries are equilibrium best responses, producing one negative and one positive value of Δv. Knife edges (C.4) are excluded from the generic rows. When the type cutoffs are interior, the three regions correspond to $\theta < \widetilde{\theta}(\kappa_L)$, $\widetilde{\theta}(\kappa_L) < \theta < \widetilde{\theta}(\kappa_H)$, and $\theta > \widetilde{\theta}(\kappa_H)$, respectively.