Revisiting Unilateral Divorce, Health, and Crime

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Abstract

Stevenson and Wolfers (2006) theorized that unilateral divorce laws shifted power from one party to a distribution of power across both parties in the marriage, effectively providing an application of Coasian Bargaining. Utilizing a two-way fixed effects (TWFE) difference-in-difference estimator, they find that unilateral divorce had significantly reduced suicide rates, domestic violence, and intimate homicide. Innovations in econometric theory have raised concerns regarding the use of TWFE with differential timing in the treatment variable, leading to biased estimation. We revisit Stevenson and Wolfers (2006) with more modern estimators for suicide and intimate homicide rates and utilize appropriate estimators to examine the effect of unilateral divorce laws on suicides and intimate partner homicide rates. In contrast to the original research, we do not find significant effects of unilateral divorce on suicide and intimate homicide rates. This indicates that *on net* these laws were not beneficial to women's outcomes.

1 Motivation

The seminal work of Stevenson and Wolfers (2006) examined the impact of unilateral divorce laws on suicide, domestic violence, and intimate homicide rates Stevenson and Wolfers (2006). The mechanism by which unilateral divorces laws are theorized to impact bargaining power from one party to the middle is through Coasian Bargaining. It is argued that the shift in power would have positive impacts on outcomes, primarily for women, wherein each party would have more equalized power to leave the marriage. Women would feel empowered to leave relationships where they were dealing with thoughts of suicide, experiencing domestic violence, or possibly at risk of becoming the victim of intimate partner homicide. They estimated that unilateral divorce laws caused decreases in all three outcomes.

Although unilateral divorce laws may shift bargaining power from the holding party to the middle, it is possible that these laws may not be monotonically good or effective. In the case of no-fault divorce the decision to be divorced is only required by one party *and* the courts will not hear that a party is at fault and therefore not offset the costs to the "not-at-fault party". Similarly, many states still require cooling-off periods which can be sped up with agreements between parties, but even in the case of no-fault divorces, a party can choose to hold out on agreeing to a speedier divorce, which would continue to create an imbalance in power.

Their paper also utilized two-way fixed effects (TWFE) estimators with differential timing in treatment. In more recent years, conversations about the validity of TWFE have risen, with papers such as Goodman-Bacon (2021) calling into question the weighting of comparison control and treatment groups in an aggregate treatment effect. Within (Goodman-Bacon, 2018) he presents a case-study of decomposing Stevenson and Wolfers (2006) using different data sources. He reveals that the estimations of Stevenson and Wolfers (2006) may be incorrect because of the problematic comparison groups within the Average Treatment Effect (ATE) of TWFE.

According to a popular AI citation search-engine, Scite, this research has been cited 214 times by multiple researchers sci (2021). If this research is inappropriately estimated, as we believe it is, then it is inappropriate research to use in supporting claims that no-fault divorce laws are wholly positive for society at large. We believe that the inappropriateness of the original author's estimation strategies coupled with the potential usage of that original research to be used in a harmful way to be strong reasons to revisit this paper.

While using new estimators, we find that there is not a positive net benefit for women's intimate homicide rates and that the benefit for women's suicide rates follows 15 years after the adoption of the laws. The former indicates that there is no net benefit of unilateral divorce laws for women's intimate homicide rate. This could be positive for some women, but on net the effect is null. The latter finding provides evidence that women's positive returns on suicide rates, may not be driven solely by unilateral divorce. Decreases in suicide rates 15 years after adoption is unlikely to be driven only by unilateral divorce laws because this is an extraordinarily long time period after adoption.

This paper revisits Stevenson and Wolfers (2006) but utilizes appropriate estimators from modern econometric theory. The paper is structured as follows: section 2 is our replication of their initial findings, section 3 is a decomposition of these estimations, section 4 presents more modern estimators across suicide rates and intimate homicide rates, and section 5 concludes.

2 Replication

In this paper, we used the same data as Stevenson and Wolfers (2006) to test the effect of unilateral divorce laws on number of suicide, domestic violence, and intimate partner homicide. Particularly, the data captures the years from 1964 through to 1996, 36 states and District of Columbia adopted unilateral divorce law, and 14 states did not adopt, which was considered as the control population. Additionally, there is a differential timing of adoption of the law across the states. Data on suicide and homicide were from the National Center for Health Statistics (NCHS) and the FBI Uniform Crime Reports (UCR). We replicate the findings of the original paper in the following sub-sections.

2.1 Two-Way Fixed Effects Method

To provide evidence that our data is similar we replicate the tables in the original paper. The original method utilized in Stevenson and Wolfers (2006) is the two-way fixed effects method difference-in-differences. This method compares treatment and control groups in the pre- and post-periods. The estimation takes on the following form:

$$Outcome_{sy} = \beta_0 + \beta_1 Unilateral + \nu_y + \kappa_s + X_{sy}\theta + \epsilon_{sy}. \tag{1}$$

Each specification includes year fixed effects (ν_y), state fixed effects (κ_s), and a set of control variables (X_{sy}). The control variables included in the original estimations include the maximum AFDC rate for a family of four, the natural log of state personal income per capita, the unemployment rate, the female-to-male employment rate, age composition variables indicating the share of states' populations aged 14 –19, and then ten-year cohorts beginning with age 20 up to a variable for 90, and the share of the state's population that is Black, White, and other Stevenson and Wolfers (2006). The variable of interest, $Unilateral$, is expected to be negative and significant across all specifications.

2.2 Replications

Column 1 of Tables 1 and 2 displays the TWFE estimation with state and year FE's in the original paper. Column 2 adds controls. Columns 3 and 4 present our replication of their TWFE results. Column 5 adds clustering at the state level, which was absent in the original paper. Abadie et al. (2017) showed that the use of normal standard errors may lead to incorrect inference and using clustering at the treatment level can improve the inference Abadie et al. (2017). Our replications find the similar result that unilateral divorce had a statistically significant and negative effect on female suicides beginning eight years after adoption. Male suicides are also impacted 9-10 years after the adoption of unilateral divorce. However, these effects become insignificant when clustered standard errors were used.

Table 1: Effect Of Unilateral Divorce Laws On Female Suicide Rates

Variable	(1)	(2)	(3)	(4)	(5)
Year of Change	1.6%	1.3%	1.6%	1.1%	1.1%
	(3.8)	(3.4)	(3.8)	(3.3)	(3.5)
1-2 years later	-1.5%	-1.4%	-1.5%	-1.2%	-1.2%
	(3.7)	(3.5)	(3.7)	(3.5)	(4.8)
3-4 years later	-1.5%	-1.1%	-1.5%	-1.1%	-1.1%
	(3.1)	(3.1)	(3.1)	(3.1)	(4.2)
5-6 years later	-3.0%	-2.0%	-3.0%	-1.9%	-1.9%
	(2.9)	(2.9)	(2.9)	(2.9)	(4.2)
7-8 years later	-8.0%	-6.6%	-8.0%***	-6.5%**	-6.5%
	(3.0)	(3.0)	(3.0)	(3.0)	(4.5)
9-10 years later	-10.0%	-8.5%	-10.0%***	-8.4%***	-8.4%*
	(3.0)	(3.0)	(3.0)	(3.0)	(4.5)
11-12 years later	-11.9%	-10.2%	-11.9%***	-10.2%***	-10.2%*
	(3.1)	(3.2)	(3.1)	(3.2)	(5.2)
13-14 years later	-12.8%	-11.1%	-12.8%***	-11.0%***	-11.0%**
	(3.2)	(3.1)	(3.2)	(3.1)	(5.0)
15-16 years later	-13.3%	-11.7%	-13.3%***	-11.7%***	-11.7%**
	(3.7)	(3.6)	(3.7)	(3.6)	(5.8)
17-18 years later	-16.4%	-13.9%	-16.4%***	-13.8%***	-13.8%**
	(3.6)	(3.6)	(3.6)	(3.6)	(5.5)
>=19 years later	-18.7%	-16.4%	-18.6%***	-16.3%***	-16.3%**
	(3.2)	(3.3)	(3.2)	(3.3)	(6.7)
F-test of joint significance	p = 0.00	p = 0.00	p = 0.00	p = 0.00	p = 0.32
Average Effect	-9.7	-8.3	-9.5	-8.1	-8.1
	(2.3)	(2.3)	()	()	()
Original Paper	Yes	Yes	No	No	No
State and Year FE	Yes	Yes	Yes	Yes	Yes
Control Variables	No	Yes	No	Yes	Yes
Robust SE	Yes	Yes	Yes	Yes	Yes
Cluster	No	No	No	No	At State

Standard Errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 2: Effect Of Unilateral Divorce Laws On Male Suicide Rates

Variable	(1)	(2)	(3)	(4)	(5)
Year of Change	-0.8%	-1.4%	-0.8%	-1.4%	-1.4%
	(2.2)	(2.1)	(2.2)	2.1	(2.5)
1-2 years later	1.2%	0.5%	1.2%	0.5%	0.5%
	(1.5)	(1.4)	(1.5)	(1.4)	(2.0)
3-4 years later	0.0%	-0.9%	0.0%	-0.9%	-0.9%
	(1.6)	(1.5)	(1.6)	(1.5)	(2.1)
5-6 years later	0.4%	-0.2%	0.4%	-0.2%	-0.2%
	(1.5)	(1.5)	(1.5)	(1.5)	(2.0)
7-8 years later	-1.0%	-1.3%	-1.0%	-1.3%	-1.3%
	(1.8)	(1.8)	(1.8)	(1.8)	(2.6)
9-10 years later	-3.5%	-3.9%	-3.5%**	-3.9%**	-3.9%
	(1.7)	(1.7)	(1.7)	(1.7)	(2.7)
11-12 years later	-2.2%	-2.6%	-2.2%	-2.7%	-2.7%
	(2.0)	(2.0)	(2.0)	(2.0)	(3.2)
13-14 years later	-3.2%	-3.6%	-3.2%	-3.6%*	-3.6%
	(2.0)	(2.0)	(2.0)	(2.0)	(3.5)
15-16 years later	-1.6%	-2.0%	-1.6%	-2.0%	-2.0%
	(2.0)	(1.9)	(2.0)	(2.0)	(3.3)
17-18 years later	-1.6%	-1.9%	-1.6%	-1.9%	-1.9%
	(2.1)	(2.0)	(2.1)	(2.0)	(3.4)
>=19 years later	-3.9%	-4.3%	-3.9%*	-4.3%**	-4.3%
	(2.0)	(2.0)	(2.0)	(2.0)	(4.2)
F-test of joint significance	p = 0.36	p = 0.37	p = 0.36	p = 0.36	p = 0.59
Average Effect	-1.5	-2.0			
	(1.3)	(1.3)	()	()	()
Original Paper	Yes	Yes	No	No	No
State and Year FE	Yes	Yes	Yes	Yes	Yes
Control Variables	No	Yes	No	Yes	Yes
Robust SE	Yes	Yes	Yes	Yes	Yes
Cluster	No	No	No	No	At State

Standard Errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3 presents our replication results for the impact of no-fault divorce laws on domestic violence. Table 3 provides estimates using OLS with state fixed effects, controls, time-varying controls, and probit estimations across overall and severe violence with male and female spouses as victims. These results are nearly identical to the original paper and indicate that domestic violence decreases in overall husband to wife violence and in severe violence between both husband to wife and wife to husband domestic violence.

Table 3: Effect of Unilateral Divorce on Domestic Violence

	Overall violence		Severe Violence	
	Husband to wife	Wife to husband	Husband to wife	Wife to husband
	Average incidence of each type of violence			
	11.7%	11.9%	3.4%	4.5%
OLS(Diffs-in-diffs)	-4.3%**	-2.7%	-1.1%	-2.9%***
	(1.9)	(1.8)	(1.3)	(1.0)
Add state fixed effects	-5.5%***	-3.2%**	-2.0%**	-3.6%
	(1.8)	(1.5)	(0.9)	(0.7)
Add individual controls	-5.0%***	-1.9%	-1.8%*	-3.4%***
	(1.8)	(1.4)	(1.0)	(0.9)
Add state-level time-varying controls	-3.6%**	-1.8%	-1.8%*	-3.0%***
	(1.5)	(1.3)	(1.0)	(0.7)
Probit with individual controls	-4.7%***	-2.0%	-1.2%*	-2.1%***
	(1.6)	(1.3)	(0.7)	(0.7)

Table 4 is a replication of the results in Stevenson Wolfers (2006), which we replicate nearly identically. We find that unilateral divorce has a significant reductions on intimate homicide against women with and without controls. This result holds across violence from spouses, family members, and known assailants. These effects do not extend to non-intimate homicides (defined as the homicide rate less the intimate homicide rate or triple difference homicides (intimate less non-intimate homicide rates). We do not estimate significant effects for men except when including controls and estimating the effect of unilateral divorce on murders of men by known assailants

Table 4: Effect of Unilateral Divorce on Intimate Homicide

	No Controls	Including Controls		
	Intimate homicide	Intimate homicide	Placebo nonintimate homicide	Diffs-in-Diffs-in-Diffs (intimate less nonintimate)
	(1)	(2)	(3)	(4)
Women Murdered by Intimates				
By Spouse	-10.5*	-11.8**	-4.3	-7.55
	(5.7)	(5.8)	(3.5)	(6.3)
By Family	-8.9**	-9.1**	-3.4	-5.7
	(4.2)	(4.4)	(4.1)	(5.5)
By Known	-8.7**	-8.9**	-3.2	-8.6
	(3.5)	(3.6)	(5.2)	(5.6)
Men Murdered by Intimates				
By Spouse	12.3	4.4	-4.0	8.4
	(8.9)	(8.5)	(2.6)	(8.3)
By Family	1.8	-4.1	-3.4	-.72
	(5.1)	(5.2)	(2.7)	(5.1)
By Known	-2.0	-5.8*	-2.8	-5.5
	(3.0)	(3.0)	(4.1)	(4.7)

clustered standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

3 Decomposition of Results

Goodman-Bacon (2021) proposed a method to decompose the average treatment effect on the treated (ATT) into different comparison groups. He identifies that certain groups (Later treated observations versus any treatment level) are problematic treatment group comparisons. Intuitively, this is because these treatment groups are comparing samples that are newly treated to samples that have been treated before. This is inappropriate because difference-in-differences econometric design should only produce estimations with comparison groups that are treated and untreated.

In addition, two other groups, which compare the earlier treatment group with the later control group and the later treatment group with the earlier control, impose stronger parallel trend assumption across time between different groups. This indicates that these comparison groups are also inappropriate because these parallel trends assumptions may not hold up, indicating an inappropriate ATE estimation.

It is also possible to think of a scenario where the heterogeneity of treatment effects may effect the estimate. If early adopting states have a smaller effect than late adopters, then the estimated effect between these two groups could be biased downwards.

The Goodman-Bacon decompositions shows the ATTs for each group and the weights assigned to them. If these inappropriate comparison groups have more weight, then the ATT should be seen with caution and is more likely to be biased. We used the Goodman-Bacon decomposition to see the distribution of ATTs of unilateral divorce on suicide rates with and without co-variates and intimate homicide with and without co-variates. We find that the ATTs from the comparison of inappropriate groups are assigned higher weights, indicating that the original treatment estimations are biased.

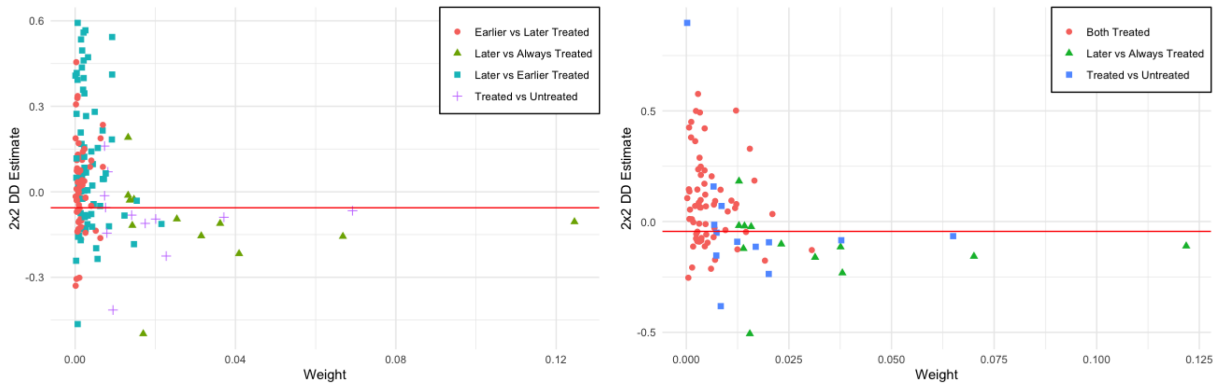
3.1 Suicide

For each of the following figures, the red line represents the ATE of TWFE estimations and each of the points represents an ATT of a different group-time cohort comparison group. These group-time cohorts are represented by different types of points for what kind of comparison group it is.

"Earlier versus Later Treated" and "Treated versus Untreated" groups are intuitive to TWFE. In each of these comparison types we're comparing treated units to either untreated or not-yet treated units. This akin to treated versus untreated units. "Later versus Always Treated" and "Later versus Earlier Treated" comparison groups are comparing two types of treated units and are inappropriate estimation groups. In these we would be comparing treated units to control units that are either always treated or treated before the treatment unit.

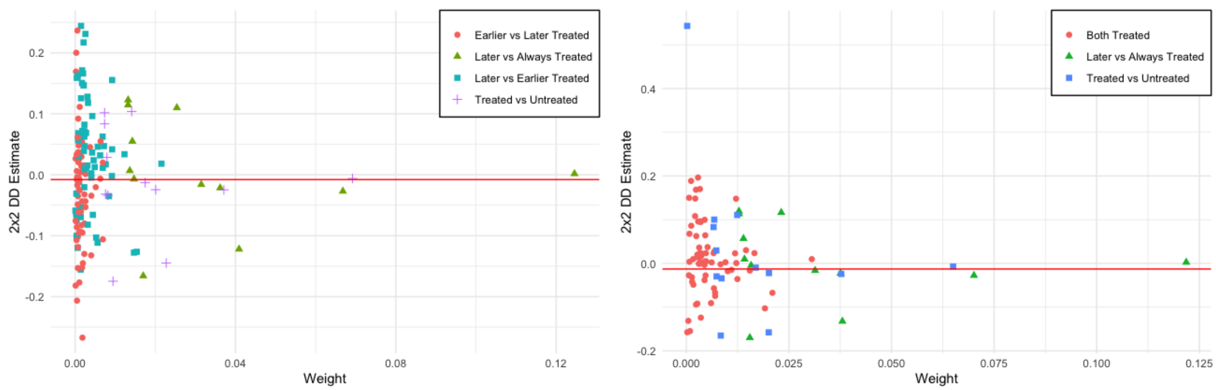
Figures 1 and 2 provide graphical representations of the group-time effects. As can be visually observed, the original estimate is influenced by biased groups that are pulling the estimate negative.

Appendix Tables A.1 and A.2 provide decompositions of effects of female and male suicide rates with and without controls, respectively. We find that the effects are driven by mostly "Later versus Always" and "Later versus Earlier Treated", as well as "Both Treated". This indicates that the original estimated effects are biased and may not indicate the correct effect of no-fault divorce on suicide rates.



Without co-variates　　　　With co-variates

Figure 1: Goodman-Bacon Decompositions by Group-Time for Female Suicide Rates



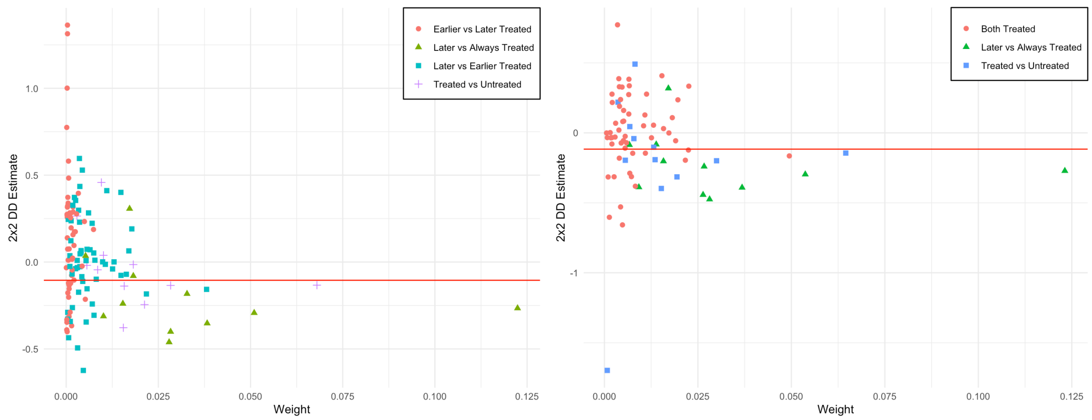
Without co-variates　　　　With co-variates

Figure 2: Goodman-Bacon Decompositions by Group-Time for Male Suicide Rates

3.2 Intimate Homicides

Figure 3 presents visualizations of ATTs by group-time cohorts. The initially estimated effects of -10.5 percentage points without co-variates and -11.8 percentage points with co-variates are presented as red lines in each figure. Both figures indicate that some inappropriate groupings are being given heavier weights and therefore are biasing the estimated results.

The "Later vs Earlier" Treated group receives the bulk of the weight in the estimated ATE when excluding co-variates. The "Earlier vs Later" Treated group receives most of the weight when including co-variates. The prevailing belief of Stevenson and Wolfers (2006) is that the estimation should be driven by treated versus untreated comparison groups, but we find that most of the estimation is driven by groups that bias estimation. This indicates that estimated effects may be driven by comparison groups that are inappropriate and require that we estimate the impact of unilateral divorce on intimate homicide using different estimators. Appendix table A.3 and A.4 present the decompositions of the effect into groups by weight and average estimate.



Without co-variates — With co-variates

Figure 3: Goodman-Bacon Decompositions by Group-Time for Female Intimate Homicide Rates

In sum, the results of the decomposition indicate that the TWFE estimators are being influenced by "Later versus Always Treated" and "Later versus Earlier Treated". This indicates the need for research to re-examine the impact of unilateral divorce laws on suicides and intimate partner homicides. In the next section, we propose new estimators to be used to estimate the true effect of no-fault divorce law adoption on suicide and intimate partner homicide.

4 New Estimators

The logic of the two-way fixed effect model does not naturally extend to differential timings Goodman-Bacon (2018); Imai and Kim (2020); Chaisemartin and D'haultfoeuille (2021). From the derivation of the linear regression, it provides us a variance weighted approximation. However, in the panel setting with staggered timing, these weights are not always appropriate. The model can put heavier weights on inappropriate comparison groups, which calls into question the interpretability of the ATE.

Recent work in econometric theory has proposed new estimation techniques for differential timing difference-in-differences, as opposed to TWFE. (Callaway and C Sant, 2020) propose an inverse probability weighting estimator which doesn't include post-treatment controls for treated units and only compares treated and either untreated or not-yet treated units for estimated ATTs. (Sant'Anna and Zhao, 2018) proposed a doubly robust estimator that combines propensity scores and linear regression providing two opportunities to correctly specify your estimation. We utilize the *did* package in R to estimate the effects of unilateral divorce on suicide and intimate homicide rates using the IPW and doubly robust estimators did (2021).

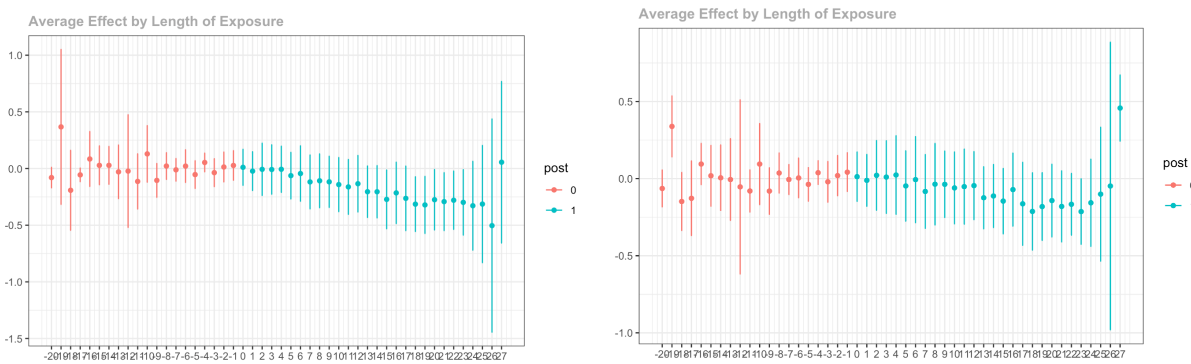
4.1 Suicide

Table 5 presents our estimation of unilateral divorce on female suicide rates with the Inverse Probability Weighting Estimator. This estimation doesn't include any co-variates. From this table we can discern that the unilateral divorce law changes had significant impacts on female suicide rates with a significant ATT of -15.23 percent, however this is without controls. However, when including controls, the ATT is insignificant.

Table 5: Group ATT of Effect of Unilateral Divorce on Female Suicide Rates

	(1)	(2)	(3)	(4)
Group	ATT	ATT	ATT	ATT
1969	-0.0984	0.0944	-0.0963*	-0.0475
1970	-0.5623	-0.3203	-0.1607	-0.0081
1971	-0.1896	-0.1531	-0.0269	-0.0182
1972	-0.1452	-0.1003	-0.0450	-0.0151
1973	-0.0834	-0.0577	-0.0211	0.0012
1974	-0.1358	-0.0018	-0.0360	0.0178
1975	-0.0798	-0.0029	-0.0398	-0.0098
1976	-0.2483*	0.0005	-0.0605	-0.0152
1977	-0.2696	-0.0994	-0.1532	-0.1038
1980	-0.1347	-0.1502	0.0229	0.0913
1984	-0.1510*	-0.1307	-0.0224	0.0109
1985	0.2763*	0.3690*	-0.0636*	-0.0576
Aggregate ATT	-0.1523*	-0.0724	-0.0497	-0.014
Gender	Female	Female	Male	Male
co-variates	No	Yes	No	Yes

Figure 4 we examine the effect of unilateral divorce on female suicide by length of exposure. Each point represents an estimate of a year either before or after treatment. Suicide rates are significantly impacted by treatment 15 years after being treated. This is a larger window that initially estimated by Stevenson and Wolfers (2006). Their initial findings and our replication find that suicide rates only become statistically significantly different from zero eight years after the law change. Therefore, other unknown factors outside the scope of Coasian argument might be in play here threatening the validity of the estimators.



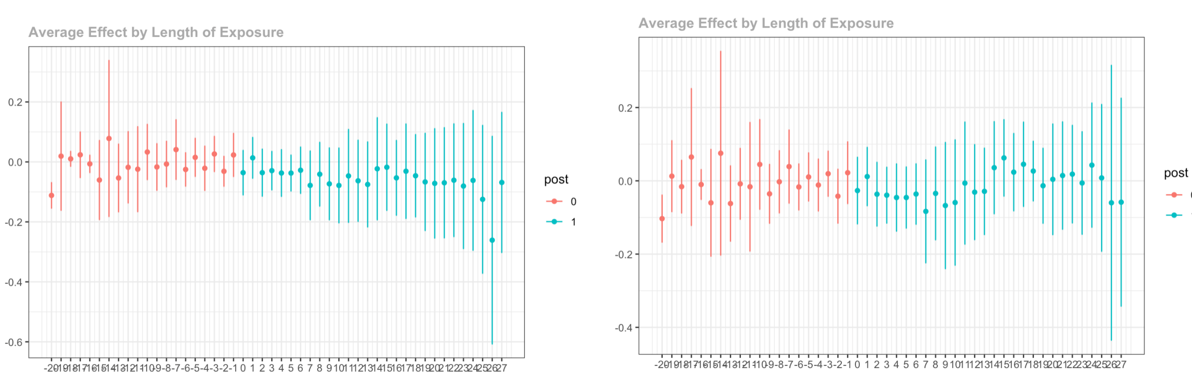
Without co-variates

With co-variates

Figure 4: Effect of Unilateral Divorce on Female Suicide by Year

Table 6 presents an estimation using a regression method with our own manual estimations of propensity scores on treatment. The *did* package in R only allows for a small number of co-variates with the inverse probability weighting and doubly robust methods, therefore we use a regression method to check that null results aren't driven by differences in co-variates. This method mimics a doubly robust estimation because we specify a propensity score method and then a linear regression method but are not constrained. From this table we still discern that the unilateral divorce law changes had no significant impact on female suicide rates when controlling for co-variates. The estimated ATT of unilateral divorce laws on female suicide rate is -4.97 percent and is insignificant.

Figure 5 presents the effect of unilateral divorce on male suicide by length of exposure while controlling for co-variates. We don't find effects for male suicide rates. This could be due to the fact that male suicide rates are already high comparatively and no-fault divorce wouldn't be strong enough to impact it.



Without co-variates

With co-variates

Figure 5: Effect of Unilateral Divorce on Male Suicide by Year

We decompose the effect of divorce laws on female suicide rates by age group using CDC Wonder data.

Our new estimators do not reveal any significant effects of divorce laws on female suicide rates when we include controls. But we do find statistically significant effect when the controls were not used. Table 6 provides evidence that the estimated significant effects are positive and driven mostly by young women (under age of 20), along with some cohorts of 20-54 year old women. This raises suspicion on the estimators as significant effect of divorce laws on younger women that are less likely to be married does not support Coasian argument. It is highly likely that these young women are experiencing positive externalities of culture change leading lower suicide rate among them rather than the input of no-fault divorce.

Table 6: Effect Of Unilateral Divorce Laws On Female Suicide Rates by Age Group

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Less than 20	20 - 34	35 - 44	45 - 54	55 - 64	65 - 74	75 - 84	Greater than 84
>=1 year later	-2.10e-05	0.000128	-0.000222	0.000243**	0.000200	-6.35e-05	-6.05e-05	3.99e-05
	(7.09e-05)	(0.000205)	(0.000146)	(0.000110)	(0.000204)	(0.000133)	(7.99e-05)	(2.91e-05)
1-2 years later	0.000249***	0.000244	-9.97e-05	0.000262*	3.62e-05	-8.47e-05	4.24e-05	3.42e-05
	(8.61e-05)	(0.000155)	(0.000122)	(0.000144)	(0.000158)	(9.60e-05)	(7.95e-05)	(4.22e-05)
3-4 years later	4.91e-05	0.000175	-0.000142	0.000150	6.64e-05	9.43e-05	4.38e-05	1.91e-06
	(7.27e-05)	(0.000158)	(0.000153)	(0.000121)	(0.000116)	(8.71e-05)	(7.65e-05)	(2.76e-05)
5-6 years later	0.000110	0.000332**	-7.06e-05	0.000281**	6.86e-05	7.67e-05	8.89e-05	3.74e-05
	(6.84e-05)	(0.000165)	(0.000124)	(0.000117)	(0.000103)	(8.30e-05)	(6.79e-05)	(2.92e-05)
7-8 years later	0.000190**	0.000361**	-0.000201	0.000170	5.59e-05	0.000132	6.41e-05	2.05e-05
	(7.42e-05)	(0.000149)	(0.000123)	(0.000118)	(0.000106)	(8.52e-05)	(6.83e-05)	(2.97e-05)
9-10 years later	0.000254***	0.000338**	-0.000189	0.000125	6.35e-05	0.000123	4.80e-05	5.17e-06
	(7.78e-05)	(0.000149)	(0.000129)	(0.000126)	(0.000110)	(8.69e-05)	(6.92e-05)	(2.92e-05)
11-12 years later	0.000254***	0.000271*	-0.000182	0.000227*	-8.70e-06	0.000160*	4.07e-05	3.06e-05
	(7.66e-05)	(0.000157)	(0.000140)	(0.000132)	(0.000111)	(9.11e-05)	(7.15e-05)	(3.08e-05)
13-14 years later	0.000292***	0.000249	-0.000180	0.000190	3.08e-06	0.000152	8.03e-05	3.33e-05
	(8.47e-05)	(0.000166)	(0.000149)	(0.000146)	(0.000117)	(9.56e-05)	(7.53e-05)	(3.17e-05)
15-16 years later	0.000285***	0.000353**	-0.000167	0.000173	-6.81e-06	0.000164	6.23e-05	3.81e-06
	(8.77e-05)	(0.000171)	(0.000158)	(0.000153)	(0.000121)	(9.98e-05)	(7.87e-05)	(3.17e-05)
17-18 years later	0.000299***	0.000297	-0.000250	0.000175	-0.000103	0.000190*	4.09e-05	6.34e-06
	(8.90e-05)	(0.000184)	(0.000180)	(0.000164)	(0.000128)	(0.000106)	(8.01e-05)	(3.25e-05)
>=19 years later	0.000285***	0.000254	-0.000246	0.000221	-8.55e-05	0.000198*	1.64e-05	2.41e-06
	(9.12e-05)	(0.000195)	(0.000175)	(0.000166)	(0.000130)	(0.000107)	(8.01e-05)	(3.26e-05)
Constant	-0.00101	0.000468	0.0190*	-0.00864	-0.000776	-4.06e-05	0.000886	-0.00174
	(0.00660)	(0.0143)	(0.0112)	(0.00962)	(0.00849)	(0.00702)	(0.00471)	(0.00166)
F-Test	0.0035	0.3033	0.8114	0.1500	0.3298	0.2814	0.3273	0.0575
Observations	785	785	785	785	785	785	785	785
R-squared	0.508	0.608	0.466	0.420	0.521	0.433	0.487	0.475

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

4.2 Intimate Homicide

Table 7 presents an estimation using the IPW method. This estimation does not include any controls. From Table 7 we discern that the unilateral divorce law changes had no significant impacts on the variables of interest, intimate homicide on women by spouses, family, or known assailants, placebos, or the triple difference dependent variable across men or women.

Table 7: Effect of Unilateral Divorce on Intimate Homicide - IPW Estimator

	Intimate homicide (1)	Placebo nonintimate homicide (2)	Diffs-in-Diffs-in-Diffs (intimate less nonintimate) (3)
		Women Murdered by Intimates	
By Spouse	-4.8	-12.1	7.2
	(16.1)	(10.5)	(15.8)
By Family	-19.2	-2.8	-16.4
	(16.5)	(9.5)	(17.5)
By Known	-6.3	-18.8	12.5
	(14.1)	(10.2)	(16.9)
		Men Murdered by Intimates	
By Spouse	15.3	-9.0	24.3*
	(13.8)	(10.9)	(13.4)
By Family	-18.0	-5.2	-12.8
	(12.7)	(14.0)	(19.2)
By Known	2.1	-21.5	23.7
	(16.3)	(29.4)	(43.0)

Bootstrapped standard errors in parentheses
*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 8 presents estimations of the Sant'anna-Zhao doubly robust method without controls. The estimates of this method don't change from the inverse probability weighting method. From this table we discern that the unilateral divorce law changes had no significant impacts on the variables of interest, intimate partner homicide on women by spouses, family, or known assailants, placebos, or the triple difference dependent variable across men or women. Estimates from the linear regression method are not included but also find insignificant effects across all specifications.

Table 8: Effect of Unilateral Divorce on Intimate Homicide - Doubly Robust Estimator

	Intimate homicide (1)	Placebo nonintimate homicide (2)	Diffs-in-Diffs-in-Diffs (intimate less nonintimate) (3)
	Women Murdered by Intimates		
By Spouse	-4.8	-12.1	7.2
	(15.7)	(10.3)	(15.3)
By Family	-19.2	-2.8	-16.9
	(15.7)	(9.7)	(16.7)
By Known	-6.3	-18.8*	-16.4
	(14.0)	(9.5)	(15.9)
	Men Murdered by Intimates		
By Spouse	15.3	-9.0	24.3*
	(14.0)	(10.1)	(12.8)
By Family	-18.0	-5.2	-12.8
	(11.9)	(13.2)	(21.6)
By Known	2.1	-21.5	23.7
	(14.5)	(30.3)	(40.7)

Bootstrapped standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

5 Conclusion

Stevenson and Wolfers (2006) published a paper that theorized that unilateral divorce laws shifted bargaining rights to the middle for couples. They used TWFE and found that unilateral divorce laws decreased female suicide rates 8 years after law adoption, decreased domestic violence, and decreased the murder of women by intimate contacts. This paper has been cited and used in support of a multitude of research, which indicates that it's an important piece of research. We have revisited their data and research question with more modern econometric techniques to study whether or not these effects are still precisely measured today. We find that although we replicated their results well using TWFE, we did not find any effects of unilateral divorce on suicide rates or intimate homicide with more appropriate estimators. This indicates that the original results were contaminated by TWFE estimations and should not be relied on for understanding the true effects of unilateral divorce on health and crime.

Within our paper, we provide a decomposition of TWFE model using the Goodman-Bacon's (2020) method and find that there are problematic weights associated with control and treatment comparison groups in the initial analysis. These inappropriately measured control and treatment groups biased the initial results and pulled their estimates downward indicating strong effects of unilateral divorce on outcomes. Our results do not necessarily prove that unilateral divorce laws and ineffective in preventing violence for *all* women, but rather that *on net* these laws are not necessarily effective. This indicates that there may be some women that were positively affected by these laws, and some women that were negatively affected. For example, many states have "cooling-off" periods, wherein an upset spouse can still keep the divorcing party beholden trapped in a marriage past their filing date.

Using more Callaway and Sant'anna (2020) methods nullifies the effects presented in the original paper. Similarly, (Stevenson and Wolfers, 2007), expressed how culture was changing at the same time as these divorce laws were being adopted. This indicates that the initial estimates, although biased, could have captured a change in culture rather than adoption of unilateral divorce laws.

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6 Appendix

Table A1: Decomposition of Effect on Female Suicide Rates

Type	Weight	Average Estimate	Weight	Average Estimate
Earlier vs Later Treated	10.62	2.3	-	-
Later vs Always Treated	41.12	-12.72	40.70	-13.04
Later vs Earlier Treated	25.41	6.18	-	-
Treated vs Untreated	22.85	-9.57	21.76	-9.24
Both Treated	-	-	37.54	5.73
Aggregate ATT	100	-5.60	100	-4.42
	Without co-variates	Without co-variates	With co-variates	With co-variates

Table A2: Decomposition of Effect on Male Suicide Rates

Type	Weight	Average Estimate	Weight	Average Estimate
Earlier vs Later Treated	10.62	-3.3	-	-
Later vs Always Treated	41.12	-0.97	40.70	-1.0
Later vs Earlier Treated	25.41	1.59	-	-
Treated vs Untreated	22.85	-1.96	21.76	-1.88
Both Treated	-	-	37.54	-0.3
Aggregate ATT	100	-0.79	100	-1.28
	Without co-variates	Without co-variates	With co-variates	With co-variates

Table A3: Decomposition of Effect on Female Spousal Homicide Rate

Type	Weight	Average Estimate
Earlier vs Later Treated	7.13	10.43
Later vs Always Treated	36.68	-25.62
Later vs Earlier Treated	35.81	0.867
Treated vs Untreated	20.38	-10.46
Aggregate ATT	100	-10.50

Table A4: Decomposition of Effect on Female Spousal Homicide Rate with co-variates

Type	Weight	Average Estimate
Earlier vs Later Treated	45.37	1.95
Later vs Always Treated	35.75	-27.49
Treated vs Untreated	18.88	-15.37
Aggregate ATT	100	-11.80