The Effects of Waiting Periods on Firearm Suicides in the U.S.

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Abstract

Suicide is often an impulsive act, and in the United States, nearly one-half of suicides involve a firearm, the most lethal and readily available method. In this paper, we use recent developments in difference-in-differences methodology to study the causal effect of waiting-period laws on firearm suicides. We find that waiting periods reduce firearm suicides among men by 1.3 deaths per 100,000 population, a statistically significant effect at the 90% confidence level. For white individuals, we observe a statistically significant reduction of 21.6 deaths per 100,000 (p-value = 0.09). While the reductions for the overall population (0.47 per 100,000) and adults aged 55 and older (26.3 per 100,000) are not statistically significant, the majority of confidence intervals lie in the negative range, suggesting protective effects. Crucially, we find no evidence of substitution toward non-firearm suicide methods following waiting period adoption. We also examine the effects of waiting period repeal and find no statistically significant increases in firearm suicides, which may indicate persistent protective effects even after policy removal. These findings provide evidence that even brief delays in firearm access can disrupt the pathway from suicidal ideation to death, suggesting that cooling-off periods may be an important policy tool for suicide prevention. **JEL**: I18; I12; K32; J17; H75.

Keywords: Firearm waiting periods; suicide prevention; gun policy; public health; difference-in-differences; event-study design.

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1 Introduction

Suicide claims nearly one million lives worldwide each year and is often an impulsive act (Lewiecki and Miller 2013). In the United States, firearms—particularly handguns—account for over half of all gun-related fatalities and more than half of all suicides, and the economic burden associated with suicide is substantial (Greenberg et al. 2015; Greenberg et al. 2021; National Center for Health Statistics 2007). Since gunshots are highly lethal and require little planning, policies that introduce a barrier between purchase and possession may be uniquely positioned to save lives. Waiting period laws create such a barrier, giving suicidal intent time before committing suicide using firearms. Therefore, it is important to causally identify whether waiting periods are effective in preventing firearm suicide.

Emerging research in economics and public health further reinforces that suicide is not solely a function of long-standing mental illness, but is highly responsive to acute shocks and the availability of means. Economic hardship—such as job loss, income volatility, or relative status decline—has been shown to significantly increase suicide rates (Breuer 2014; Christian, Hensel, and Roth 2019; Daly, Wilson, and Johnson 2013). The results support the idea that the implementation of policy measures aimed at impulsive moments, such as waiting periods for firearm purchases, can significantly influence the outcomes by disrupting critical time frames that could otherwise lead to deadly actions.

Moreover, impulsive-aggressive behavior has consistently been identified as a risk factor for suicide across the life course, with adolescents and young adults being particularly vulnerable (Anestis et al. 2014; McGirr et al. 2008). Many suicide attempts are driven by transient states of hopelessness and distress rather than long-term ideation. Restricting immediate access to lethal means could create a critical window in which intense suicidal urges can fade, and lifesaving intervention can be provided.

While many studies examine gun control policies and suicide rates, few identify causal effects. The empirical literature notably lacks causal analyses of US waiting-period laws on firearm suicides. The existing literature is generally limited to individual states, short periods following legislative actions, or relies on broad national metrics, thereby complicating efforts to disentangle causal estimates. To our knowledge, this will consequently be the first paper aiming to causally estimate the effect of waiting periods on suicides.

A substantial body of evidence links easy access to firearms with an increased risk of suicide. International comparisons find strong correlations between household gun ownership and suicide rates, with no signs that people simply switch to other means when guns are less available (Killias 1993). In the United States, Grossman et al. (2005) show that unloaded guns and separate ammunition storage are

associated with markedly lower odds of suicide by youth. These findings align with clinical observations that many suicide attempts suddenly arise during moments of acute psychological distress (Lewiecki and Miller 2013).

Cross-national policy evaluations reinforce the value of restricting rapid access to firearms. Following the tightening of gun laws in 1992, New Zealand saw a 46% decrease in firearm suicides among the general population and a 66% decrease among individuals aged 15–24 (Beautrais, Fergusson, and Horwood 2006). The 1996 Australian National Firearms Agreement, which combined large gun buybacks with stricter licensing, has also been associated with subsequent declines in firearm suicide (Baker and McPhedran 2007).[1] In the United States, the laws on firearm removal based on risk ('red flag') enacted in Connecticut and Indiana were followed by measurable reductions in statewide suicide rates (Kivisto and Phalen 2018).

Firearm-related injuries and suicides among youth remain a significant concern in the United States. Chaudhary et al. (2024) find that mental health diagnoses often precede youth suicides, underscoring the need for earlier identification and intervention strategies within healthcare systems. The financial burden of firearm injuries, both fatal and nonfatal, has been substantial. Injuries are estimated to cost the healthcare system and the economy, through lost productivity, billions of dollars (Miller et al. 2024). These challenges are compounded by persistent trends of firearm-related harm in children and adolescents, emphasizing the urgent need for prevention and harm-reduction approaches (Kaufman et al. 2021; Lee et al. 2022). This urgency is magnified by psychological research that emphasizes the impulsive nature of many youth suicides, where access to firearms dramatically increases the risk of fatal outcomes (Anestis et al. 2014; McGirr et al. 2008).

Efforts to reduce firearm-related injuries must also focus on storage practices and perceptions of accessibility. Miller et al. (2025) highlight the critical role of secure firearm storage in reducing the risk of suicide, particularly in households with adolescents. However, firearm storage practices vary widely, and older adults and parents often underestimate the extent to which firearms are accessible to youth (Carter et al. 2022; Hastings et al. 2025). Even when parents report using safe storage methods, teens may still perceive firearms as accessible, suggesting that education and behavioral interventions must account for both the parental and youth perspectives (Hastings et al. 2025). We contribute to this literature by

[1] The Australian National Firearms Agreement's buyback provision was a mandatory government purchase program in 1996-1997 that collected approximately 650,000 prohibited firearms from civilians at market value, funded by a temporary Medicare levy. Participation was compulsory, with criminal penalties for non-compliance after the amnesty period.

evaluating the efficacy of waiting periods on firearm suicides.

We link six decades of county-level mortality data from the National Vital Statistics System with the longitudinal RAND State Firearm Law Database, which codes the exact timing of every waiting period statute enacted between 1813 and 2015 (Cherney et al. 2022; National Center for Health Statistics 2007). To enable full use of the county-level mortality data, we create a crosswalk that accounts for all county mergers and splits occurring after 1959, using information from Bailey and Goodman-Bacon (2015) and Forstall (1994). This produces a balanced panel of counties with harmonized Federal Information Processing Standards (FIPS) codes spanning 1959 to 2019. The staggered adoption of waiting-period laws across states creates a natural experiment that allows us to compare suicide trajectories in treated, not-yet-treated, and never-treated states.

We use the recent developments in the difference-in-differences literature to implement an event study design that exploits both cross-state and within-state overtime variation in exposure to waiting periods. The specification includes county-fixed effects to absorb time-invariant local confounders and year-fixed effects to control for common shocks.

We find a sharp and sustained decline in firearm suicides among men. For the general population, adults aged 55+, and White individuals, waiting periods reduce firearm suicides by approximately 0.5 deaths per 100,000 population—a 7% drop relative to baseline—though estimates are statistically insignificant with confidence intervals largely in negative territory. The effects are larger for men and adults 55+, two groups accounting for the majority of firearm suicides. Crucially, we detect no significant change in non-firearm suicides among these groups, consistent with waiting periods curbing fatalities by delaying access to a uniquely lethal method rather than reducing overall suicide intent.

We also estimate the effect of repealing waiting periods on suicides. We find that states that moved from having waiting periods to repealing them did not experience an increase in firearm suicides afterward. Although most point estimates are negative, they are imprecisely estimated, making it difficult to draw strong conclusions. However, the absence of detectable increases following repeal suggests that waiting periods may create lasting behavioral changes or shift social norms around impulsive firearm purchases in ways that persist even after the formal policy is removed.

We also investigate whether the protective effects of waiting periods operate through a delay mechanism. We estimate models in which the independent variable is the number of days a purchaser must wait. We find evidence of a dose-response relationship: each additional day of mandatory waiting is associated with reductions in firearm suicides, with particularly large effects among older adults and

white individuals. These findings suggest that waiting periods prevent deaths by allowing time for acute suicidal crises to subside, and that states considering waiting period legislation should attend not only to whether a waiting period exists but also to its duration.

2 Data

We use two main data sources. To measure the effect of waiting period on firearm suicides, we use mortality data from the National Vital Statistics System (NVSS) death files for county-level suicides. We also use RAND's state firearm law dataset for waiting periods.

Firearm suicide data

To measure the effect of waiting periods on firearm suicides in the United States, we use mortality data from the National Vital Statistics System (NVSS) covering the years 1959 to 2019 (National Center for Health Statistics 2007). Our outcome of interest is the firearm suicide rate, defined as the number of firearm suicides per 100,000 population in each county and year. We specifically use the Multiple Cause of Death files, which provide detailed causes of death for each death recorded in the US at the county level using ICD-10 codes. The ICD-10 codes allow for the identification of specific causes of death, including suicide. Suicides are further broken down into several categories, allowing for a more detailed analysis of different types of suicide, including those involving firearms.[2] In addition, the data include a range of socioeconomic characteristics of the deceased, such as age, sex, race, marital status, and education level. For each death, we also have information on the county of occurrence, county of residence, and county population size.

Using information on counties that merged and split from Bailey and Goodman-Bacon (2015) and Forstall (1994), we recombine all counties that split or merged after 1959 to produce a crosswalk for the Multiple Cause-of-Death files, creating a balanced panel of all counties from 1959 to 2019. We also harmonize Federal Information Processing Standards (FIPS) codes across counties. This crosswalk represents a methodological contribution, as it had not been previously available, which would allow for

[2] The ICD-10 codes used to defined underlying causes of death due to suicide were X60-X84 (intentional self-harm), and Y87.0 (Sequelae of intentional self-harm). X60 to X69 correspond to intentional self-poisoning, while X70 to X84 correspond to intentional self-harm by other and unspecified means, including drowning, hanging, strangulation, and suffocation, smoke, sharp object, etc. Suicide by firearms was categorized using three specific codes: X72 (intentional self-harm by handgun discharge), X73 (intentional self-harm by rifle, shotgun and larger firearm discharge), and X74 (intentional self-harm by other and unspecified firearm discharge).

the full usage of the county-level death files.

State firearm law

The second dataset we use is the RAND State Firearm Law Database, developed as part of the Gun Policy in America initiative launched in 2016. It is a longitudinal database that tracks all gun laws by state from 1813 to the present (Cherney et al. 2022). The database covers various categories of gun laws, including background check requirements for handguns and long guns, firearm sales restrictions, minimum age requirements, and the presence of waiting periods, defined as the time a seller must wait between the purchase and delivery of a firearm.

We construct the waiting period treatment variable based on the presence of a waiting period in a given state and year using the RAND database. Several states are considered "never-takers," meaning that they never implemented waiting periods for firearm purchases. These states include Colorado, Delaware, Iowa, Massachusetts, Michigan, Missouri, Nebraska, Nevada, New York, North Carolina, Ohio, South Carolina, Utah, and Virginia. Among the states that implemented waiting periods, two scenarios arise: (1) states with a single policy transition, meaning they switched from no waiting period to having the one only once, and (2) states with multiple transitions between implementation and non-implementation. States with a single policy transition include California, the District of Columbia (DC), Florida, Hawaii, Illinois, Maryland, Minnesota, Mississippi, New Jersey, Rhode Island, and Washington. The remaining states did not experience policy transitions. We present different cohorts of states in Table 3 and a map of all states with different treatment sequences in Figure 1.

Sample construction

The sample is composed of states that experience either one policy transition from no-treatment to treatment, or none at all. This yields a final sample of 23 states: 10 treatment states that adopted a waiting period at some point before or during the sample period, and 13 control states that never adopted a waiting period for firearm purchases. In terms of counties, the final sample is composed of 1,306 counties. A total of 932 counties as part of the 'never-takers', six counties adopted a waiting period prior to 1959, and 374 counties adopted a waiting period between 1959 and 2019. We present the states and counties included in our final sample, along with their adoption status of the waiting period policies, in Figures 2 and 3. We show the staggered adoption of waiting periods in Figure 2, while we present the corresponding county

count in Figure 3. For the purpose of our analysis, the comparison group is composed of never-takers and yet-to-be-treated. For population and demographic information, we combined our final dataset with data from the Decennial US Census (U.S. Census Bureau 2020).

We also construct a second sample to estimate the effects of repealing waiting periods. This sample includes states that either transitioned from treatment to no-treatment or were always treated. This yields a final sample of 14 states: 8 treatment states that repealed their waiting period law during the sample period, and 6 control states that maintained a waiting period throughout. We present the staggered repeal of waiting periods and the county counts in Figures 4 and **??** respectively.

Outcome variable

The main outcome variable is the suicide rate by firearm per 100,000 people. We identify suicides related to firearms using ICD-10 codes for the underlying causes of death.[3]

We then calculated county-level suicide rates by aggregating individual-level mortality data for each year and county. Specifically, we sum the number of firearm suicides in each county-year and divide by the corresponding county population, multiplying by 100,000 to calculate the suicide rate by firearm in a given county and year. You can find trends in firearm and non-firearm suicide rates from 1959 to 2019 in Figures 6 and 7. There is a substantial heterogeneity in firearm suicide rates by treatment status (Figure 8), with consistently lower rates observed in states that implemented waiting period policies, as well as marked demographic disparities (Figure 9), with particularly elevated rates among men and people over 55+ years of age.

Summary statistics

We present the summary statistics in Table 1. Counties that adopted a mandatory waiting period tend to have lower total suicide rates per 100,000 people, overall firearm suicide rates, men, older adults (age 55 and older) and white individual firearm suicide than counties that never adopted a mandatory waiting period. In addition, counties that adopted a mandatory waiting period tend to have a similar share of the female population, college educated, and tend to have a lower proportion of their population living below the poverty line. To assess the comparability of treated and control counties, we present covariate

[3] We use the following ICD-10 codes from the NVSS for firearm suicides: X72 (intentional self-harm by handgun discharge), X73 (intentional self-harm by rifle, shotgun, and larger firearm discharge), and X74 (intentional self-harm by other and unspecified firearm discharge).

balance statistics in Table 2. The table shows that while treated and control counties are similar on several dimensions, there are statistically significant differences in the share of college-educated residents and the proportion living below the poverty line, with control counties having higher rates of both. Notably, treated counties also exhibit significantly lower firearm suicide rates among adults aged 55 and older prior to treatment.

3 Empirical strategy

We exploit two sources of variation in exposure to the waiting period to assess its effect on firearm suicide rates. The first source of variation comes from the state-level adoption of firearm purchase waiting periods, which allows us to compare between counties that have adopted waiting periods and those that have not. The second source of variation is based on the timing of firearm suicides, with some suicides that occurred before and after the adoption of waiting periods within counties in the adopting states. The first source of variation compares outcomes between counties in states that have adopted waiting periods and those that have not. The second source of variation is based on the timing of suicides, comparing suicides that occurred in counties before and after the adoption of waiting periods within adopting states. The identifying assumption is that absent the adoption of waiting periods, counties that adopted waiting periods and states that never adopted them would have followed similar trends in suicide rates by firearms (parallel trends assumption), that no individual does not anticipate treatment (no anticipation assumption), and Stable Unit Treatment Value Assumption (SUTVA).

We estimate an event study model that allows for unrestricted treatment effect heterogeneity across counties using Callaway and Sant'Anna (2021). Let y_{ist} denote the firearm suicide rate per 100,000 people in county i in state s at time t:

$$y_{ist} = \sum_{l=-K}^{L} \beta_l \mathbf{1}\{t - E_s = l\} + \theta_i + \lambda_t + \varepsilon_{ist} \tag{1}$$

where E_s is the year in which state s adopted a waiting period, and $\mathbf{1}\{t - E_s = l\}$ is an indicator variable equal to one when time t is l periods away from treatment adoption. For example, Florida adopted a waiting period on January 10, 1991. In 1988, three years before adoption, $t - E_s = 1988 - 1991 = -3$, so the indicator $\mathbf{1}\{t - E_s = -3\}$ equals one. The coefficients β_l capture treatment effects from K periods before to L periods after adoption. All regressions include county fixed effects (θ_i) and year fixed effects (λ_t). We use counties that are not-yet-treated and never-treated as the control group. Standard errors are

cluster-bootstrapped at the state level (the treatment variation level).

The coefficients of interest are β_l. For $l < 0$, these coefficients provide a test of the parallel trends assumption: if β_l are statistically insignificant, this supports the assumption that treated and control counties would have followed similar trends in the absence of treatment. The coefficient β_{-1} also provides a test of the assumption of no anticipation; statistical insignificance suggests that people in an adopting county did not alter their behavior in anticipation of the policy. This assumption is plausible given that suicides are often impulsive and acute, making it unlikely that individuals would systematically time suicide attempts based on anticipated future gun policy changes. For $l \geq 0$, the coefficients β_l capture the dynamic post-treatment effects of waiting periods on firearm suicide rates.

To evaluate whether waiting period laws simply shift methods of suicide rather than reducing overall suicide rates, we analyze their impact on nonfirearm suicide rates. If waiting periods genuinely decrease total suicides, we would expect no significant change in non-firearm suicide methods. Using the same analytical approach (equation 1) with non-firearm suicide rates as our outcome variable, we test this relationship. The absence of any significant effect on nonfirearm suicides would strengthen the validity of our core assumption that waiting periods specifically prevent firearm-related suicides rather than causing method substitution.

For the analysis examining states that move out of treatment (i.e., states that repeal their waiting period laws), the treatment group consists of states that transition from having a waiting period to not having one, while the control group consists of states that consistently maintain waiting periods throughout the study period. This reverse treatment design allows us to examine whether the removal of waiting periods leads to increases in firearm suicide rates, providing additional evidence for the causal effect of these policies.

4 Results

The Effects of Waiting Period Adoption on Firearm Suicide Rates

Overall Effect on Firearm Suicides. We present the results of estimating equation 1 in Figure 10. We show the estimates 10 years before the adoption of waiting periods and 10 years after. For the 10 years before adoption, we present point estimates and their associated 95% confidence intervals, which correspond to pre-periods. We can use these pre-treatment estimates to assess the parallel trends assumption. For 10 years after adoption, we present the point estimates and their associated 95% confidence intervals for the post-adoption periods. These estimates correspond to the treatment effects.

We find a statistically insignificant decrease in the suicide rate by firearm in the periods following the adoption of waiting periods. Specifically, adopting waiting periods reduces the suicide rate by firearms by about 0.47 deaths per 100,000 people, though the results are statistically insignificant. Although the negative point estimates are statistically insignificant, the majority of the confidence intervals lie on the negative side. This pattern suggests a modest protective effect of waiting periods on overall firearm suicide rates, though the evidence remains inconclusive due to imprecise estimation. The lack of statistical significance may reflect heterogeneity in treatment effects across different demographic groups or insufficient statistical power when pooling across all populations.

Effect on men. We present the results of estimating equation 1 in Figure 11 when restricting the sample to men only. In the post-treatment period, we observe an average reduction in firearm suicides among men by 1.3 deaths per 100,000 after the adoption of waiting periods. We find this effect statistically significant at the 90% confidence level. This finding is particularly important given that men account for the vast majority of firearm suicides in the United States. The statistically significant reduction among men, in contrast to the null overall effect, suggests that waiting periods may be most effective for populations with higher baseline rates of firearm suicide. The mechanism likely operates through disrupting impulsive suicide attempts, which research suggests are more common among men who use firearms.

Effect on adults aged 55 and older. We present the results of estimating equation 1 on suicides among individuals 55 years and older in Figure 12. We find that the pre-treatment estimates are not statistically significant from zero, supporting the assumption of parallel trends. After the adoption of waiting periods, we observe a statistically insignificant reduction in firearm suicides for adults 55 years and older by 26.3 deaths per 100,000 (p-value = 0.18). Despite the lack of statistical significance, the magnitude of this point estimate is substantial and represents one of the largest reductions we observe across demographic subgroups. This age group merits particular attention in suicide prevention efforts, as older adults who attempt suicide are more likely to die from their attempts, making the potential protective effect of waiting periods especially valuable for this population.

Effect on white individuals. We show the causal effect of waiting period laws on firearm suicides among white individuals in Figure 13. The pretreatment estimates support the assumption of parallel trends. We find a statistically significant reduction in firearm suicides among white individuals by 21.6 deaths per 100,000 (p-value = 0.09), which represents a substantial decrease in the suicide rate by firearm among this population. This large and statistically significant effect is noteworthy because white

individuals have historically had the highest firearm suicide rates in the United States. The magnitude of the reduction suggests that waiting periods may be particularly effective in communities with higher baseline firearm ownership rates and greater cultural acceptance of firearm use. This finding provides strong evidence that mandatory waiting periods can significantly reduce firearm suicides in high-risk populations.

Effect on other causes of suicide. We show the causal effect of waiting period laws on other causes of suicide (i.e., all causes of suicide excluding suicides by firearm) in Figures 14–17. Once again, the pretreatment estimates support the assumption of parallel trends. We find that adopting waiting periods is associated with an increase of 0.40 deaths per 100,000 (p-value = 0.135). However, this effect is not statistically significant from zero, confirming that people did not substitute towards suicide using other methods after a state adopts waiting periods. We observe the same pattern when estimating the effect of waiting periods on other causes of suicide among men and adults 55 years of age and find a significant decrease among white individuals. The absence of substitution effects is crucial for interpreting the policy impact of waiting periods. If individuals simply switched to other suicide methods, the net public health benefit would be diminished. Our results suggest that waiting periods reduce overall suicide mortality rather than merely redirecting individuals toward alternative methods, supporting the hypothesis that many firearm suicide attempts are impulsive and that introducing a delay can prevent deaths.

The Effects of Waiting Period Repeal on Firearm Suicide Rates

Overall effect of repeal on firearm suicides. We show the results of the estimation in Figure 18. We present the estimates 5 years before the repeal of waiting periods and 5 years after. For the 5 years before repeal, we show point estimates and their associated 95% confidence intervals, which correspond to pre-periods. We can use these pre-treatment estimates to assess the parallel trends assumption. For 5 years after repeal, we present the point estimates and their associated 95% confidence intervals for the post-adoption periods. These estimates correspond to the treatment effects of repealing waiting periods on firearm suicides. We find a statistically insignificant decrease in the suicide rate by firearm in the periods following the repeal of waiting periods. Although most of the point estimates are negative, they are all imprecisely estimated. The imprecision of these estimates makes it difficult to draw strong conclusions about the effects of repealing waiting periods. Interestingly, the negative point estimates suggest that waiting periods may have persistent protective effects that endure even after repeal. This could occur through social norms around impulsive firearm purchases may have shifted in lasting ways.

Effect on men. We present the results of the event study estimates when restricting the sample to men only in Figure 19. In the post-repeal period, we observe a statistically insignificant reduction in firearm suicides among men. We find that the point estimates are imprecisely estimated. The continued absence of increases in firearm suicides among men following repeal is noteworthy given that our adoption analysis showed statistically significant protective effects of waiting periods for this group. The pattern is consistent with the hypothesis that waiting periods create lasting behavioral changes or institutional arrangements that continue to protect against impulsive firearm suicides even after the formal policy is removed.

Effect on adults aged 55 and older. We show the results of the estimation for adults 55 years and older in Figure 20. We find that estimates before repeal are not statistically significant from zero, supporting the assumption of parallel trends. After the repeal of waiting periods, we observe statistically insignificant point estimates. Given the substantial (though imprecise) protective effects we observed for this age group under adoption, the absence of detectable increases following repeal is consistent with persistent effects of prior waiting period exposure.

Effect on white individuals. We show the results of the estimation for white individuals in Figure 21. We find that estimates before repeal are statistically insignificant, supporting the assumption of parallel trends. Although the point estimate at $t = -1$ is statistically significant, it is unlikely that this represents a violation of the no anticipation assumption since the estimate is positive rather than negative, which would be expected if individuals were anticipating the repeal by increasing firearm suicides before the policy change. After the repeal of waiting periods, we observe statistically insignificant point estimates.

Effect on other causes of suicide. We show the causal effect of removing waiting periods on other causes of suicide (i.e., all causes of suicide that exclude suicides by firearm) in Figures 22–25. Once again, the pretreatment estimates support the assumption of parallel trends. We find that the repeal of waiting periods is associated with a decrease of 0.33 deaths per 100,000 (p-value = 0.142). We observe the same pattern when estimating the effect of the repeal of waiting periods on other causes of suicide among men and adults ages 55+, and white individuals. The consistent lack of substitution effects in the repeal analysis mirrors our findings from the adoption analysis, providing further evidence that waiting periods do not simply shift individuals toward alternative suicide methods. The absence of increases in non-firearm suicides following repeal, combined with the absence of increases in firearm suicides, suggests that the protective effects established during the waiting period era may persist even after policy removal.

Placebo and Other Estimators

One potential concern is that the timing of waiting period adoption could coincide with other unobserved factors that differentially affect suicide rates. To address this issue, we examine dynamic treatment effects for firearm suicides among women in Figure 26. As we show in Figure 9, women are substantially less likely than men to use firearms when committing suicide. This makes women a good placebo group: if broader social or policy changes occurring at the same time as waiting period adoption were driving the results, we would expect to observe similar declines in women's suicides. Instead, we find that the post-treatment estimates are consistently non-negative, with an average effect of about 0.33 suicides per 100,000 women ($p = 0.031$). The fact that women's firearm suicide rates do not fall after the adoption of waiting periods provides strong evidence that the reductions we observe in the overall population are not driven by spurious correlations or contemporaneous policy shifts, but rather reflect the causal impact of waiting period laws. This placebo test strengthens our confidence in attributing the observed reductions among men and other high-risk groups specifically to the waiting period policy rather than to confounding trends. The null (or slightly positive) effect for women also suggests that waiting periods may operate primarily by disrupting impulsive suicide attempts in populations with high baseline firearm access and use, which characterizes men far more than women.

Furthermore, recent methodological advances in difference-in-differences estimation have underscored potential biases in traditional two-way fixed effects (TWFE) models when treatment timing is staggered across units. To evaluate the robustness of our findings, we compare treatment effect estimates from multiple estimators specifically designed for staggered adoption settings in Figure 27. Alongside TWFE, we report estimates from Callaway and Sant'Anna (2021), Sun and Abraham (2021), Gardner (2022), Roth and Sant'Anna (2023), and Borusyak, Jaravel, and Spiess (2024). Across these approaches, we find that the results consistently indicate that waiting periods reduce firearm suicide rates. The point estimates are similar in magnitude, and the confidence intervals substantially overlap across methods. This convergence of evidence from methodologically distinct estimators provides strong support for the robustness of the estimated treatment effect, indicating that our findings are not driven by the choice of identification strategy. The consistency across estimators is particularly important given recent methodological debates about the validity of TWFE models in staggered adoption settings. Our results demonstrate that the protective effects of waiting periods are not artifacts of potentially biased TWFE estimation but rather represent genuine causal effects that persist across alternative, more robust estimation approaches. This robustness check substantially strengthens the credibility of our core findings and their

policy implications.

Mechanism: The Role of Waiting Period Length

A key question for policy design is whether the protective effects of waiting periods operate through a delay mechanism—that is, whether longer waiting periods provide greater protection against impulsive firearm suicides. To investigate this mechanism, we estimate a two-way fixed effects model in which the treatment variable is the number of days a purchaser must wait between buying and receiving a firearm, rather than a binary indicator for any waiting period.

Table 4 presents the results. We find that each additional day of mandatory waiting is associated with a reduction in firearm suicides across all demographic groups examined. For the overall population, an additional waiting day reduces firearm suicides by 0.063 deaths per 100,000, representing a 0.74 percent decline relative to the baseline mean of 8.45 deaths per 100,000. The effect is statistically significant at the 10 percent level.

The protective effects of longer waiting periods are substantially larger for older adults and white individuals—two groups with elevated baseline firearm suicide rates. Among adults aged 55 and older, each additional waiting day is associated with a reduction of 5.63 deaths per 100,000, a 22.5 percent decline relative to the baseline mean of 24.99. For white individuals, the corresponding reduction is 4.12 deaths per 100,000, or 25.6 percent of the baseline rate. Both estimates are statistically significant at the 5 percent level. Among men, we observe a reduction of 0.096 deaths per 100,000 per waiting day, though this estimate is imprecisely estimated.

These findings provide direct evidence that the delay mechanism is central to the protective effects of waiting period laws. The dose-response relationship—whereby longer waiting periods yield larger reductions in firearm suicides—is consistent with the hypothesis that waiting periods prevent deaths by allowing time for acute suicidal crises to subside. The particularly large effects among older adults and white individuals suggest that these populations may be especially responsive to delays in firearm access, potentially because their suicide attempts are more likely to be impulsive or because they have higher baseline access to firearms.

From a policy perspective, these results suggest that states considering waiting period legislation should attend not only to whether a waiting period exists but also to its duration. A three-day waiting period, for example, may yield meaningfully different public health outcomes than a seven-day or fourteen-day requirement. The substantial per-day effects we observe indicate that even modest extensions to

existing waiting periods could generate additional reductions in firearm suicides.

5 Conclusion

This study provides the most extensive evidence to date that waiting periods for firearm purchases are an effective population-level suicide prevention tool. Leveraging six decades of county-by-year mortality data and the full historical record of state gun laws, we find that waiting-period laws reduce firearm-suicide rates by roughly 0.5 deaths per 100,000 population, or about 7 percent relative to baseline. The effects are strongest for men and adults aged 55 and older, the two groups that account for the majority of firearm suicides. Crucially, we detect no significant change in non-firearm suicides, consistent with the interpretation that waiting periods save lives by limiting access to a uniquely lethal method rather than causing substitution to other means.

Our analysis of waiting period repeals provides additional insight into the dynamics of these policies. States that repealed their waiting periods did not experience statistically significant increases in firearm suicides, though the point estimates are imprecisely estimated. While this might initially seem inconsistent with the protective effects observed under adoption, the absence of detectable increases following repeal suggests that waiting periods may generate lasting behavioral changes or shift social norms around impulsive firearm purchases in ways that persist even after the formal policy is removed. This pattern of persistent effects was consistent across demographic subgroups, including those that showed the strongest protective effects under adoption.

Our placebo analysis further support this causal interpretation. Women, who are substantially less likely to use firearms in suicide attempts, serve as a natural placebo group. Their firearm suicide rates did not decline after the adoption of waiting periods, strengthening the case that the reductions observed among men and other high-risk groups reflect the true effect of waiting period laws rather than contemporaneous unobserved changes or spurious correlations. Moreover, robustness checks using recently developed difference-in-differences estimators for staggered adoption consistently indicate negative treatment effects, reinforcing the credibility of our findings.

We also provide direct evidence on the mechanism through which waiting periods reduce firearm suicides. By estimating models in which the treatment variable is the number of mandatory waiting days rather than a binary indicator, we find a clear dose-response relationship: each additional day of mandatory waiting is associated with further reductions in firearm suicides. The effects are particularly pronounced

among older adults and white individuals, with each additional waiting day reducing firearm suicides by over 20 percent relative to baseline for these groups. These findings support the hypothesis that waiting periods save lives by allowing time for acute suicidal crises to subside, and suggest that policymakers should consider not only whether to implement a waiting period but also its optimal duration.

From a policy perspective, these results highlight the preventive potential of waiting periods as a low-cost intervention. Unlike broader restrictions on firearm ownership, waiting periods impose only a temporary delay on purchases while providing a crucial buffer against impulsive, high-lethality acts of self-harm. The findings suggest that relatively modest regulatory measures can yield substantial public health benefits. Policymakers debating gun violence interventions often face a trade-off between political feasibility and measurable impact. Waiting periods appear to offer an unusually favorable balance: they impose minimal costs on lawful purchasers while generating sizable reductions in mortality.

Future research should extend this work by examining heterogeneity in effects across urban and rural contexts, racial and ethnic groups, and by considering potential interactions with complementary interventions such as extreme risk protection orders and safe storage laws. In addition, comparative analyses with other "cooling-off" regulations—such as waiting periods for prescription opioid refills or other lethal means—may provide further insight into the broader applicability of time-based barriers in suicide prevention.

Overall, this study demonstrates that waiting periods are a powerful tool for reducing firearm suicides, particularly among the groups most at risk. By creating a critical pause between purchase and possession, waiting-period laws save lives in contexts where minutes and hours can make the difference between a temporary crisis and a permanent tragedy.

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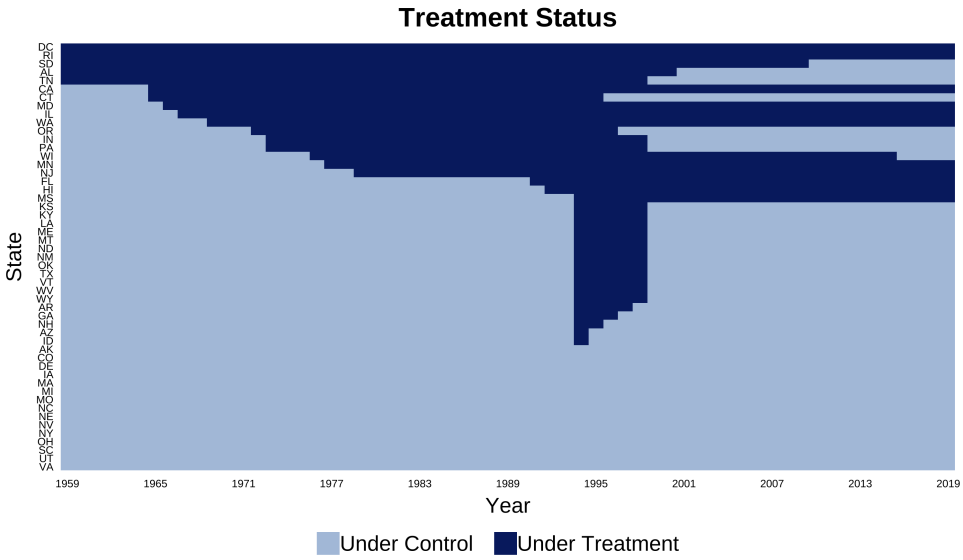
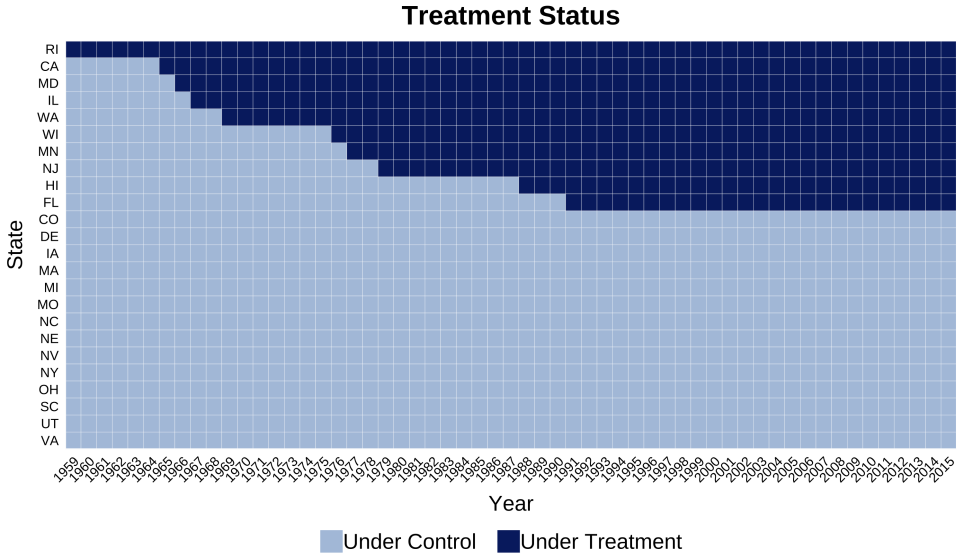
Fig. 1. Timing of Waiting Period Policy Adoption Across All States



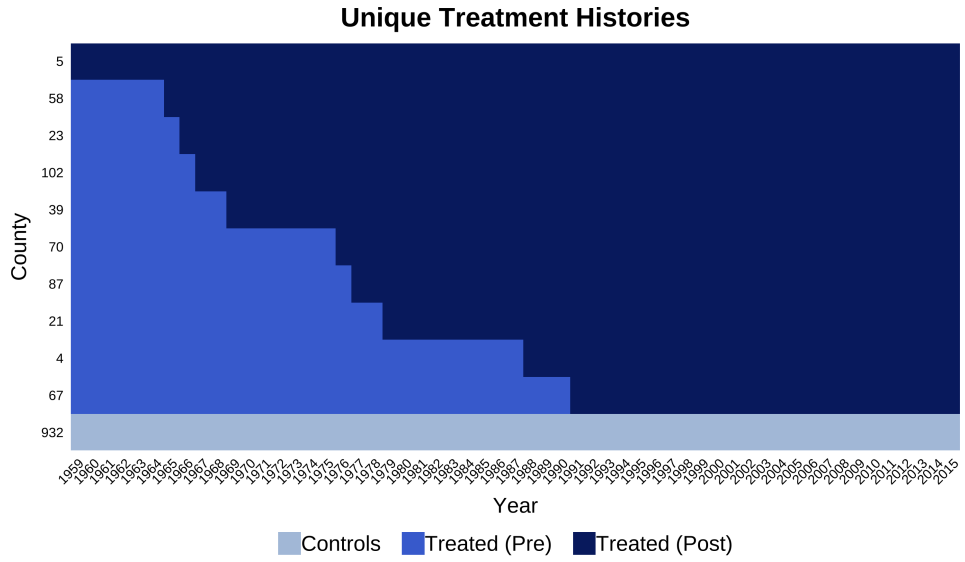
Note: This staggered adoption panel view illustrates the year of waiting period policy implementation or exit for each state. It provides visual clarity on treatment timing across states, which is crucial for interpreting the event study estimates and understanding the source of identifying variation.
Source: RAND State Firearm Law Database, 1813–2015.

Fig. 2. Timing of Waiting Period Policy Adoption Across States



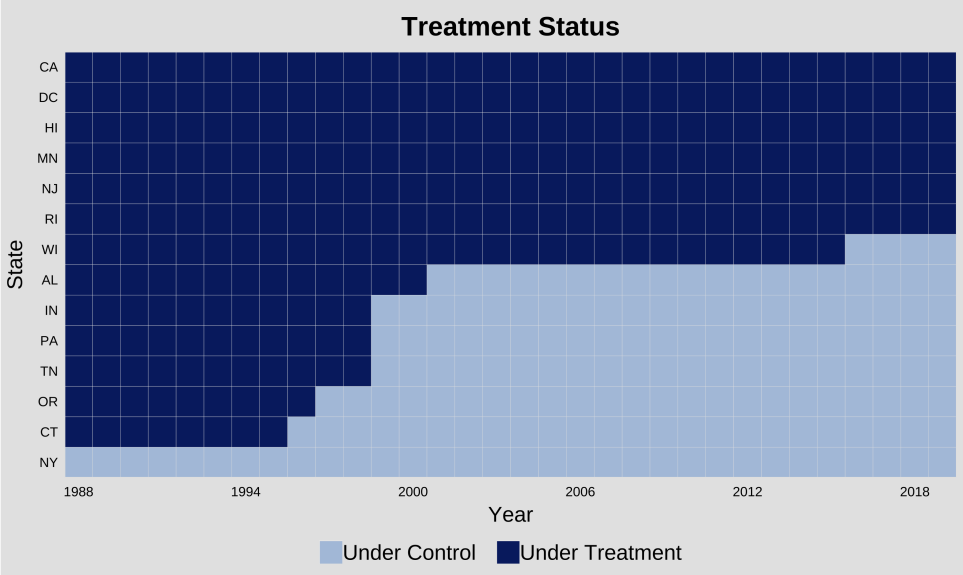
Note: This staggered adoption panel view illustrates the year of waiting period policy implementation for each state included in the study. It provides visual clarity on treatment timing across states, which is crucial for interpreting the event study estimates and understanding the source of identifying variation.
Source: RAND State Firearm Law Database, 1813–2015.

Fig. 3. Timing of Waiting Period Policy Adoption Across States: Number of Counties



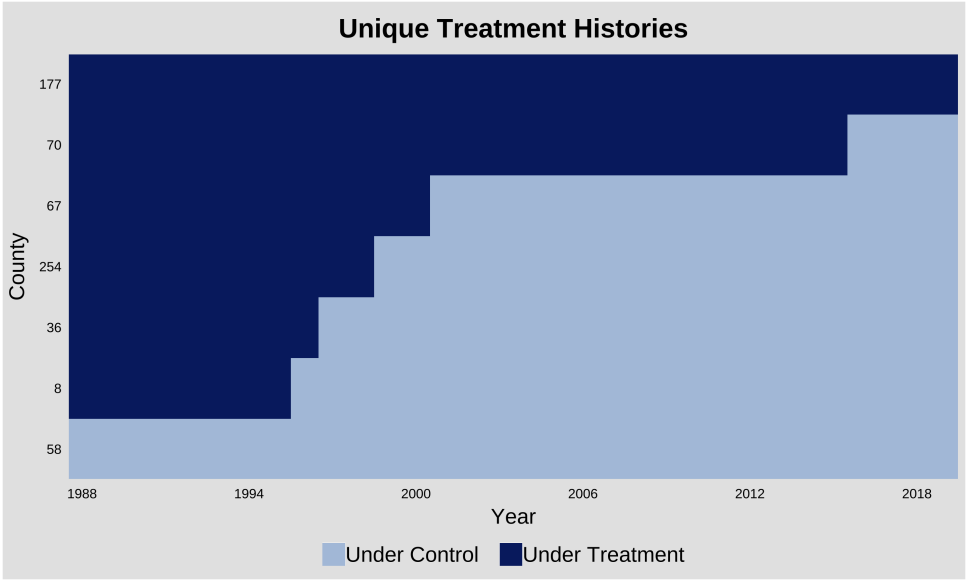
Note: This alternative view of policy adoption timing complements Figure 2. It emphasizes the distribution of treated versus control counties over time, helping to validate the use of staggered treatment timing in the empirical strategy.
Source: RAND State Firearm Law Database, 1813–2015.

Fig. 4. Timing of Waiting Period Policy Adoption Across States That Moved Out of Treatment



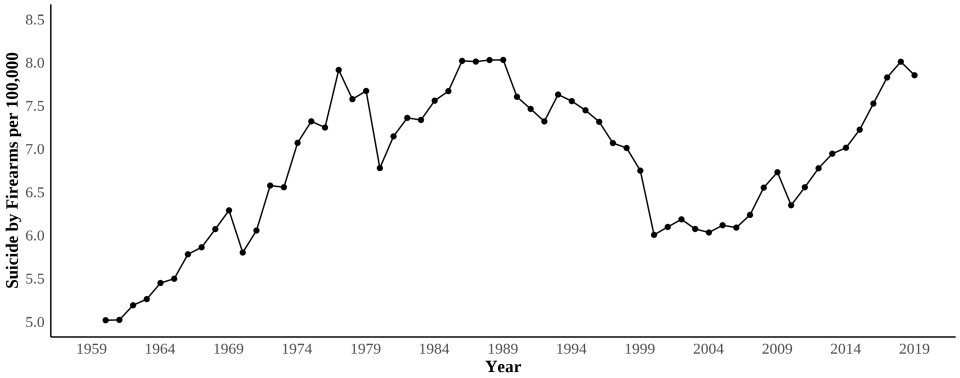
Note: This staggered adoption panel view illustrates the year of waiting period policy implementation for each state included in the study. It provides visual clarity on treatment timing across states, which is crucial for interpreting the event study estimates and understanding the source of identifying variation.
Source: RAND State Firearm Law Database, 1813–2015.

Fig. 5. Timing of Waiting Period Policy Adoption Across States: Number of Counties



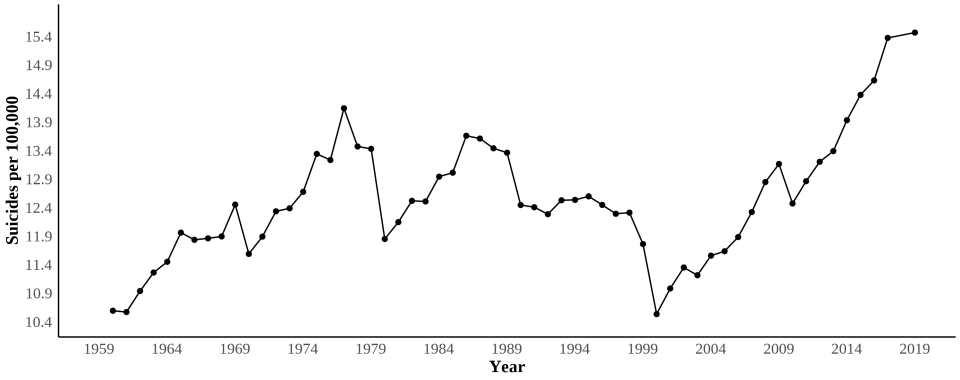
Note: This alternative view of policy adoption timing complements Figure 4. It emphasizes the distribution of treated versus control counties over time, helping to validate the use of staggered treatment timing in the empirical strategy.
Source: RAND State Firearm Law Database, 1813–2015.

Fig. 6. Trends in Firearm Suicide Rates Across Counties, 1959–2019



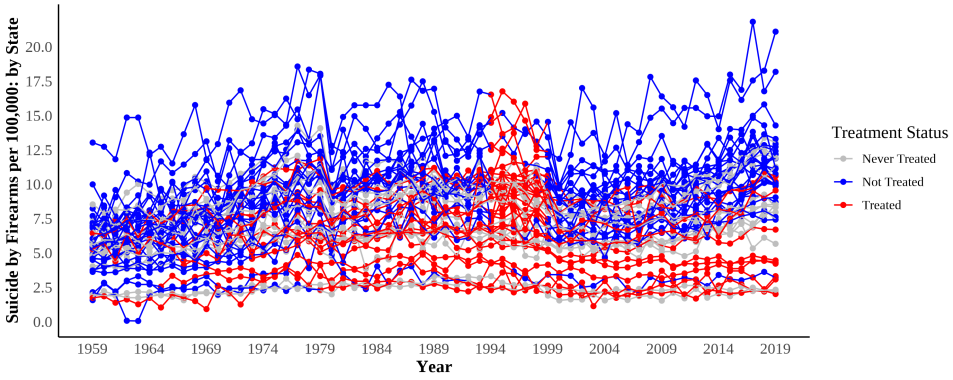
Note: This figure shows the annual trend in suicide rates by firearm (per 100,000 population) across US counties from 1959 to 2019. The figure highlights the long-term trajectory of firearm suicides, providing historical context for analyzing the impact of waiting period laws introduced in different years and states.
Source: National Vital Statistics System (NVSS), Multiple Cause of Death Files, 1959–2019.

Fig. 7. Trends in Overall Suicide Rates Across Counties, 1959–2019



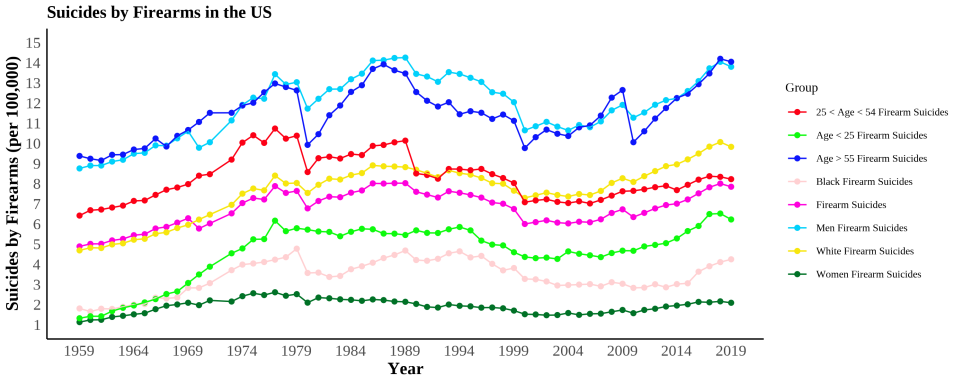
Note: This figure displays the overall suicide rate (all causes, per 100,000 population) from 1959 to 2019. It offers a comparison benchmark for firearm-specific suicides and helps evaluate whether general suicide trends might confound the estimated effects of waiting period policies.
Source: National Vital Statistics System (NVSS), Multiple Cause of Death Files, 1959–2019.

Fig. 8. Firearm Suicide Rates by State Treatment Status, 1959–2019



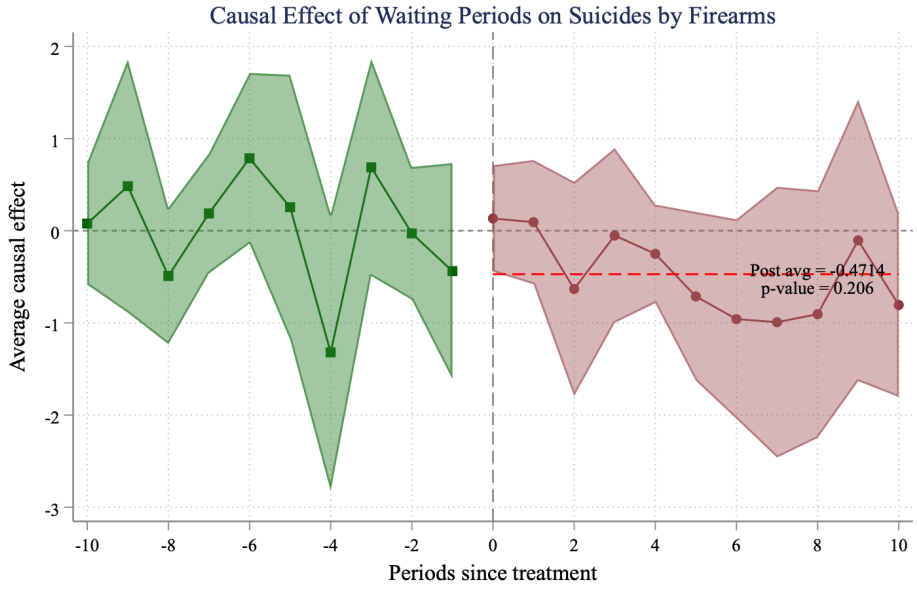
Note: This figure displays firearm suicide rates (per 100,000 population) from 1959 to 2019 across states, categorized by treatment status. "Treated" states implemented waiting period policies, "Not Treated" states never adopted such policies during the study period, and "Never Treated" states represent the control group. The consistently lower rates in treated states suggest a potential protective effect of waiting period legislation on firearm suicide mortality.
Source: National Vital Statistics System (NVSS), Multiple Cause of Death Files, 1959–2019.

Fig. 9. Firearm Suicide Rates by Demographic Group in the US, 1959–2019



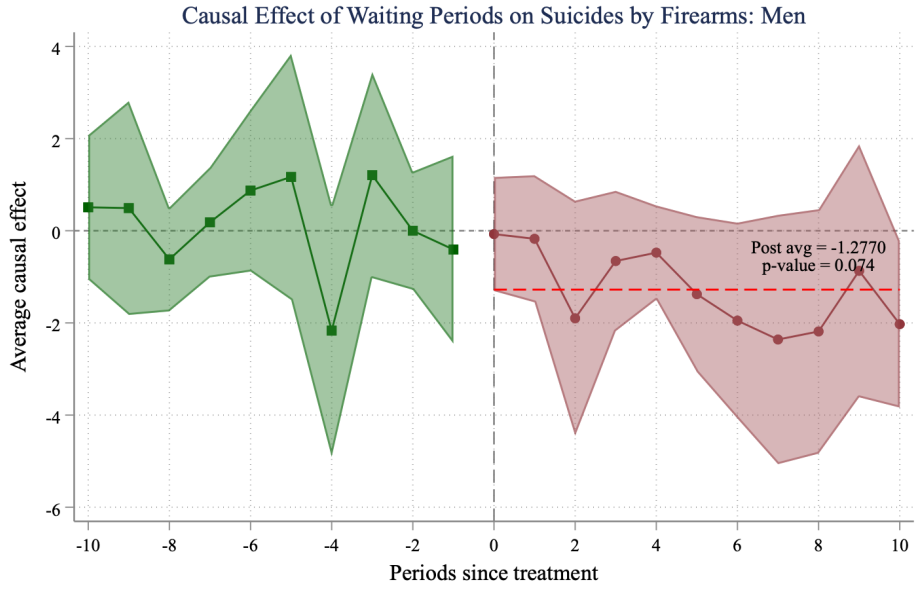
Note: This figure illustrates firearm suicide rates (per 100,000 population) across demographic categories from 1959 to 2019. The substantial differences between men and women, age groups, and racial categories highlight the importance of demographic-specific approaches to suicide prevention. The recent increases across multiple groups after 2010 suggest concerning trends that may warrant targeted intervention strategies.
Source: National Vital Statistics System (NVSS), Multiple Cause of Death Files, 1959–2019.

Fig. 10. Estimated Effect of Waiting Periods on Overall Firearm Suicide Rates



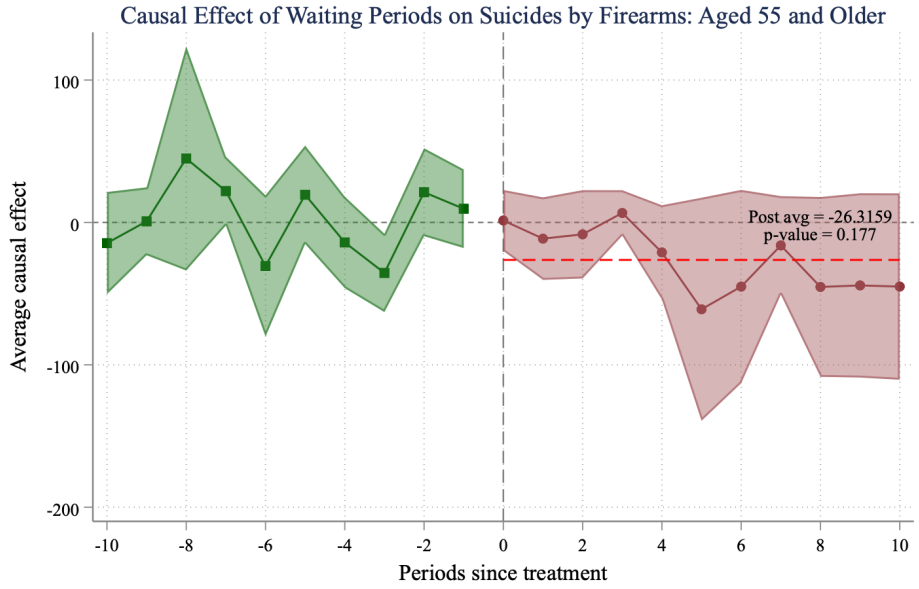
Note: This figure shows the dynamic effects of waiting period laws on firearm suicide rates across US counties. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. Standard errors are bootstrapped and clustered at the state level.

Fig. 11. Effect of Waiting Periods on Firearm Suicide Rates Among Men



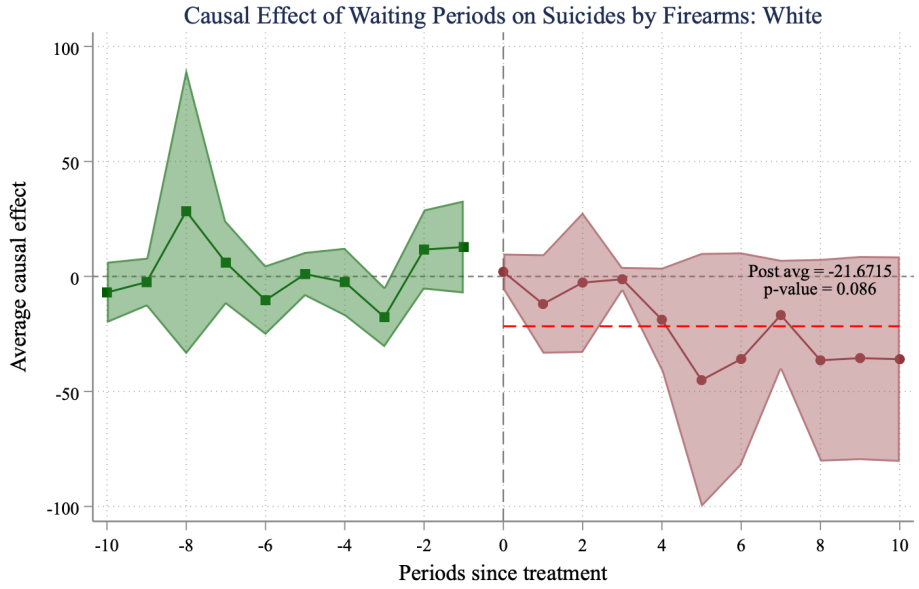
Note: This figure focuses on the male population, showing how waiting period laws affect firearm suicide rates for men specifically. Standard errors are bootstrapped and clustered at the state level.

Fig. 12. Effect of Waiting Periods on Firearm Suicide Rates Among Adults Aged 55+



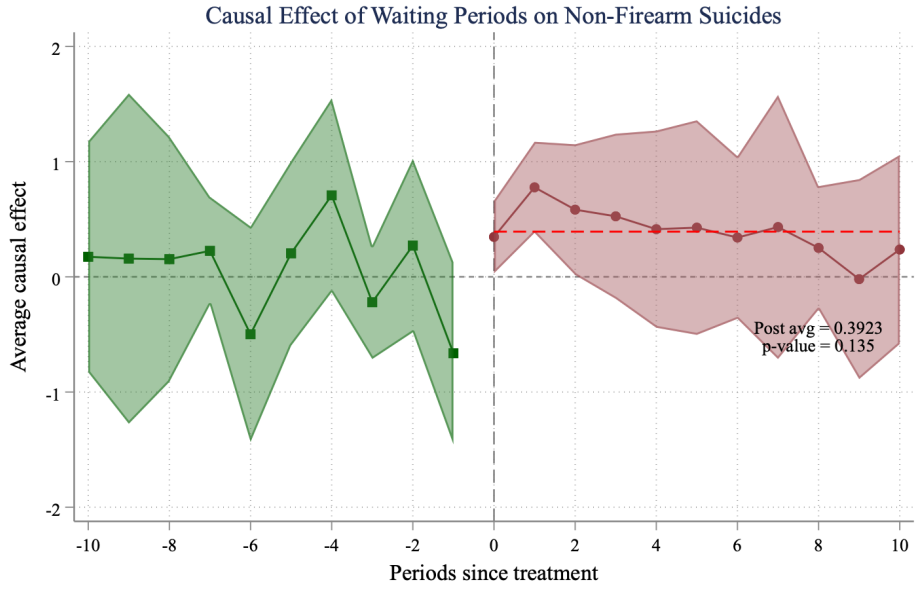
Note: This event study estimates the policy effect on older adults, a group at elevated suicide risk. Standard errors are bootstrapped and clustered at the state level.

Fig. 13. Effect of Waiting Periods on Firearm Suicide Rates Among White Individuals



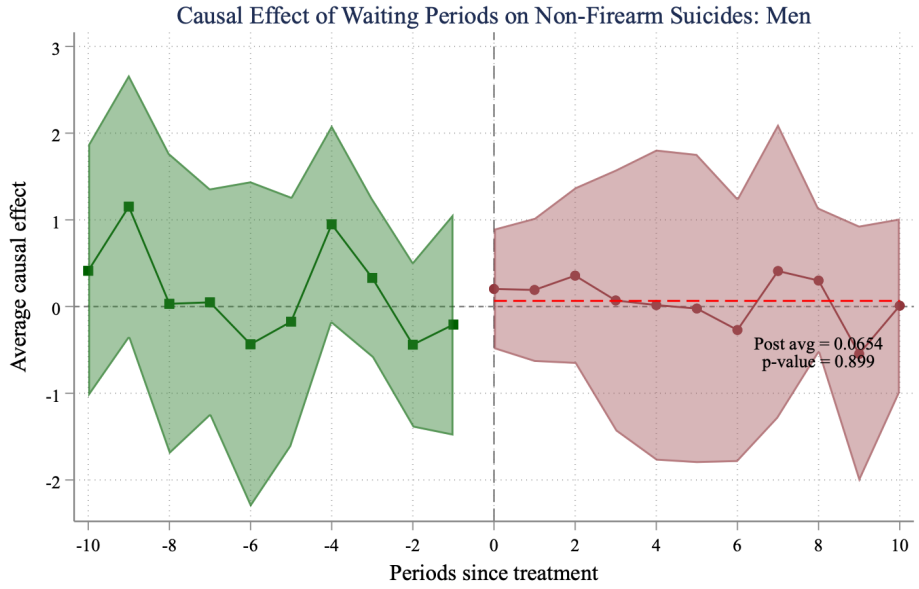
Note: This figure examines firearm suicide trends for white individuals. Standard errors are bootstrapped and clustered at the state level.

Fig. 14. Effect of Waiting Periods on Non-Firearm Suicide Rates



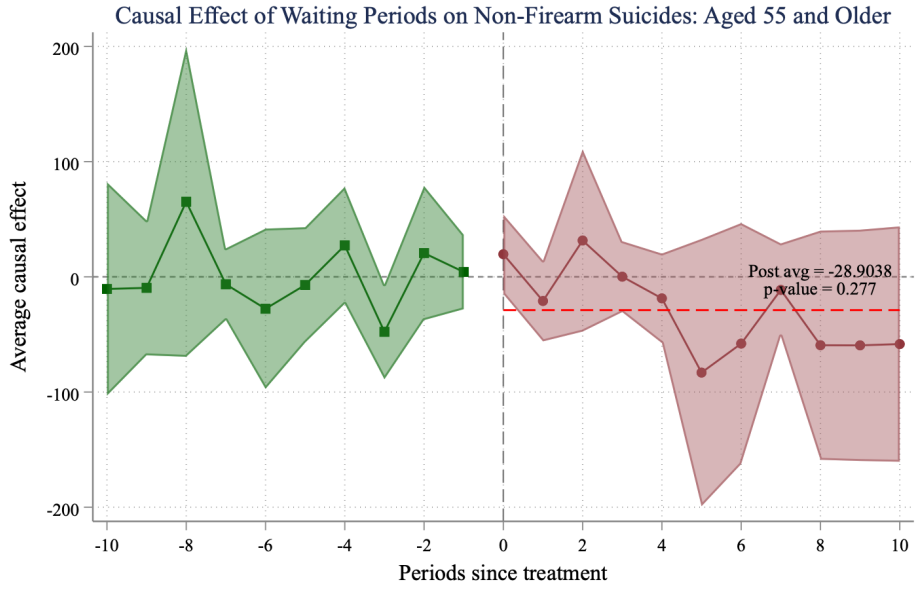
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. The absence of significant differences in pre-treatment periods supports the parallel trends assumption. Standard errors are bootstrapped and clustered at the state level.

Fig. 15. Effect of Waiting Periods on Non-Firearm Suicide Rates Among Men



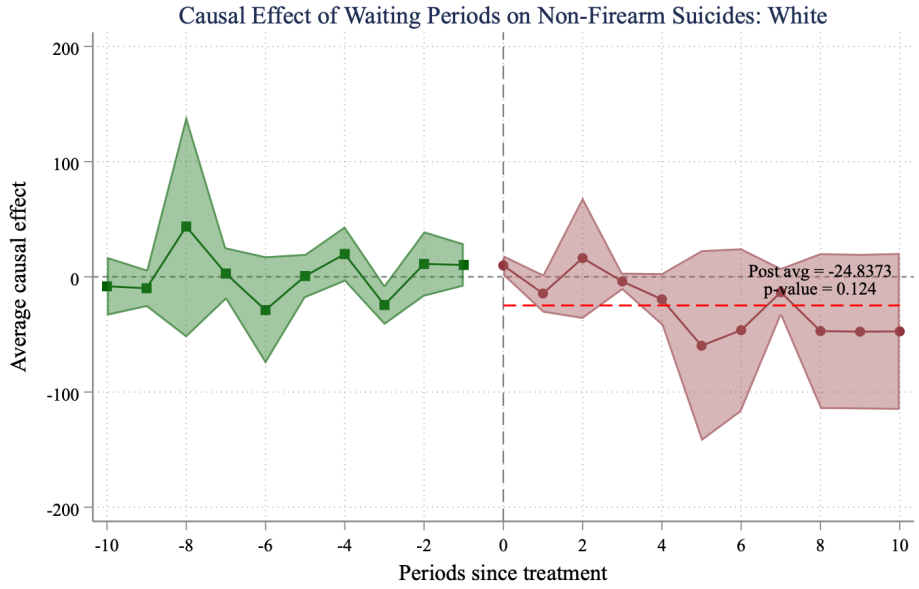
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among men. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. The absence of significant differences in pre-treatment periods supports the parallel trends assumption. Standard errors are bootstrapped and clustered at the state level.

Fig. 16. Effect of Waiting Periods on Non-Firearm Suicide Rates Among Adults Aged 55+



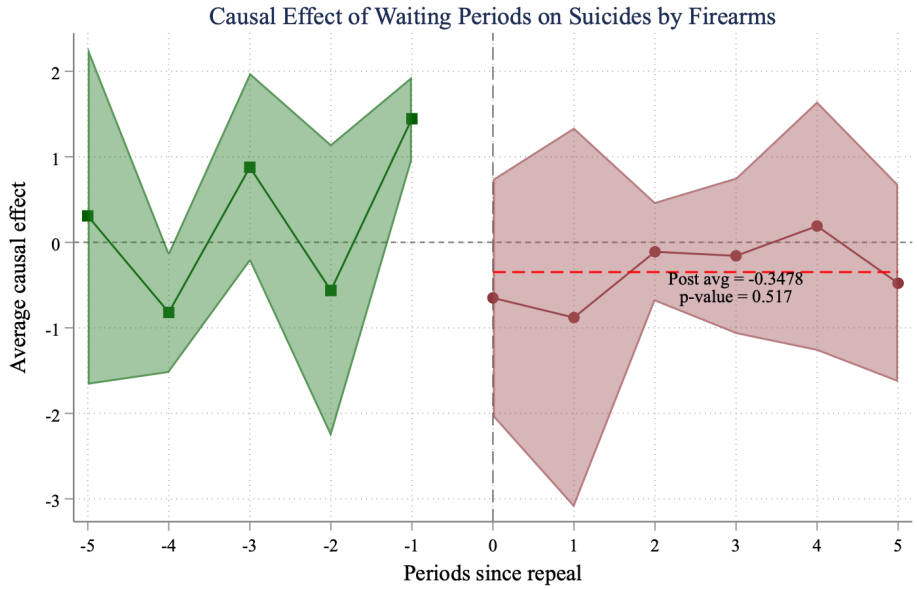
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among adults aged 55+. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. The absence of significant differences in pre-treatment periods supports the parallel trends assumption. Standard errors are bootstrapped and clustered at the state level.

Fig. 17. Effect of Waiting Periods on Non-Firearm Suicide Rates Among White Individuals



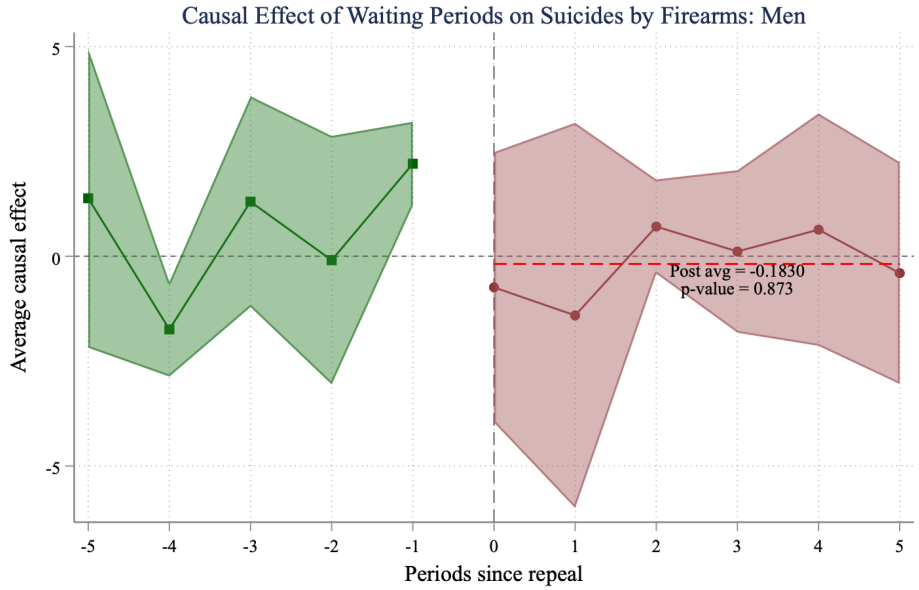
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among White individuals. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. The absence of significant differences in pre-treatment periods supports the parallel trends assumption. Standard errors are bootstrapped and clustered at the state level.

Fig. 18. Estimated Effect of Waiting Periods on Overall Firearm Suicide Rates: Out of Treatment Sample



Note: This figure shows the dynamic effects of waiting period laws on firearm suicide rates across US counties using a sample of states that move out of treatment. The sample is composed of states that experience either one policy transition from treatment to no-treatment, or always treated, yielding 23 states (13 treatment, 13 control). Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. Standard errors are bootstrapped and clustered at the state level.

Fig. 19. Effect of Waiting Periods on Firearm Suicide Rates Among Men: Out of Treatment Sample

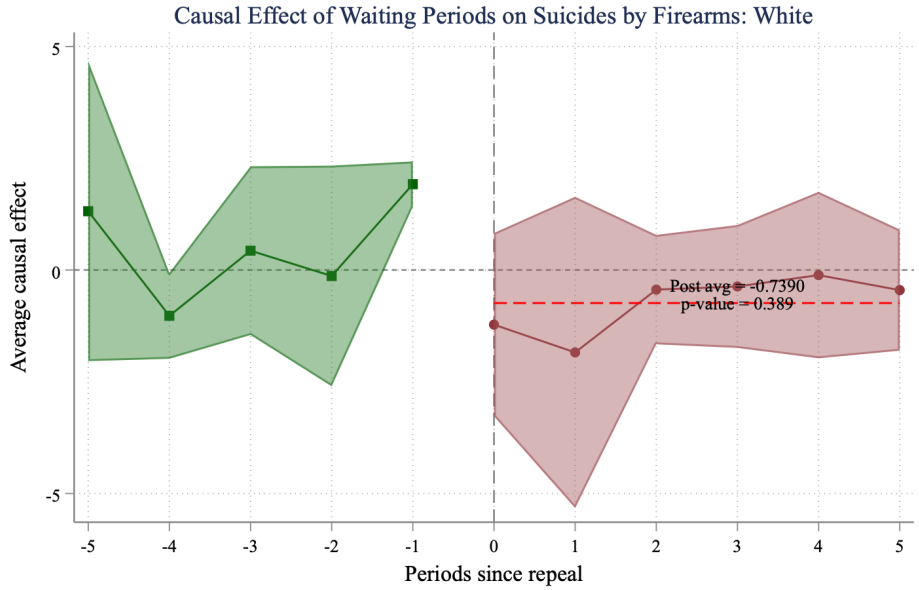


Note: This figure focuses on the male population using a sample of states that move out of treatment, showing how waiting period laws affect firearm suicide rates for men specifically. The sample includes 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 20. Effect of Waiting Periods on Firearm Suicide Rates Among Adults Aged 55+: Out of Treatment Sample

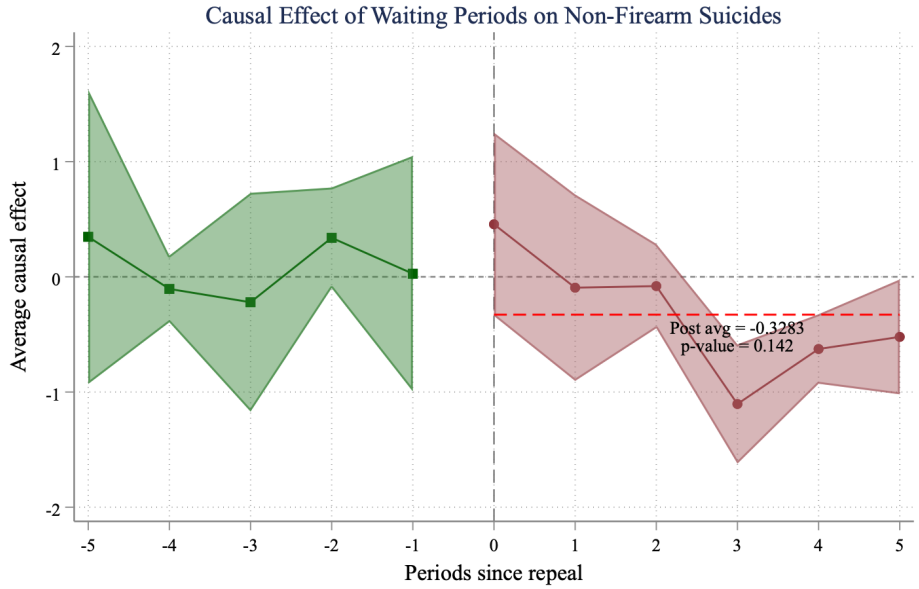
Note: This event study estimates the policy effect on older adults using a sample of states that move out of treatment, focusing on a group at elevated suicide risk. The sample comprises 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 21. Effect of Waiting Periods on Firearm Suicide Rates Among White Individuals: Out of Treatment Sample



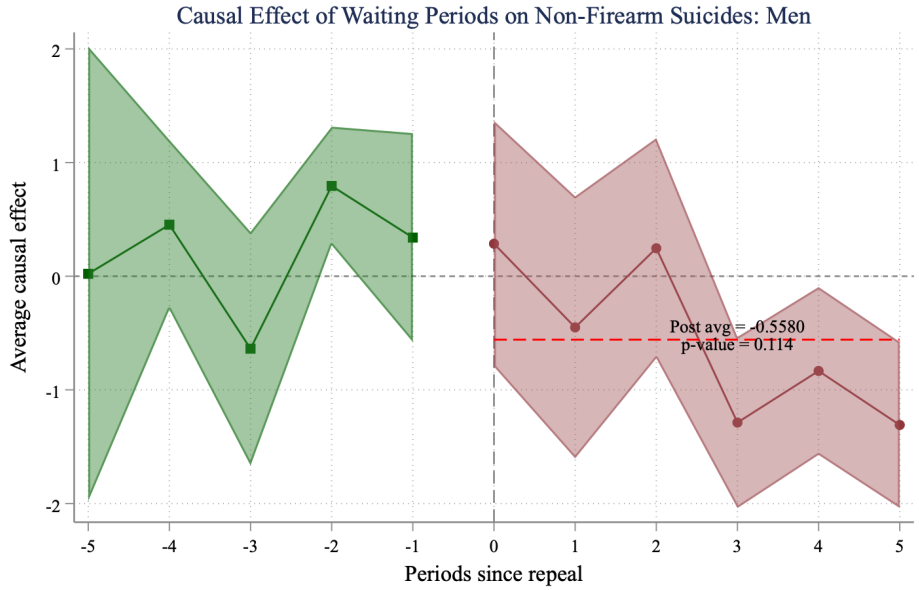
Note: This figure examines firearm suicide trends for white individuals using a sample of states that move out of treatment. The sample includes 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 22. Effect of Waiting Periods on Non-Firearm Suicide Rates: Out of Treatment Sample



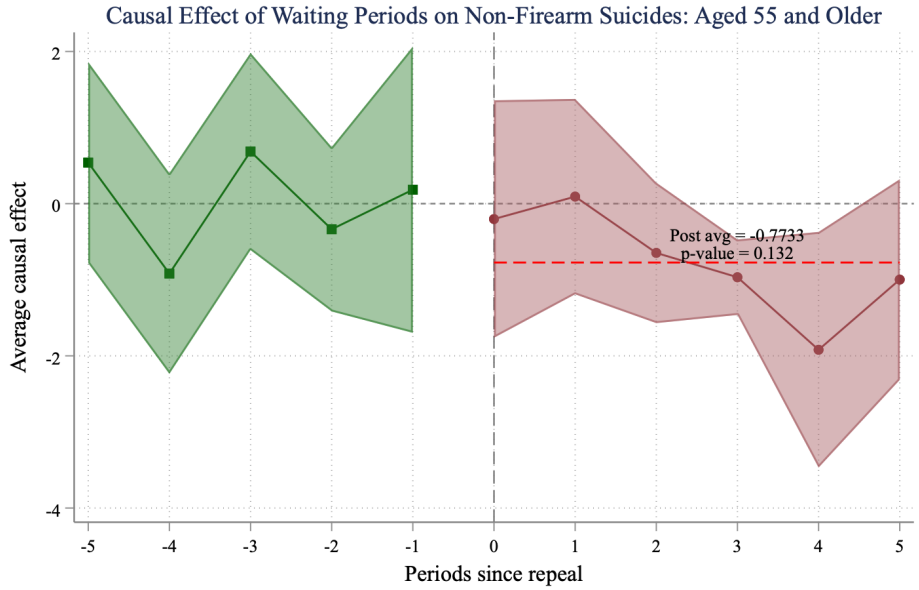
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties using a sample of states that move out of treatment. The sample comprises 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 23. Effect of Waiting Periods on Non-Firearm Suicide Rates Among Men: Out of Treatment Sample



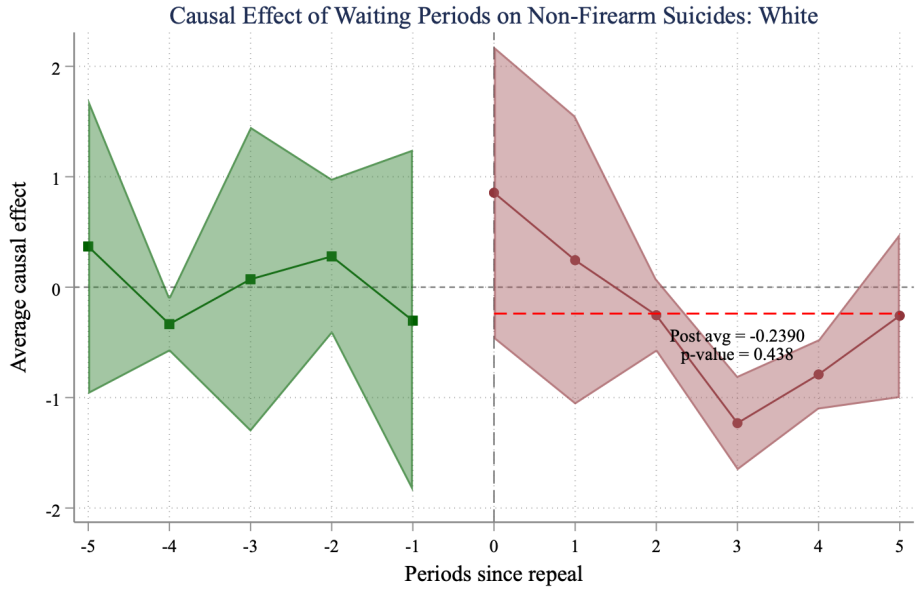
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among men using a sample of states that move out of treatment. The sample includes 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 24. Effect of Waiting Periods on Non-Firearm Suicide Rates Among Adults Aged 55+: Out of Treatment Sample



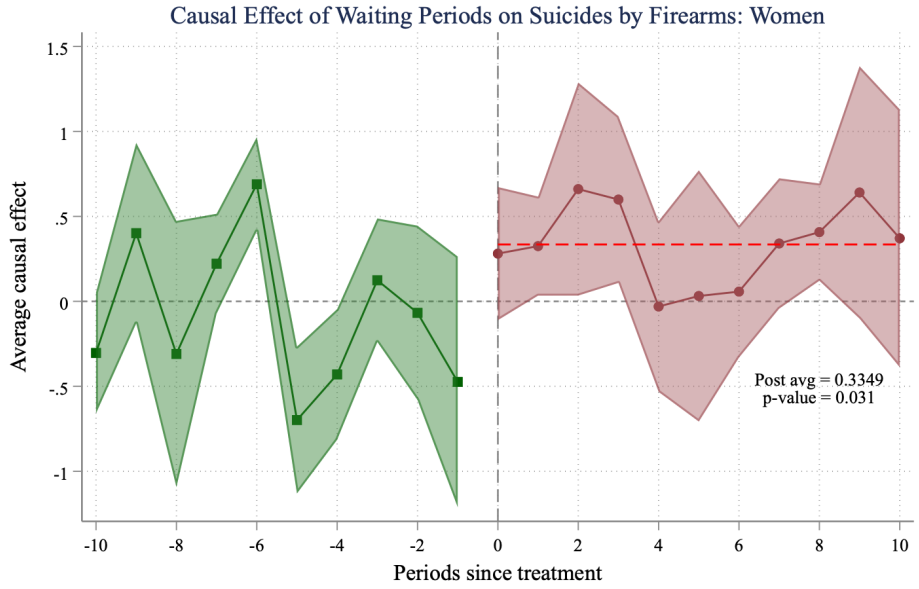
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among adults aged 55+ using a sample of states that move out of treatment. The sample comprises 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 25. Effect of Waiting Periods on Non-Firearm Suicide Rates Among White Individuals: Out of Treatment Sample



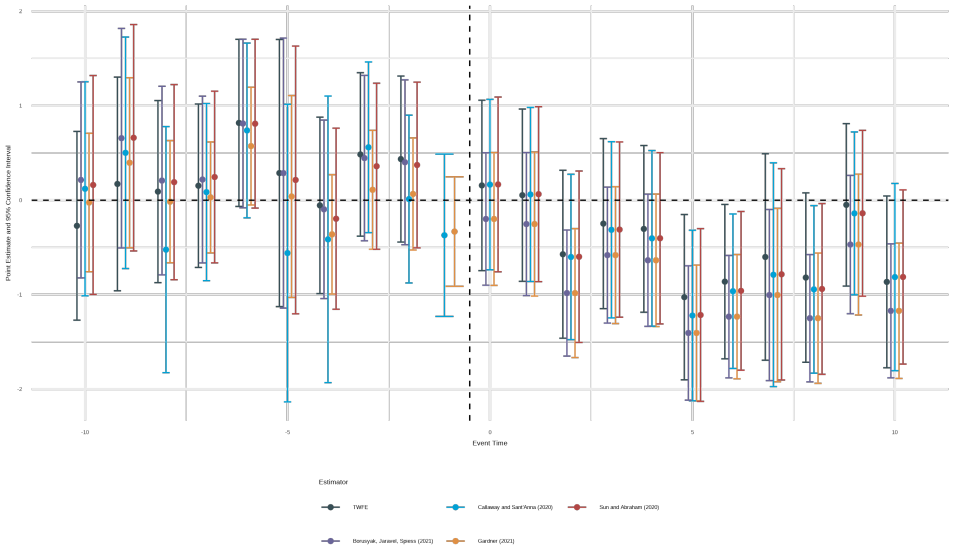
This figure shows the dynamic effects of waiting period laws on all other suicide rates (non-firearm) across US counties among White individuals using a sample of states that move out of treatment. The sample includes 23 states (13 treatment, 13 control) that experience either one policy transition from treatment to no-treatment, or always treated. Standard errors are bootstrapped and clustered at the state level.

Fig. 26. Effect of Waiting Periods on Firearm Suicide Rates Among Women



This figure shows the dynamic effects of waiting period laws on firearm suicide rates among women across US counties. Each point represents the estimated difference in firearm suicide rates relative to the year of policy adoption (year 0), with 95% confidence intervals. Standard errors are bootstrapped and clustered at the state level.

Fig. 27. Effect of Waiting Periods on Firearm Suicide Rates: Different Types of Estimators



This figure compares treatment effect estimates from multiple difference-in-differences estimators for the effect of waiting periods on firearm suicides. Different estimators include Callaway and Sant'Anna (2021), Roth and Sant'Anna (2023), Borusyak, Jaravel, and Spiess (2024), Gardner (2022), Sun and Abraham (2021), and two-way fixed effects (TWFE) to assess sensitivity of results to methodological choices in staggered adoption designs.

Table 1: Summary Statistics for County-Year Data

	Mean	SD	N
Total Suicide Rate (per 100k)	13.64	14.14	80,313
Firearm Suicide Rate (per 100k)	8.45	11.00	80,313
Men's Firearm Suicide Rate (per 100k)	14.86	19.92	80,313
Firearm Suicide Rate Aged 55+ (per 100k)	24.99	221.42	80,313
White Individuals' Firearm Suicide Rate (per 100k)	16.11	97.85	80,313
Non-Firearm Suicide Rate (per 100k)	5.19	7.55	80,313
Female Population Share	0.50	0.02	80,313

Notes: Data come from the National Vital Statistics System (NVSS), 1959–2019. Suicide rates are expressed per 100,000 population. "Female Population Share", "College Educated", and "Below Poverty Line" are proportions. SD denotes standard deviation. N is the number of county–year observations (80,313).

Table 2: Covariate Balance: Treated vs. Control Groups

Variable	Control (N = 60,300)	Treated (N = 20,013)	Difference
Total Suicide Rate (per 100k)	13.752	13.314	-0.437 (1.453)
Firearm Suicide Rate (per 100k)	8.720	7.639	-1.081 (0.990)
Men's Firearm Suicide Rate (per 100k)	15.320	13.461	-1.859 (1.582)
Firearm Suicide Rate Aged 55+ (per 100k)	27.112	18.612	-8.500* (5.145)
White Individuals' Firearm Suicide Rate (per 100k)	17.282	12.561	-4.721 (3.517)
Non-Firearm Suicide Rate (per 100k)	5.032	5.675	0.644 (0.566)
Female Population Share	0.505	0.503	-0.002 (0.003)
College Educated Share	0.133	0.090	-0.043** (0.020)
Below Poverty Line Share	0.282	0.124	-0.157*** (0.040)

Note: Data come from the National Vital Statistics System (NVSS), 1959–2019. Suicide rates are expressed per 100,000 population. "Female Population Share", "College Educated", and "Below Poverty Line" are proportions. Standard errors clustered at the state level. Significance: * p$<$0.1, ** p$<$0.05, *** p$<$0.01.

Table 3: Treatment Cohorts by State

Treatment Cohort	State Count	Percent of States	State Abbreviations
1965 Cohort	2	3.92%	CA, CT
1966 Cohort	1	1.96%	MD
1967 Cohort	1	1.96%	IL
1969 Cohort	1	1.96%	WA
1973 Cohort	3	5.88%	IN, OR, PA
1976 Cohort	1	1.96%	WI
1977 Cohort	1	1.96%	MN
1979 Cohort	1	1.96%	NJ
1988 Cohort	1	1.96%	HI
1991 Cohort	1	1.96%	FL
1994 Cohort	19	37.25%	AK, AZ, AR, GA, ID, KS, KY, LA, ME, MS, MT, NH, NM, ND, OK, TX, VT, WV, WY
Always Treated	5	9.8%	AL, DC, RI, SD, TN
Never Treated	14	27.45%	CO, DE, IA, MA, MI, MO, NE, NV, NY, NC, OH, SC, UT, VA
Total	51	100%	

Notes: Data from RAND State Firearm Law Database. Always treated are the states that had a waiting period law in place before at 1959, the first year of vital statistics data that is available. They include states that might have moved to a less restrictive waiting periods or abolished them.

Table 4: Delay Mechanism: Waiting Days and Firearm Suicides (TWFE)

	(1) All	(2) Men	(3) Age 55+	(4) White
Waiting days	-0.063^{*}	-0.096	-5.630^{**}	-4.123^{**}
	(0.036)	(0.060)	(2.617)	(1.879)
Baseline mean	8.45	14.86	24.99	16.11
Pct change per day (%)	-0.74	-0.64	-22.53	-25.60
N	80,313	80,313	80,313	80,313

County and year fixed effects; SEs clustered at the state level.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$