Examining the Association between Racialized Economic Threat and White Suicide in the United States, 2000–2016

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Abstract

Suicide is steadily rising. Many blamed worsening economic conditions for this trend. Sociological theory established clear pathways between joblessness and suicide, focused on status threat, shame, and consequent disruption of social relationships. However, recent empirical research provides little support for a link between unemployment and suicide. I attempt to reconcile this contradiction by focusing on white suicide and white employment-to-population ratio. Whiteness is not just a default category, but a pervasive ideology that amplifies the effects of status loss. The white employment-to-population ratio represents a form of racialized economic threat and accounts for discouraged workers who have exited the labor force. I use longitudinal hybrid models with U.S. state-level data, 2000 to 2016 and find that decreasing employment is associated with increasing suicide among the white population and white men. I discuss this study's contributions to the literature on suicide and joblessness and the emerging scholarship on whiteness and health.

Keywords

employment, health, status threat, suicide, whiteness

Since the late 1990s, the U.S. has experienced a steady increase in suicides. This rise has been particularly pronounced among non-Hispanic white Americans, a group already experiencing very high rates of suicide, and has contributed to the decline and stagnation of life expectancy (Harper, Riddell, and King 2021). In light of this trend, suicide attracted growing attention from scholars and the general public, with the prevailing narrative centering on worsening economic conditions.

Sociological literature on suicide has long established the role of financial strain and joblessness in particular as a risk factor (Stack 2000; Stack and Haas 1984; Wray, Colen, and Pescosolido 2011). More recent contributions examined the theoretical mechanisms linking joblessness to suicide, focusing on the emotional responses triggered by status loss and the cultural context that constrains these responses (Abrutyn 2019; Abrutyn and Mueller 2014, 2018). However, researchers have provided mixed evidence about the association between unemployment and suicide (DeFina and Hannon 2015). Analyses of the years immediately following the Great Recession support the harmful effects of unemployment (Phillips and Nugent 2014; Reeves et al. 2012), but several recent studies find only partial or moderating effects, or no effect at all (Dow et al. 2020; Gertner, Rotter, and Shafer 2019; Houle and Light 2017; Kaufman et al. 2020; Rambotti 2020).

To reconcile this apparent contradiction, I address two considerations: the racialized nature of suicide and the measurement of unemployment. Suicide varies by race, but the literature on suicide does not adequately consider the system of racialized hierarchies governing the U.S. Whiteness is not just a reference category, but also a pervasive ideology influencing the social, cultural, and emotional foundations of suicide (Abrutyn 2019; Abrutyn and Mueller 2014, 2018). Joblessness is likely more damning for white Americans who are at the top of the

racialized hierarchy. Thus, I suggest considering race-specific explanations for race-specific outcomes: here, I focus on white Americans. Second, the vast majority of suicide research measures joblessness with the unemployment rate, but this measure has important limitations. Thus, I use the employment-to-population rate. Employment is a more concrete concept than unemployment (Leon 1981). Unlike the unemployment rate, the employment-to-population ratio includes "discouraged workers," who have given up searching for work. Their inclusion is fundamental, because while the unemployment rate recovered after the Great Recession, the employment-to-population ratio did not, leading analysts to believe that the decline in unemployment is masking "continued weakness in the labor market" (Kalleberg and Wachter 2017:3).

I combine these considerations by analyzing the relationship between white suicide rate and white employment-to-population ratio, which has been conceptualized as a racialized economic threat (Bjorklund, Davis, and Pfaffendorf 2018; Eitle, D'Alessio, and Stolzenberg 2002). I use longitudinal hybrid models with U.S. state data in 2000 to 2016. I find that the declining share of employed white Americans is associated with an increase in white suicide. The association is consistent for both-sex and male suicide rates, less so for female rates. The relationship is robust across several model specifications. Post-regression estimates based only on white men show that a 1% increase in the white male employment-to-population ratio is associated with about 2,850 fewer suicides throughout the entire study period (2000 to 16), and a one-standard deviation increase is associated with about 13,535 fewer suicides.

BACKGROUND

Suicide on The Rise: Worrisome Trends and Renewed Sociological Interest

In 2015, life expectancy at birth declined in the U.S. for the first time since 1993 (Xu et al. 2016). The country's life expectancy, which is already significantly lower than other rich democracies, has yet to recover. While there is no single cause, it is clear that suicide contributes to this worrisome trend (Murphy et al. 2018). Unlike most other rich democracies, the U.S. experienced a clear rise in suicides during the past few years (Stack 2018). Suicide is constantly among the top 10 causes of deaths in the U.S.: It increased 33% between 1999 and 2019, and in 2019 alone, it accounted for 47,500 deaths, which corresponds to 1 death every 11 minutes (CDC 2021).

While sociology, and in particular the classic work of Émile Durkheim (1897), was central to the early days of suicidology, this interest seemed to vanish, based on the low number of papers on suicide published in sociology compared to other disciplines (Stack and Bowman 2011). However, the recent rise in suicide attracted significant attention in the scholarly community and in the general public, often in conjunction with the discussion of the so-called "deaths of despair" (Case and Deaton 2015), and led to a renewed interest in the topic, which resulted in important empirical and theoretical contributions.

In short, Durkheim theorized that suicide rates are a function of the structure of social relationships and that they reflect a "collective breakdown" of social groups' "integrative and regulative functions" (Mueller et al. 2021:2). Consistent with how social scientists currently approach population health, he suggested that an optimal way to prevent suicide was to implement societal changes that foster healthy relationships. Durkheim's theory has recently been expanded to integrate meso-level explanations derived from adjacent subfields, such as

cultural sociology and the sociology of emotions (Abrutyn 2019; Abrutyn and Mueller 2014, 2018). These arguments provide valuable context for macro-level research, such as the analysis proposed in this study.

Suicide and Joblessness: Mechanisms, Evidence, and a Puzzle

Sociological literature on suicide has long emphasized how job loss represents a major (if not the major) cause of financial strain, which may increase the risk of suicidal behavior (Stack 2000; Wray et al. 2011). Unemployment may lead to suicide via several mechanisms, which may also affect those who have not personally lost a job, such as fear of unemployment or induced stress in relationships (Stack and Haas 1984). Abrutyn and Mueller (2014) delve deeper into the emotional foundations of suicide, drawing direct connections to Durkheim's theory. They find that job loss represents "a real, perceived, or threatened status change . . . that is incongruent with one's own self-conception" (Abrutyn and Mueller 2014:336). This status threat usually triggers feelings of shame, which are often repressed and replaced by anger. These emotions disrupt the individual's "social relationships with individuals, groups, or society as a whole" (Abrutyn 2019:113). When anger becomes intolerable, it can target others or the self. In sum, job loss is associated with "anomic suicides . . . driven by shame and anger at real or imagined violations of social expectations that threaten one's sense of self" (Abrutyn and Mueller 2014:336).

These notions often found empirical support, especially in times of economic decline, such as the recession of 1981 to '82 (Stack and Haas 1984) and the Great Recession of 2007 to 2009 (Phillips and Nugent 2014; Reeves et al. 2012). However, several recent analyses have found inconsistent evidence, especially when focusing on regional variation within the U.S. Several studies found no direct link between the unemployment rate and suicide (Gertner et al. 2019; Houle and Light 2017; Rambotti 2020). Other studies found only evidence of a partial

(Dow et al. 2020) or moderating (Kaufman et al. 2020) effect of unemployment. This last study (Kaufman et al. 2020) is particularly interesting: It finds a significant interaction between unemployment and minimum wage, but the interaction shows some surprising patterns. The lowest suicide rates occur at the highest levels of minimum wage, as one would expect, and unemployment, contrary to one's expectations. DeFina and Hannon (2015) found evidence of a changing relationship between unemployment and suicide. Analyzing data from the years 1979 to 2010, they find no effect pre-1995 and a harmful effect post-1995, which declines but remains significant during the Great Recession.

These studies suggest that the relationship between suicide and unemployment in the U.S. is more complex than one might expect and may be subject to peculiar dynamics. For example, Stack (2018) proposes that the particular configurations of income inequality and welfare policies in the U.S. may explain why suicide is rising in the U.S. but falling in most European countries. I propose two additional considerations which may account for the inconsistent findings on unemployment and suicide: One has to do with race, while the other is a measurement issue.

Racialized Economic Threat: A Solution to the Puzzle?

The U.S. is a society organized around racialized symbolic hierarchies that are extremely powerful and articulated in ways that aim to deny overt racism while contributing to the persistence of racial inequalities (Bonilla-Silva 2006). In the context of these hierarchies, perceptions of one's economic achievement or decline, and the arguments used to explain these outcomes, draw significantly on different racial experiences (MacLeod 1987). Losing or being unable to find employment may carry different meanings and trigger different emotional responses if a person belongs to the racially dominant group, as it is harder to blame society or

discrimination (Stack 2000). Declining employment among the dominant group (that is, white Americans) may be all the more damning, as it intensifies the loss of privilege and social standing. Abrutyn and Mueller (2014) propose that people are more vulnerable to shame and anomie when their disrupted identity or status are strongly salient and pervasive, and the salience of whiteness is unmistakable (Doane and Bonilla-Silva 2003).

Furthermore, in the context of racialized social hierarchies, it may be worth conceptualizing labor market outcomes as a (real or imagined) competition between racial and ethnic groups. After all, narratives construed to protect American workers often center around the threat presented by non-white workers who are "stealing" jobs through immigration, outsourcing, or affirmative action (Brown 2013; Harper and Reskin 2005; Santa Ana 1999; Wang and Liu 2018). While these narratives may provide arguments to blame others rather than the self, they likely intensify the perception of status decline among white Americans. Thus, it is important to focus on the association between race-specific unemployment and race-specific suicide.

The second consideration pertains to measurement. Most existing research has focused on unemployment rate (Oyesanya, Lopez-Morinigo, and Dutta 2015). I propose that the employment-to-population ratio is a better measure. I provide more details in the data section, but it is important to understand that, while obviously related, these measures differ in one meaningful way. The employment-to-population ratio accounts for people who are not in the labor force. This category includes people who were not employed at the time of the survey, were not on temporary layoff, and were not actively looking for work during the four weeks prior to the survey. Among them, there are the so-called "discouraged workers," who report discouragement about job opportunities as the main reason for not looking for work. The Bureau

of Labor Statistics lists the following as common answers given by discouraged workers (BLS 2021):

There are no jobs available or none for which they would qualify.

They have been unable to find work in the past.

They lack the education, training, or experience needed for available jobs.

Employers think that they are too young or too old, or they are subject to some other type of discrimination.

It is reasonable to believe that discouraged workers may be particularly vulnerable to the emotional responses associated with suicidal behavior because duration of unemployment is associated with suicide (Stack and Haas 1984). Abrutyn and Mueller (2014) argue that prolonged feelings of shame may lead to sadness and isolation and that these mixed types of ego-anomic suicides are likely more common than pure types. A recent review of suicide during economic recessions (Oyesanya et al. 2015) found only one study that used the employment-to-population ratio in 22 former Soviet Bloc countries (Brainerd 2001). The lack of research on this measure is problematic because discouraged workers are not counted in the official unemployment rate. It is also interesting to notice that while the unemployment rate returned to the levels recorded before the Great Recession, the (white) employment-to-population ratio failed to recover (see the online appendix: Figure A1).

To address the two points discussed in this section, I focus my attention on the relationship between the white suicide rate and white employment-to-population ratio. White Americans are the dominant racial group in the U.S. Despite this privileged position, their suicide rate is historically high, only lower than American Indians and Alaska Natives (Harper et al. 2021). The white employment-to-population ratio has been conceptualized as a form of

racialized economic threat (Bjorklund et al. 2018; Eitle et al. 2002) because declines in white labor force participation represent a disruption of the racialized hierarchy. Perceived white economic decline has been linked to punitive state policies and conservative/white backlash (Bjorklund et al. 2018; Willer, Feinberg, and Wetts 2016). I ask whether it is also linked to suicide, hence my first hypothesis:

Hypothesis 1: White employment-to-population ratio will be negatively associated with white suicide rate.

Furthermore, because suicide rates vary dramatically by gender, I disaggregate the previous hypothesis using not only race-specific but also gender-specific measures:

Hypothesis 1a: White male employment-to-population ratio will be negatively associated with white male suicide rate.

Hypothesis 1b: White female employment-to-population ratio will be negatively associated with white female suicide rate.

Lastly, the gendered nature of suicide is not simply a function of different rates. Abrutyn and Mueller (2018:60) discuss how current conceptions of masculinity generate "narrow views about what it means to be a man (or a white man) and what is expected" and additional pressures on men. Failure to meet such strict expectations may lead men to feel engulfed with shame, exposing them to fatalistic suicides (Abrutyn and Mueller 2014). This leads to my last hypothesis:

Hypothesis 2: The relationship between gender-specific employment-to-population ratio and suicide rate is stronger among white men than white women.

DATA AND METHODS

Data

The data for this study came from multiple sources. I collected repeated observations of the variables described below for the 50 U.S. states over a 17-year period (2000 to 2016). I chose this time period (2000 to 2016) based on several considerations. First, the choice was constrained by data availability for all the variables of interest. Furthermore, this time period is interesting for three reasons. First, there has been a notable increase in suicide rates in the U.S. during this time period. Second, the Great Recession, which occurred between 2007 and 2009, introduced an economic shock that created the economic conditions for a rise in suicides (Chang et al. 2013). Finally, this period included the election (and re-election) of Barack Obama as the first black president of the U.S. During this time, race became more salient in political discourse and political beliefs became more polarized by racial attitudes (Tesler 2016). These sentiments ultimately favored the election of Donald Trump in 2016 (Abramowitz and McCoy 2019; Hooghe and Dassonneville 2018).

Outcome

My outcome was *white suicide rate*. Suicide is patterned by gender. For this reason, I used overall, male, and female white suicide rates. I collected these measures from the Centers for Disease Control and Prevention. The rates represent the number of deaths by suicide or self-harm (codes: U03, X60-X84, and Y87.0) per 100,000 people identified as white. To be consistent with the focal independent variable (see below), the rate was not disaggregated by ethnicity. The rates were not adjusted by age.

Focal Independent Variable

Drawing on scholarship on racialized economic threat (Bjorklund et al. 2018; Eitle et al. 2002), I used the state-level *white employment-to-population ratio* as my focal independent variable. The source was the Bureau of Labor Statistics. The formula for the employment-to-population ratio is: (employed ÷ civilian noninstitutional population) × 100. The civilian noninstitutional population includes both people in and out of the labor force. I used the default denominator, which includes the civilian noninstitutional population of age 16 or above.

The employment-to-population ratio presents several advantages (Leon 1981). It has a straightforward interpretation: "it is the share of adults who are employed," and it "gauges employment against *population*, and not the *labor force*" (Donovan 2015:5). Analysts observed that "being employed is an observable experience, while being unemployed often lacks that same concreteness" (Leon 1981:36). The employment-to-population ratio is preferrable to the official unemployment rate because the latter excludes people who are not currently in the labor force, such as "discouraged workers." Instead, the employment-to-population ratio includes discouraged workers, who at the time of the survey were neither employed nor looking for work during the last four weeks. This is important because the unemployment rate returned to pre-Great Recession levels, but the employment-to-population did not (see the online appendix: Figure A1). It is still debated whether "falling unemployment rate indicates a strengthening labor market, or signals an exodus of discouraged workers from the labor force" (Donovan 2015:5).

As I did for the outcome, I collected overall and gender-specific rates for the white employment-to-population ratio. The Bureau of Labor Statistics disaggregates this measure by race or by ethnicity, not by race and ethnicity combined. Thus, I could only collect information on this variable for people who identify as white, whether of Hispanic origin or not. While not

ideal, I ensured that the outcome and focal independent variables were consistent by collecting suicide rates for white people of any ethnicity (see above).

Covariates

I included a comprehensive set of covariates that were relevant to the relationship between racialized economic threat and suicide. I used the *revised 1960 to 2016 citizen ideology series* (Berry et al. 1998) to measure the political orientation of each state's citizens on a liberal–conservative continuum. The measure varies between 0 and 100, and larger values indicate more liberal ideology. More details about the measure can be found in Berry et al. (1998).

Four variables measured important socio-economic characteristics of the states: income inequality, measured as the *Gini coefficient* and available through the World Inequality Database (Alvaredo et al. n.d.; Frank et al. 2015); *poverty rate* as the % of state population living below the Federal poverty line (source: U.S. Census); all-industry *GDP* per capita, chained at 2009 dollars (source: U.S. Bureau of Economic Analysis); and share of the state population receiving *SNAP* or *Food Stamp* benefits (University of Kentucky Center for Poverty Research 2016). I included SNAP or Food Stamp benefits because previous research found a link between SNAP and suicide (Rambotti 2020) and because recent studies showed that SNAP substantially reduces extreme poverty (Brady and Parolin 2020; Parolin and Brady 2019).

Next, I included a measure of *education*, the share of population aged 25 and older with a bachelor's degree or higher, from IPUMS USA (Ruggles et al. 2017). To do this, I collected data at the individual level from the American Community Survey between 2000 and 2016 (more than 41 million data points). I weighted these data and aggregated them at the state-year level. I followed this procedure for all the variables collected from IPUMS USA, including several variables described below.

Suicide is sensitive to household composition: Thus, I controlled for the share of *single-person households* and the *average household size* (source: IPUMS USA). Additional risk factors included access to guns, measured as *household firearm ownership* rate (Schell et al. 2020), alcohol consumption, which I attempted to control for by including *adult binge drinking* (source: Behavioral Risk Factor Surveillance System Prevalence and Trends Data), and *male/female ratio* (source: IPUMS USA). Suicide is also sensitive to *population* size (source: U.S. Census Bureau) and *population density* (source: the American Community Survey). Both of these measures were highly skewed, thus I used them in logged form.

I included several measures of the population composition. First, I used two variables describing the state racial-ethnic breakdown, specifically *% white* and *% Hispanic* (source: IPUMS USA). I controlled for the share of population who identify as white (vs. the reference category: nonwhite) for two reasons. First, one could believe that the state-level suicide rate is merely a function of the state white population. Second, one could assume that a shrinking white population may be conceptualized as another dimension of racialized threat. Instead, I added the share of population identifying as Latino or Hispanic (vs. not Latino or Hispanic) to the models because as mentioned above, the outcome and key predictors provide rates for white people of any ethnicity: Thus, it is important to control for it.

Finally, I introduced *age* as a covariate by using the state share of population in three distinct age groups: under 18, 18 to 34, and over 64 (reference category: age 35 to 64). Often researchers adjust only mortality rates by age (that is, the outcome). Instead, using this approach ensured that the whole model is adjusted by age (Milyo and Mellor 2003).

Analytical Strategy

Longitudinal models are well equipped for the kind of pooled cross-sectional data used in this study. Because my models included several variables that present little to no variation within states, I used hybrid longitudinal models (Allison 2009), which I estimated using the user-written *xthybrid* command in Stata 16 (Schunck and Perales 2017).

Hybrid models are a particular type of clustered models, which are suitable for panel data. In fact, the data of this study can be conceptualized as nested, where level 1 is represented by state-years (n = 850) and level 2 by states (n = 50). Hybrid models involve particularly flexible modeling specifications as they split variable effects into within-cluster and between-cluster effects. After model estimation, it is possible to test whether within and between effects are identical using a Wald test. A null hypothesis suggests that the within and between effects do not differ significantly: Thus, researchers should use the more efficient random effects.

Hybrid models are a useful alternative to random or fixed effects. Random effects average out the effects of a predictor across and within clusters (e.g., states), while fixed effects ignore variation between clusters. Furthermore, while fixed effects models have been a typical choice for longitudinal data, they are subject to increasing scrutiny (Hill et al. 2019; Kropko and Kubinec 2020). Hill et al. (2019) discuss several limitations of fixed effects models, some of which are directly relevant to this study. For example, fixed effects models "are notoriously susceptible to attenuation bias from measurement error" (Angrist and Pischke 2008:225), which may especially occur when using aggregate data.

Furthermore, fixed effects models need predictors that present adequate variation over time because "one cannot predict a variable from a constant and will do a poor job of trying to predict a variable from a near constant" (Treiman 2009:370). Unfortunately, the literature offers

no exact threshold, and researchers rarely acknowledge this issue. To avoid overestimating the within-state effect of variables that present minimal change over time, I set a minimum threshold for within-state variance to 5%. For variables that present variation within states below this threshold, hybrid models estimated only random effects. The command *xthybrid* reports the variance within clusters: For example, approximately 0% of the total variance in population density was within clusters as opposed to around 26% for the white employment-to-population ratio.

I modeled time in two ways. First, I included years as a continuous variable in both the simple and multiple models. Second, I added a random slope for years. In hybrid models, random slopes "allow the (within) effects of the level-one variables to vary between clusters" (Schunck and Perales 2017:98). In other words, the random slope for time allows the (within-state) effect of time on suicide to vary across states.

I used robust standard errors clustered by state to address potential issues with heteroscedasticity and autocorrelation (Hoechle 2007). Finally, I performed several robustness checks to assess whether the findings were sensitive to alternative specifications.

RESULTS

Table 1 reports the descriptive statistics of the variables used in this study. The number of observations for each variable varies between 850 and 771 (50 states over 17 years). The white female suicide rate is the only variable with missing observations. If there are fewer than 20 suicides, the rate is unreliable; while if there are fewer than 10 suicides, the CDC suppresses the rate to ensure anonymity.

[Table 1 about here]

For descriptive purposes, I show the geographical distribution of my outcome and focal predictor: white suicide rate and white employment-to-population ratio. For brevity, I show only the maps reporting data for both sexes. Figure 1 shows the change in white suicide rates from 2000 to 2016. White suicide increased the most in the West. The largest increase is in Wyoming (8.5 units per 100,000) followed by Montana, Idaho, and New Hampshire (8.4 each). Only one state shows a decline: In Hawaii, the rate is 0.4 units lower in 2016 than 2000. However, this may be misleading, because white suicide rate in Hawaii in 2015 was 5.9 units higher than in 2000.

[Figure 1 about here]

The largest declines in the white employment-to-population ratio occurred across multiple regions, with the South being particularly represented (see Figure 2). The largest decline is in Hawaii (10.1%), followed by Mississippi (9.2%), Nevada (8.9%), Georgia (8.8%), and Alabama (8.3%). Only North Dakota has a slight increase from 2000 to 2016 (0.6%).

[Figure 2 about here]

Figure 3 shows the pooled cross-sectional association between the white suicide rate and white employment-to-population ratio across 850 state-years. As expected, the variables are negatively and statistically significantly correlated ($p < 0.001$), although their association is weak ($r = -0.30$).

[Figure 3 about here]

It may be interesting to see a visualization of the within-cluster component on this relationship. In Figure 4, I map the correlation between white suicide rate and white employment-to-population ratio by state. For each of these correlation coefficients, N is 17 years. The negative relationship observed in Figure 3 is consistent across almost every state. Only one

state shows a positive but not statistically significant correlation (North Dakota, $r = 0.23$). The correlation is negative but not significant in Louisiana (−0.38) and Nebraska (−0.46). In the remaining 47 states, the correlation between white suicide rate and white employment-to-population ratio is negative and significant at the 0.05 level, with the strongest values in Texas and Alabama (−0.90). Eighteen states representing every U.S. region have a correlation equal to or stronger than −0.80. Figure A2 in the online appendix shows the linear trend of this relationship by state.

[Figure 4 about here]

Table 2 reports the results of the regression models predicting white suicide rate. The simple model includes only the key predictor and year, while the multiple models include the full set of covariates. I report within- and between-state effects only for variables whose effects are significantly different according to the Wald chi-square test. If the test does not yield statistically significant results, I only report the more efficient random effects (Schunck and Perales 2017). Two-tailed t statistics are in parentheses. I used robust standard errors clustered by state to account for heteroscedasticity and autocorrelation (Hoechle 2007).

The findings support Hypothesis 1. In the simple model, the white employment-to-population ratio is statistically significant and negatively associated with the white suicide rate: a 1% increase in the predictor is associated with a 0.09 decrease in deaths by suicide per 100,000 people. In the full model, not only does this association hold, but it also increases slightly in magnitude ($b = -0.14$, $p < 0.001$). This suggests the presence of a mild suppression effect caused by one or more of the confounders. Interestingly, the effect of time also increases in magnitude (the within-cluster coefficient for year goes from 0.29 to 0.38). The year random slope shows that the within-state effect of time varies significantly across states. It is also noteworthy that

time is the only variable whose coefficient must be split into within- and between-cluster effects because there is no variation between states[1], while the random coefficient is recommended by the Wald chi-square for all other variables. Two other variables are statistically significant and negatively associated with the outcome: the state share of SNAP recipients and the population density. Both of these effects align with previously published work on suicide (Baller and Richardson 2002; Rambotti 2020).

[Table 2 about here]

In Table 3, I report the models estimating suicide rate for white men only. The results support Hypothesis 1a, and are similar to those pertaining to the white population as a whole. The employment-to-population ratio for white men is statistically significant and negatively associated with white male suicide both in the simple ($b = -0.11$, $p < 0.05$) and in the multiple model ($b = -0.14$, $p < 0.01$). The effect of time and population density is consistent with what we observed above: White male suicide increases over time and is more prevalent in sparsely populated areas. Several differences appear: For example, the share of SNAP recipients is not statistically significant, although the p-value is just above the 0.05 alpha level ($p = 0.067$). Consistent with literature that has documented the saliency of firearm use for suicide among men (Stack and Wasserman 2009), the household firearm ownership rate is positively associated with white male suicide ($b = 0.06$, $p < 0.01$). The full model also shows two variables with significantly different effects. States with larger shares of single-person households and population below 18 years of age have lower suicide rates among white men, but changes in these two variables within states are not associated with changes in the outcome.

[Table 3 about here]

Table 4 reports the results of regression models predicting the white female suicide rate. In this case, the results differ from what I presented above and provide limited support for Hypothesis 1b. In the simple model, there is no association between the outcome and the random effect of white female employment-to-population ratio ($b = -0.04$, $p > 0.05$). In the full model, the between and within effects differ: While the between effect is negatively associated with the outcome ($b = -0.18$, $p < 0.01$), the within effect is not ($b = -0.02$, $p > 0.05$). This means that states where a larger share of white women are employed have lower suicide rates among white women, but change in this share is not associated with change in the relative suicide rates, although these rates increased over time as indicated by the year effect[2]. Four variables show statistically significant associations between their random effects and the outcome: GDP, % of SNAP recipients, population, and population density. Holding all the covariates constant, states with higher values of these variables have lower suicide rates. Two age groups show distinct effects: States with larger shares of population below 18 years of age and above 64 have lower rates of white female suicide rates; within-state change in the share of the youngest population is not associated with the outcome, while within-state increases in the share of the oldest population are associated with increased suicide. Finally, it should also be noted that the white female suicide rate has missing values in eight states: Alaska, Delaware, Hawaii, North Dakota, Rhode Island, South Dakota, Vermont, and Wyoming. I reran the analysis after dropping these states. The full model only required the main predictor's random effect, which was not statistically significant.

[Table 4 about here]

I conducted several sensitivity analyses to test whether the findings are susceptible to different model specifications. I estimated fully adjusted models for each outcome with the

following alternative specifications: state fixed effects, state and year fixed effects, random effects, and hybrid models with a 10% threshold for minimum within-state variance. In the online appendix (Figures A3 to A5), I report the findings for my main predictors using coefficient plots, which show the regression coefficient estimates and 95% confidence intervals. The main findings are strongly consistent under these specifications when analyzing white overall and male suicide. Instead, the results for white female suicide depend largely on model specification. The regression coefficients for the main predictor are largely nonsignificant when using one-way (state) fixed effects and two way (state and year) fixed effects. This is similar to the within-state effect observed in the hybrid models with a 5% threshold for within-state variation (Table 4) and with a 10% threshold (Figure A5). When using random effects models, the coefficient is only marginally significant ($p = 0.077$). The results shown in Figure A5 effectively highlight a major strength of hybrid models, that is, their ability to separate effects that occur between or within clusters, provided that these differ significantly. For researchers who value robustness across model specifications, these sensitivity analyses suggest using caution when making claims about the association observed among white women. In this sense, Hypothesis 2, which posits that the relationship between the gender-specific employment-to-population ratio and suicide rate is stronger in white men than white women, finds support.

Additional analyses not reported here show that adding the official unemployment rate does not fundamentally alter the picture. Finally, while I focus on white Americans, I also conducted the same analysis using outcomes and key predictors pertaining to Black Americans. While the sample size decreases substantially, especially for women, I find no significant association, consistent with previous studies (Siddiqi et al. 2019).

Lastly, it is important to assess the real-world implications of these findings. To do so, I estimated the potential number of avoidable suicides associated with an increase in the main predictor. I focused on the estimates for white men, which is the group exhibiting the steadiest association. In the full model, a 1% increase in the white male employment-to-population ratio is associated with a predicted 0.14 decrease in suicide rate: This translates to about 2,850 fewer suicides during the study period (2000 to 16). A one standard deviation increase (approximately 4.75%) in the white male employment-to-population ratio is associated with about 13,535 fewer suicides in the same time period.

DISCUSSION

Previous sociological literature on suicide provides convincing arguments for why joblessness should increase the risk of suicide (Stack 2000; Stack and Haas 1984; Wray et al. 2011) by triggering emotional responses of shame and anger that disrupt one's social relationships and sense of self (Abrutyn 2019; Abrutyn and Mueller 2014, 2018). However, the evidence supporting this relationship is mixed (Oyesanya et al. 2015), and several recent studies found no direct association (Gertner et al. 2019; Houle and Light 2017; Rambotti 2020) or only a partial (Dow et al. 2020) or moderating association (Kaufman et al. 2020) between unemployment and suicide.

In this paper, I propose that a better understanding of the relationship between suicide and joblessness emerges by considering the context of racial hierarchies that govern social life in the U.S., and by focusing not on unemployment but on employment, a more concrete and "observable experience" (Leon 1981:36). Focusing the attention on the racially dominant group, white Americans, I find that racialized economic threat (Bjorklund et al. 2018; Eitle et al. 2002), measured as the share of the white population that is actively employed, is significantly and

negatively associated with the suicide rates of white Americans. This relationship is steadier among white men, whose expectations for success, help-seeking, and emotional responses are severely constrained by pervasive norms of masculinity (Abrutyn and Mueller 2018; Cleary 2012).

This study makes two major contributions. First, it contributes to the scholarship on joblessness and suicide by using the employment-to-population ratio. This measure includes discouraged workers, who may be particularly vulnerable to suicide (Stack and Haas 1984). I only found one paper that explicitly uses this measure in an analysis of former Soviet Bloc countries (Brainerd 2001). Future research may test whether the employment-to-population ratio is consistent across different geographical and temporal contexts or population subgroups. Surely, the trend in the employment-to-population ratio fits the upward trend in suicide. As a recent review pointed out, the evidence suggests that suicides increased almost continuously "through unemployment cycling up and down during the past 25 years makes it difficult to square with recent trends" (Harper et al. 2021:395). This contradiction disappears when using the employment-to-population ratio, which has steadily declined over the last two decades.

Second, this study adds to a recent but already influential framework that investigates "how a racialized social system of whiteness benefits whites' health in some ways while adversely affecting whites' health in other ways" (Malat, Mayorga-Gallo, and Williams 2018:148). Whites represent the dominant racial group in the U.S., but their overall health is far from excellent. Malat et al. (2018) posit that racial capitalism in the U.S. influences a number of societal conditions (most importantly, social policies), individual experiences, and psychological responses that affect whites' health positively and negatively. The association between lower levels of employment and higher levels of suicide among white Americans, presented in this

paper, is certainly mediated by narratives and experiences built around ideas of whiteness, such as shame for failing to meet stringent cultural expectations (Abrutyn and Mueller 2018), strong identification with gun culture (Carlson 2015; Metzl 2019), feelings of disconnection from one's own land (Hochschild 2018), and lack of narratives of redemption (Keyes 2009).

This study presents some limitations. First and foremost, ecological data limit the possibility to test mechanisms discussed in the literature. Ideally, one would work with a large, individual-level random sample. This is a well-known data limitation in the literature on suicide, which may be circumvented by investigating different aspects of suicidal behavior, such as ideation, or by focusing on institutions other than the state, such as high school students (Abrutyn, Mueller, and Osborne 2020; Mueller and Abrutyn 2016). Second, the non-experimental nature of this analysis should prevent the reader from interpreting association as causation. Concerns about reverse causality should be mitigated by considering that employment is highly likely a direct function of the labor market's strength. However, one could hypothesize that awareness of increasing suicides may impact employment, for example by contributing to the perception of white decline and thus, discouraging workers from seeking employment. Future research may adopt an experimental design to specifically test this hypothesis.

Another limitation of observational analysis is possible omitted variable bias. I attempt to address this by adding an extensive set of covariates. In the sensitivity analysis, I also use longitudinal fixed effects models, which account for time-invariant unobserved characteristics. However, even these models cannot account for unobserved time-varying traits (Hill et al. 2019). Third, ecological data may be exposed to attenuation bias because of measurement error. Measurement errors bias the regression coefficient towards zero (Meng, Wu, and Zhan 2016). In an additional analysis using errors-in-variables regression (Lockwood and McCaffrey 2020), I

found tentative evidence of this bias, which does not affect the validity of the findings, but may underestimate the predictor's effect.

There are other ways in which future research can build upon this study. For example, it may be possible to create more comprehensive measures of racial economic threat, similar to Homan's (2019) measurement of structural sexism. Another possibility could involve the use of Google searches, a method that has already been successfully applied to the study of population health (Chae et al. 2015). Future research could also explore the conditions under which economic threat leads to suicide. Analysis of racial subgroups may become quickly unfeasible once the absolute number of deaths by suicide makes the rate unreliable, but an alternative could include exploring the statistical interaction between relevant variables (DeFina and Hannon 2015; Kaufman et al. 2020).

Furthermore, future research should investigate recent trends in suicide. From 2018 to 2019, suicide rates declined among white Americans but not people of color (Stone, Jones, and Mack 2021). Recent analyses find that this trend continued during the COVID-19 pandemic (Pattani 2021). A crucial question is what accounts for this trend? Research largely neglected to consider how reparative (Goffman 1967) or integrative (Braithwaite 1989) rituals reduce anomie. For example, the Republican vote has often been interpreted as an indicator of status threat (Siddiqi et al. 2019), but one may hypothesize that Trump's 2016 victory may have triggered a sense of collective effervescence among his supporters, sustained through numerous rallies over the following years. Perhaps, these rituals—largely amplified by the media—partially offset the harmful effect of racialized threat on suicide, providing some Trump supporters with a renewed sense of social primacy and thus, contributing to a decline in white suicide.

Reducing suicide among white Americans remains challenging. Research shows that generous welfare policies are generally associated with better health (Montez et al. 2020) and linked to lower suicide rates. However, racial resentment towards "underserving" people of color decreases support for these policies (Metzl 2019). The challenge for policymakers is to implement policies that may reduce suicide while also fostering racial justice and avoiding white backlash. Ambitious interventions designed to revert decades of an astounding rise in economic disparities and declining upward mobility (Chetty et al. 2017; Kenworthy 2018) are more necessary than ever.

SUPPLEMENTAL MATERIAL

Additional supporting information may be found in the online version of this article.

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NOTES

1. It is important to clarify a question about the effect of time: Unless there are missing observations, time varies only within states and not between states. For this reason, year's between-effect is omitted in Tables 2 and 3.
2. Because several states have missing values for white female suicide rates, some variation in the effect of year also exists between the states (see previous note). In this case, the Wald test supports the use of random effect, which makes it unfeasible to add a year random slope. In fact, in these hybrid models, a random slope can be added only to within-cluster effects (Schunck and Perales 2017). I performed a robustness check where I split the effect into between- and within-state and added the year random slope. The results are substantially consistent. When I reran the analysis after dropping the states with missing values, the (within) effect of time is consistent with the (random) effect reported in Table 4.

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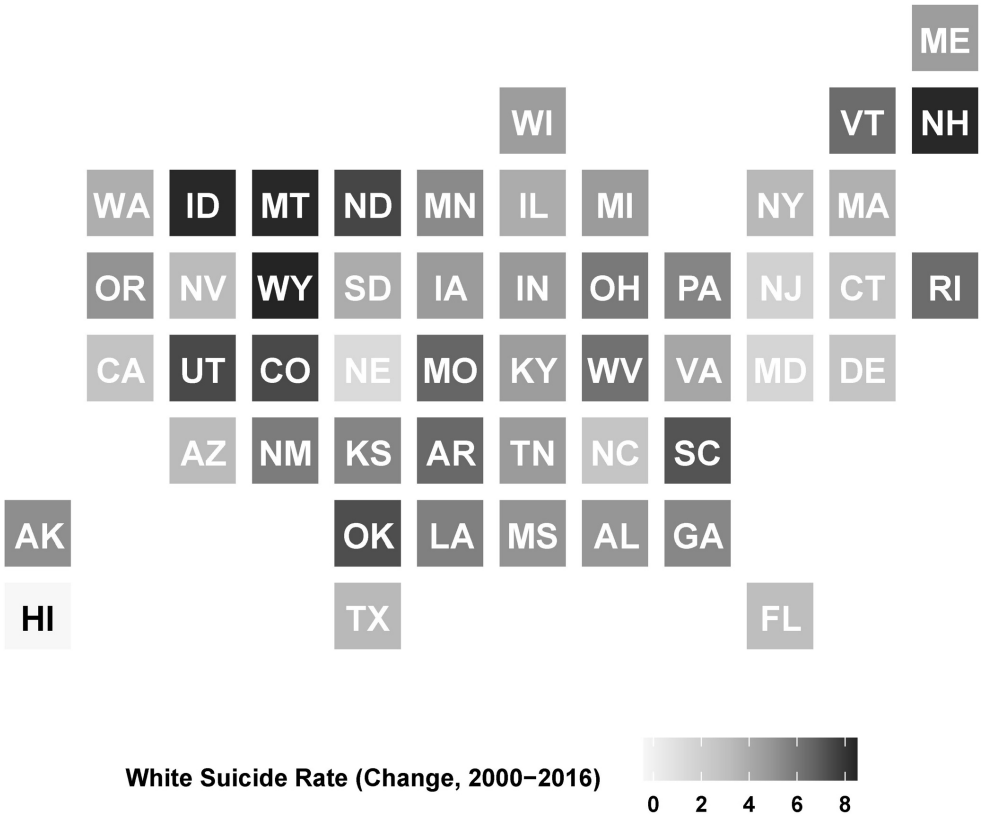
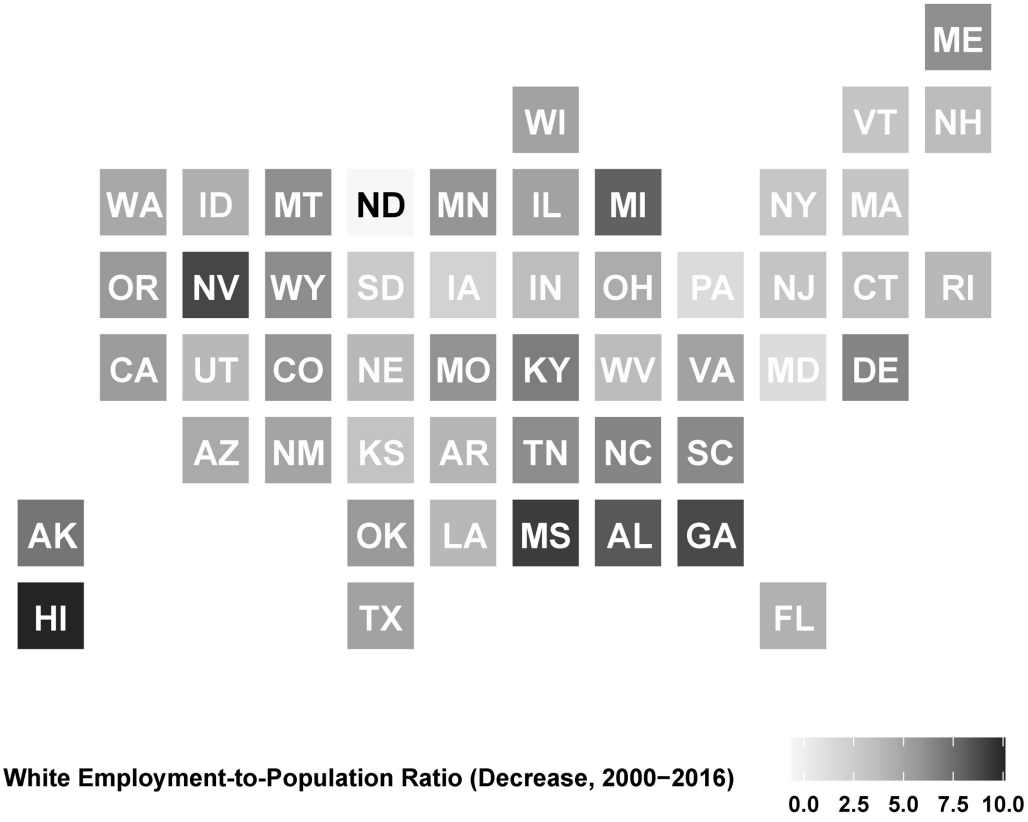
TABLES AND FIGURES

Figure 1: White Suicide Rate, Change 2000–2016



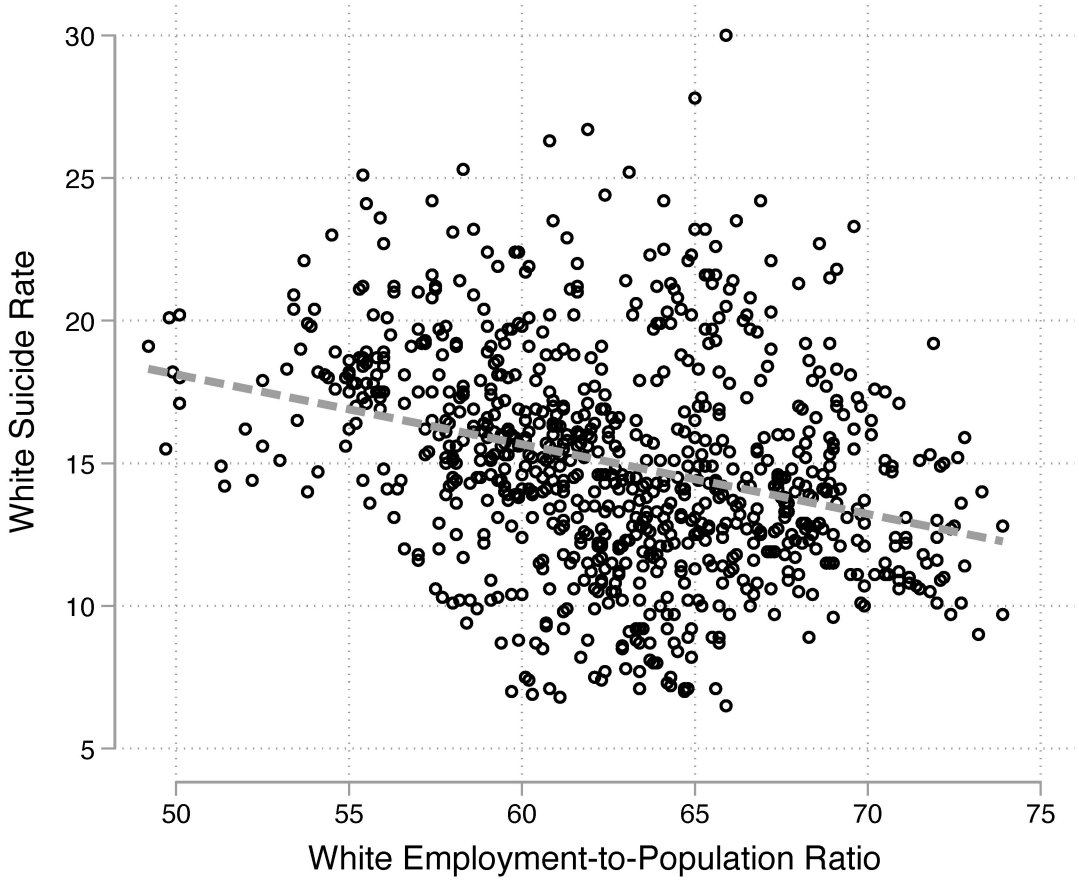
Data Source: CDC.

Figure 2: White Employment-to-Population Ratio, Decrease 2000–2016



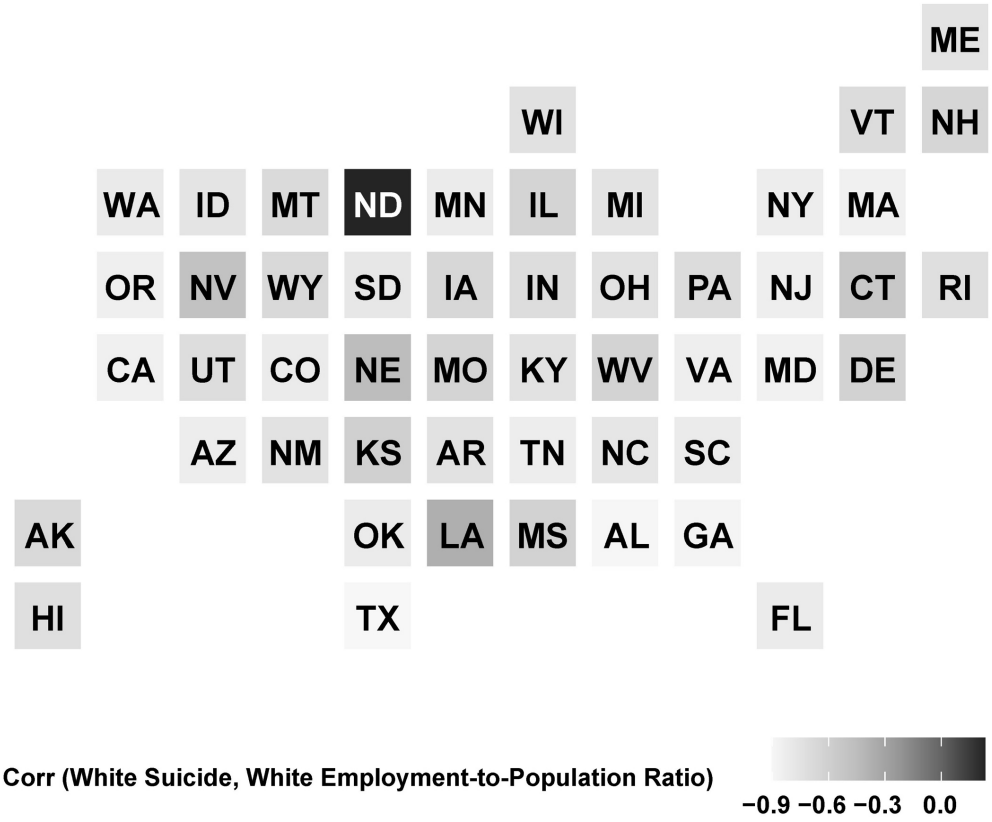
Data Source: U.S. Bureau of Labor Statistics.

Figure 3: White Suicide Rate by White Employment-to-Population Ratio (*n* = 850 State-Years)



Data Sources: CDC and U.S. Bureau of Labor Statistics.

Figure 4: Correlation Map of White Suicide Rate and White Employment-to-Population Ratio



Data Sources: CDC and U.S. Bureau of Labor Statistics.

Table 1: Descriptive Statistics

	N	Mean	SD	Min.	Max.
White Suicide	850	14.98	3.80	6.50	30.00
White Male Suicide	850	23.91	5.82	10.10	47.70
White Female Suicide	771	6.27	1.97	2.00	12.90
White Employment-Pop. Ratio	850	62.78	4.66	49.20	73.90
White Male Employment–Pop. Ratio	850	68.97	4.75	53.60	80.10
White Female Employment–Pop. Ratio	850	56.85	5.08	44.80	69.60
State Citizen Ideology	850	51.09	15.88	8.45	97.00
Gini	850	45.05	2.17	38.41	51.38
Poverty Rate	850	12.61	3.34	4.50	23.10
GDP	850	48939.41	9508.11	30564.00	79894.00
% SNAP Recipients	850	10.63	4.50	2.83	22.65
Education	850	27.43	5.09	14.91	43.37
% Single-person Households	850	12.41	1.93	5.41	18.09
Average Household Size	850	3.27	.22	2.86	4.22
Household Firearm Ownership	850	38.89	13.64	3.40	69.40
Male/Female Ratio	850	.99	.03	.93	1.15
Adult Binge Drinking	850	15.74	3.28	6.60	26.35
Population (log)	850	15.13	1.01	13.11	17.49
Population Density (log)	850	4.46	1.40	.09	7.10
% Age < 18	850	24.30	1.94	19.43	32.18
% Age 18–34	850	23.16	1.36	19.63	28.86
% Age > 64	850	13.11	1.86	5.69	19.05
% White	850	80.89	11.46	38.79	98.30
% Hispanic	850	9.83	9.82	.55	48.55

Data Sources: CDC, U.S. Bureau of Labor Statistics, Berry et al. (1998), World Inequality Database, U.S. Census Bureau, U.S. Bureau of Economic Analysis, University of Kentucky Center for Poverty Research, IPUMS USA, Schell et al. (2020), BRFSS, American Community Survey.

Table 2: Hybrid Models Predicting White Suicide Death Rate 2000–16

	(1)		(2)	
White Employment-to-Pop Ratio	−.09**	(−2.72)	−.14***	(−2.48)
State Citizen Ideology			−.01	(−1.80)
Gini			−.04	(−.60)
Poverty Rate			.00	(.04)
GDP			−.00	(−.69)
% SNAP Recipients			−.09*	(−2.30)
Education			−.09	(−1.42)
% Single-person Households			−.09	(−1.11)
Average Household Size			−1.41	(−1.11)
Household Firearm Ownership			.02	(1.36)
Male/Female Ratio			2.31	(.32)
Adult Binge Drinking			.01	(.49)
Population (log)			−.55	(−1.69)
Population Density (log)			−1.24***	(−2.80)
% Age < 18			.15	(.75)
% Age 18–34			.16	(1.11)
% Age > 64			.31	(1.59)
% White			−.04	(−1.69)
% Hispanic			.02	(.55)
Year				
Between Effect	Omitted		Omitted	
Within Effect	.29***	(16.11)	.38***	(7.15)
Constant	20.83***	(9.56)	38.35*	(2.44)
Random Components				
State-Level Random Intercept	10.30***	(4.87)	3.61***	(3.33)
Year Random Slope	.01***	(4.99)	.01**	(3.19)
N	850		850	
Log Pseudolikelihood	−1434.36		−1387.60	
Wald Chi-square	484.94		1290.38	
p	.000		.000	
AIC	2880.72		2823.20	

Data Sources: CDC, U.S. Bureau of Labor Statistics, Berry et al. (1998), World Inequality Database, U.S. Census Bureau, U.S. Bureau of Economic Analysis, University of Kentucky Center for Poverty Research, IPUMS USA, Schell et al. (2020), BRFSS, American Community Survey.

Note: Level 1: 850 state-years; Level 2: 50 states; Two-tailed t statistics in parentheses; Robust SE clustered by state; Unless otherwise noted, the coefficients represent random effects.

* $p < .05$, ** $p < .01$, *** $p < .001$

Table 3: Hybrid Models Predicting White Male Suicide Death Rate 2000–16

	(1)		(2)	
White Male Employment-to-Pop Ratio	−.11*	(−2.45)	−.14**	(−2.12)
State Citizen Ideology			−.03	(−1.90)
Gini			−.02	(−.19)
Poverty Rate			.01	(.22)
GDP			−.00	(−.59)
% SNAP Recipients			−.12	(−1.83)
Education			−.18	(−1.82)
% Single-person Households				
Between Effect			−1.95***	(−2.12)
Within Effect			−.05	(−.41)
Average Household Size			−2.91	(−1.88)
Household Firearm Ownership			.06**	(2.62)
Male/Female Ratio			−2.32	(−.27)
Adult Binge Drinking			.03	(.45)
Population (log)			−.43	(−.93)
Population Density (log)			−2.43***	(−2.35)
% Age < 18				
Between Effect			−1.04*	(−2.43)
Within Effect			.65	(1.61)
% Age 18–34			.45	(1.96)
% Age > 64			.52	(1.74)
% White			−.03	(−1.17)
% Hispanic			.01	(.18)
Year				
Between Effect	Omitted		Omitted	
Within Effect	.39***	(14.68)	.58***	(6.44)
Constant	31.41***	(10.44)	108.98***	(5.10)
Random Components				
State-Level Random Intercept	24.69***	(4.85)	5.73***	(3.82)
Year Random Slope	.02***	(3.98)	.01**	(2.85)
N	850		850	
Log Pseudolikelihood	−1882.19		−1830.25	
Wald Chi-square	368.96		2014.68	
p	.000		.000	
AIC	3776.39		3712.49	

Data Sources: CDC, U.S. Bureau of Labor Statistics, Berry et al. (1998), World Inequality Database, U.S. Census Bureau, U.S. Bureau of Economic Analysis, University of Kentucky Center for Poverty Research, IPUMS USA, Schell et al. (2020), BRFSS, American Community Survey.

Note: Level 1: 850 state-years; Level 2: 50 states; Two-tailed t statistics in parentheses; Robust SE clustered by state; Unless otherwise noted, the coefficients represent random effects.

* $p < .05$, ** $p < .01$, *** $p < .001$

Table 4: Hybrid Models Predicting White Female Suicide Death Rate 2000–16

	(1)		(2)	
White Female Employment-to-Pop Ratio				
Random Effect	−.04	(−1.64)		
Between Effect			−.18**	(−2.86)
Within Effect			−.02	(−.60)
State Citizen Ideology			−.01	(−1.33)
Gini			−.08	(−1.71)
Poverty Rate			−.01	(−.29)
GDP			−.00*	(−2.21)
% SNAP Recipients			−.05*	(−2.21)
Education			−.02	(−.39)
% Single-person Households			−.02	(−.30)
Average Household Size			−.14	(−.12)
Household Firearm Ownership			.00	(.02)
Male/Female Ratio			2.65	(.75)
Adult Binge Drinking			.01	(.56)
Population (log)			−.60**	(−2.63)
Population Density (log)			−.50*	(−2.56)
% Age < 18				
Between Effect			−.34*	(−1.99)
Within Effect			.08	(.50)
% Age 18–34			−.12	(−.93)
% Age > 64				
Between Effect			−.41*	(−2.27)
Within Effect			.37**	(2.76)
% White			−.02	(−.84)
% Hispanic			.03	(1.85)
Year	.18***	(14.82)	.21***	(5.46)
Constant	−257.31***	(−14.00)	−276.36***	(−2.95)
Random Components				
State-Level Random Intercept	2.48***	(5.12)	.81***	(4.62)
N	771		771	
Log Pseudolikelihood	−287.98		−235.38	
Wald Chi-square	422.01		1372.07	
p	.000		.000	
AIC	1985.97		1922.75	

Data Sources: CDC, U.S. Bureau of Labor Statistics, Berry et al. (1998), World Inequality Database, U.S. Census Bureau, U.S. Bureau of Economic Analysis, University of Kentucky Center for Poverty Research, IPUMS USA, Schell et al. (2020), BRFSS, American Community Survey.

Note: Level 1: 771 state-years; Level 2: 50 states; Two-tailed t statistics in parentheses; Robust SE clustered by state; Unless otherwise noted, the coefficients represent random effects.

* $p < .05$, ** $p < .01$, *** $p < .001$