



Whites' Racial Resentment and Perceived Relative Discrimination Interactively Predict Participation

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Abstract

Scholarship in American politics finds whites' racial resentment and status threat predict their vote choice. However, research in social identity indicates that such variables should negatively correlate with participation, attenuating resentful whites' political power. We resolve this contradiction by studying these variables interactively, using data from the 2012, 2016, and 2020 American National Election Studies. Our primary independent variables are racial resentment and a measure of relative discrimination that captures whites' perceptions of inequality relative to other racial/ethnic groups. Both constituent variables are negatively associated with participation. Interactively, they are positively associated with political participation. These relationships persist when we predict voter turnout, an index measuring non-electoral participation, and an index measuring civic engagement. In 2012, our interaction term moderates the negative impact of racial resentment in relative discrimination. In elections contested by Donald Trump, our interaction term has a positive substantive effect. Our findings advance scholarship on white political behavior. First, they identify a relationship between whites' racial attitudes and participation, advancing a research program that primarily examines vote choice. Second, they clarify the relationship between different white racial attitudes. Next, they detail the conditional relationship between whites' prejudice and politics and how it has changed over time.

Keywords White racial attitudes · Political participation · Voter turnout · Prejudice

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Introduction

The ¹ 2016 election and subsequent violent political mobilization of whites² ignited academic and public interest in the role of racial attitudes in white political behavior. While some scholarship traces the origins of whites' support for Donald Trump, other research examines how whites' racial attitudes shape their partisanship, public opinion, and vote choice since Barack Obama's 2008 election. This flurry of scholarship identifies two principal constructs that predict whites' political behavior: out-group antipathy and in-group attachment (Sides et al., 2019; Berry et al., 2019; Jardina, 2020). These findings broadly suggest that American whites may, increasingly, behave as a politically conscious bloc motivated by a sense of status loss (Gest, 2016; Jardina, 2019) and racial resentment (Tesler, 2016). While this research examines vote choice, those effects only matter if prejudiced and threatened whites participate in politics. Research into prejudice and Social Identity Theory (SIT) suggest that prejudice and threats to whites' group status may both *attenuate* participation among whites. In short, research on white racial attitudes and identity suggests the potential for greater political coherence among American whites, while research into SIT and prejudice predict the opposite.

We address this contradiction by examining how white racial prejudice and perceived status threat *interactively* predict white political participation. We use data from the 2012, 2016, and 2020 American National Election Studies (ANES). Our independent variables are racial resentment and perceived relative group discrimination (hereafter relative discrimination), which captures whites' sense of their racial group's status relative to other racial or pan-ethnic groups. We find evidence that racial resentment and relative discrimination have independent, negative, and significant effects on white political participation. These findings are consistent across measures of turnout, non-voting electoral participation, and civic engagement. Yet, when these variables are interacted, the depressive effects of racial resentment and relative discrimination diminish. In some years, prejudiced and status-concerned whites only participate as much as non-prejudiced whites without status anxieties; in other years, they are *more* politically active. Together, these results suggest that whites' racial resentment or sense of status loss, alone, may not be enough to prompt whites' political participation. To participate, prejudiced whites may need to connect their negative stereotypes about people of color to beliefs that whites are losing status because of those perceived norm violations.

Here, we advance understanding of white political behavior in three ways. First, we clarify the relationship between whites' racial attitudes and political participation, advancing a research program that almost exclusively examines vote choice. Specifically, we present evidence that whites are demobilized by their senses of prejudice and status loss, but that whites who report the most prejudice *and* status loss, together, are not demobilized—and are often more politically active. Second, we clarify the relationship between various constructs that measure white racial

¹ Replication materials for this study, including code and data, are available on the *Political Behavior* Dataverse: <https://doi.org/10.7910/DVN/IISRW5>.

² Throughout this article, we refer to non-Hispanic whites in the U.S. as “whites.”

attitudes: white out-group animus and in-group solidarity are conceptually related but have analytically distinct influences on American politics (Jardina, 2020). Finally, we identify the conditional relationship between whites' prejudice and their political participation. Many whites may consider their racial prejudices to be without clear political redress or an insignificant political obstacle to their racial group. To be driven toward white identity politics, whites must also believe they are losing ground relative to the targets of their bigotry.

Linking Prejudice, Identity, and Participation among Whites

White Americans' racial attitudes shape their political attitudes and behavior. This relationship has been documented for decades (i.e., Kinder & Sears, 1981), but scholarship conducted since Obama's 2008 victory emphasizes that whites' racial attitudes, increasingly, predict their partisanship, public opinion, and vote choice (Abrajano & Hajnal, 2015; Chudy, 2020; Craig & Richeson, 2014; Jardina, 2019; Mutz, 2018; Sides et al., 2019; Tesler, 2016). This research program, broadly, employs two kinds of constructs: one measures whites' out-group antipathy, the other measures their in-group solidarity (Jardina, 2020). We review these constructs after discussing research on political participation in the U.S.

Theories of Political Participation

Extensive research examines why people participate in politics, especially by turning out to vote. Foundational work on participation drew on microeconomic theory to predict low turnout (Riker & Ordeshook, 1968). Scholars developed “resource” models that attribute participation—including non-voting participation—to material resources and civic skills. Both of those variables are bolstered by education, which has become a central predictor in models of participation (Verba et al., 1995; Wolfinger & Rosenstone, 1980). While education and wealth are associated with increased participation, some groups vote, participate, and protest more than their socioeconomic status predicts. Scholars identify two principal factors to explain increased participation among these groups. First, survey and experimental research highlight the important role of mobilization in participation (Rosenstone & Hansen, 1993; Han, 2009; García Bedolla & Michelson, 2012; Valenzuela & Michelson, 2016; Walker, 2020). Second, social identities can increase voting (Tate, 1994; Dawson, 1998; Barreto, 2010), political participation (García Bedolla, 2005; Wong, et al., 2011), and protest (Chong, 1991; Zepeda-Millán, 2017).

Racial Resentment and Participation

Other variables shape whites' politics. Researchers identify whites' negative attitudes toward out-groups as an animating factor in their political behavior. Classic studies of negative out-group affect include authoritarianism (Hetherington & Weiler, 2009), ethnocentrism (Kinder & Kam, 2009), and old-fashioned racism

(Piston, 2010; Tesler, 2013). Researchers employ multiple measures of prejudice (eds. Sears et al., 2000), most frequently white anti-Black prejudice (Hutchings & Valentino, 2004). Research into “new” or “symbolic” racism measures whites’ beliefs that African Americans contravene norms around work and individualism, and thus exploit the welfare state (DeSante, 2013; Gilens, 1999; Kinder & Sanders, 1996; Kinder & Sears, 1981; Mendelberg, 2001). Such research—and the studies we draw on most here—operationalize “symbolic racism” or “racial resentment” as explanatory variables (Kinder & Sanders, 1996; Kinder & Sears, 1981; but see: DeSante & Smith, 2020; Smith et al., 2019). Racial resentment predicts whites’ policy attitudes (Gilens, 1999; Kinder & Sanders, 1996; Kinder & Sears, 1981; Winter, 2008), partisanship (Enders & Scott, 2018; Tesler, 2016), candidate evaluations (Kinder & Dale-Riddle, 2012; Tesler & Sears, 2010), and vote choice (Hopkins, 2019; Jardina, 2020; Knuckey & Kim, 2015; Sides et al., 2019; Tesler, 2016).

The relationship between racial resentment and white political participation is less studied, but no less consequential: racial resentment’s influence on political attitudes only matters if the racially resentful participate in politics. Few studies examine the relationship between prejudice and participation and their findings are inconsistent. Racial prejudice can influence whites’ turnout in elections contested by Black candidates (Pasek et al., 2009; Petrow, 2010), and those contested between white candidates when race is especially salient (Luttig, 2017; Banda & Cassese, 2020). Evidence suggests that racial resentment demobilized white voters when Obama ran for president (Luttig, 2017; Pasek et al., 2014); however, the relationship may be conditioned by partisanship (Weller & Junn, 2018). Racial prejudice, racial resentment, and symbolic racism are positively associated with Republican turnout in 2008 (Pasek et al., 2009) and 2010 (Luttig, 2017), but not in 2016 (Banda & Cassese, 2020). Racial resentment is negatively associated with turnout among white Democrats, because it creates a cross-pressure between partisanship and racial attitudes (Krupnikov & Piston, 2015; Luttig, 2017; Banda & Cassese, 2020) making voting more cognitively burdensome. The negative relationship between racial prejudice and participation among white Democrats is consistent across distinct measures of racial prejudice, elections, candidates’ race (Krupnikov & Piston, 2015; Luttig, 2017), and multiple measures of political participation (Banda & Cassese, 2020).

These researchers find that the relationship between racial resentment and participation is conditioned by partisanship. We draw on other scholars to theorize an unconditional relationship. Whites who express racial resentment—due either to frustrations with growing racial equality or lower levels of education—also express less trust in government and lower evaluations of government (Filindra et al., 2022), which may decrease participation and civic engagement. Racially resentful whites are more likely to be alienated from democratic politics (Gest, 2016; Miller & Davis, 2020), support extreme parties (Gest, 2016), and express skepticism of state action (Kinder & Sanders, 1996). When controlling for white consciousness, whites who expressed higher levels of racial resentment were less likely to participate in the 2012 and 2016 elections (Berry et al., 2019). Therefore, the null and inconsistent associations between white prejudice and participation may stem from studies’ not modelling the distinct effects of in-group attachment. Whites who hold racially

resentful views towards Blacks, but do not connect those beliefs to threats to whites' status, may not consider their anti-Black stereotypes political, or capable of political redress. As such, we expect that once status threat is controlled for, racial resentment will have a negative relationship with white political participation.

Whites' Social Identities and Participation

Whites' racial attitudes are not limited to anti-Black prejudice. Other scholars have drawn on research into SIT to theorize how whites' racial identities shape their political behavior. Whites' concern for their dominant group status in the U.S. can have similar observable implications to prejudice (Jardina, 2019; Pérez et al., 2021). However, it is operationalized differently: as whites' sense of attachment to and commonality with other whites. Scholars drawing on SIT argue that attachment to social groups is a basic impulse (Tajfel, 1970) which prompts group members to pursue collective action to better the group's status (Tajfel, et al., 1971). Individuals may identify with multiple groups (race, gender, etc.) (Huddy, 2001), but due to social and cognitive constraints, tend to prioritize a few (Brewer, 1991; Gaertner et al., 1993; Roccas & Brewer, 2002; Garcia-Rios, Pedraza, & Wilcox-Archuleta, 2019; Yadon & Ostfeld, 2020). Most importantly, the choice about which identities to prioritize is driven by a desire to maximize self-esteem (Tajfel, 1970; Tajfel & Turner, 1979). Thus, scholars drawing on SIT emphasize that whites' identification with co-racials emerges from their awareness of, and attachment to, their group's dominant status (Wong & Cho, 2005; Branscombe et al., 2007; Craig & Richeson, 2014; Lopez Bunyasi, 2015; Schildkraut, 2017; Masuoka & Junn, 2013; Jardina, 2019; Pérez, et al. 2021; Yadon & Ostfeld, 2020). In some circumstances, when whites—incorrectly—perceive their racial identities are no longer an avenue to status, they withdraw from political participation (Gest, 2016; Schildkraut, 2010). In drawing on SIT, we provide reasons to expect that status threat may reduce whites' political participation.

Perceived threats to whites' status may signify to whites that their racial identities are no longer an avenue to self-esteem. Scholars note whites' anxieties about their eroding numerical (Abascal, 2020; Craig & Richeson, 2014; Enos, 2017), political (Parker & Barreto, 2013), economic (Gest, 2016; Mutz, 2018), and cultural (Abrajano & Hajnal, 2015) dominance. Whites are threatened by assertions that they benefit from privilege (Branscombe et al., 2007; Eibach & Keegan, 2006; Lowery et al., 2007; Powell et al., 2005) or racial inequality (Eibach & Keegan, 2006; Taylor Phillips & Lowery, 2015). In these studies, dependent variables include vote choice (Berry et al., 2020; Mutz, 2018), partisanship (Jardina, 2019), symbolic ideology (Craig & Richeson, 2014), and opposition to immigration (Abrajano & Hajnal, 2015; Gest, 2016; Hopkins, 2010). In short, there are extensive studies that show that whites perceive illusory threats to their racial group's status, and these perceived threats may shape white politics. We draw on the works in SIT cited above to deduce a second expectation: perceptions of threat to whites' group status, *ceteris paribus*, will be negatively associated with political participation. Whites' racial

identities flow from their status as the dominant racial group in the US. Perceptions of diminished status will decrease the group attachments that drive participation.

Connecting Prejudice and Privilege

Whites' racial identifications and their perceptions of Blacks are intertwined: whites' privileged group status is specifically implicated in historical subjection of African Americans (Gross, 2008; Lipsitz, 2006) and current expropriation of Blacks (Acharya et al., 2018). The research cited here presents further evidence of this relationship: whites' understanding of intergroup inequality is filtered through their attitudes toward Blacks (Kinder & Sanders, 1996; Simmons & Bobo, 2018). Earlier works on prejudice conceptualized individual-level stereotypes and group-level inequalities as intertwined preconditions for prejudice (Adorno et al., 1950, 150; Allport, 1958, 476; Blumer, 1958). Newer research into SIT supports this understanding: whites' attachment to their dominance (Branscombe et al., 2007) can drive their opposition to movements for racial equality (Wilkins & Kaiser, 2014; Eibach & Keegan, 2006), leading them to blame people of color for inequality (Taylor & Lowry, 2015; Kinder & Sanders, 1996). Applying this framework, we argue that whites who are high in racial resentment but perceive no racial group status threat, are less likely to be politically mobilized. As stated above, they do not feel that they (or other whites) are losing ground in employment, social standing, or access to scarce state resources through imagined Black violations of norms around work. When whites are high in perceived status threat but low in racial resentment, we argue that they will be unlikely to channel those anxieties into political participation. Racial resentment captures contemporary partisan-ideological contestation over race policy and racial inequality in America (Tesler, 2016): whites who express status anxiety but not racial resentment are unlikely to be mobilized by elites who focus on winning the support of racially resentful Americans (Jardina, 2020; Sides et al., 2019).

However, when whites are high in both racial resentment and status threat, we expect them to be more easily mobilized. These whites espouse anxieties about their group's dominant position *and* connect those threats to doubts about the deservingness of Blacks. Foundational studies on prejudice show that members of a dominant group who feel their privileges are eroding relative to groups perceived as inferior and undeserving espouse the most prejudice (Adorno et al., 1950; Allport, 1958; Blumer, 1958; Sidanius & Pratto, 1999; Yadon & Ostfeld, 2020). We expect prejudice to stoke political participation. We argue that whites who harbor both racial resentment toward Blacks *and* perceive white group interests are threatened will be more likely to believe that whites suffer discrimination *because* Blacks are contravening norms around work and individualism to which their group adheres. Because they connect such stereotypes to group-level inequality, these whites are most likely to believe that collective action should remedy the inequality they perceive whites face (Lee, 2007; Simmons & Bobo, 2018). Moreover, whites who express high levels of both racial resentment and relative discrimination will be conscious of the racial sorting of the parties and electorate, and thus perceive political participation

as a necessary strategy to remedy their illusory perceptions of white marginalization. Below, we discuss the data and methods with which we will test these expectations.

Data and Methods

To test our theory of the interactive relationship between racial resentment, perceived status threat, and white political participation, we use data from the 2012, 2016, and 2020 ANES. The ANES is, generally, the best publicly available dataset to examine the effects of white racial attitudes because it includes measures of in-group and out-group white racial attitudes. Numerous scholars examine the racialization of whites' political attitudes and behavior following the 2008 presidential election. These scholars, broadly speaking, argue that Obama's electoral success in 2008 (Tesler, 2016), the Tea Party's success in mobilizing reactionary whites (Parker & Barreto, 2013), demographic forecasts in the late 2000s (Jardina, 2019), or shifting campaign strategies (Sides et al., 2019) aligned whites' racial and political attitudes after 2008. Many of these researchers argue that Trump's election is a further manifestation of these trends. Thus, we also include data from 2020.

Using the 2012, 2016, and 2020 ANES allows us to construct a measure of white status threat, which we term relative discrimination (Berry, et al., 2020).³ For each white respondent, we create an index ranging from 0 to 1 measuring whether they believe that whites experience more or less discrimination than other racial groups. This, however, precludes us from using data before 2012: before the 2012 ANES, respondents were not asked about racial group discrimination. In 2012, the reference groups are whites, Blacks, and Latinos; after 2012, the reference groups are whites, Asian Americans, Blacks, and Latinos. Thus, in 2012, the index can take the following values: 0 (whites experience the least discrimination), 0.25 (whites experience less discrimination than one group, but the same as another), 0.5 (whites experience the same discrimination as other groups), 0.75 (whites experience more discrimination than one group, but the same as another), or 1 (whites experience the most discrimination). In 2016, the index functions similarly, but with finer gradations, as it includes another reference group (Asian Americans). Social scientists have used perceptions of discrimination to measure individuals' perceptions of racial inequality for decades (Chong & Kim, 2006; Conover, 1988; Gest, 2016; Schildkraut, 2010; Wong, 2018, etc.) validating our choice of the measure to operationalize perceived threats to white group status. Furthermore, scholarship on intergroup conflict stresses the importance of measuring individuals' perceptions of inequality *relative* to other groups in the polity (Gurr, 1970; Vanneman & Pettigrew, 1972; Kim, 2003; Masuoka & Junn, 2013; Newman, 2016; Ash, 2019). Our other principal independent variable is racial resentment. Racial resentment taps whites' skepticism that state

³ We do not argue that whites' perceptions of discrimination capture the reality of American racial inequality: whites continue to benefit from institutionalized white supremacy. Rather, we argue that whites' misperceptions of group inequality are politically meaningful. Extensive research indicates that misperceptions shape political behavior, for example: Cramer (2016) and Enders & Armaly (2019).

intervention to ameliorate racial inequality would be fair, because whites, allegedly, worked their way up “without any special favors.” (Kinder & Sanders, 1996). We recode the racial resentment index to a 0–1 scale to enable comparison with our similarly-coded measure of relative discrimination.

We use three dependent variables in our analyses. The first is a dichotomous, self-reported measure of turnout (e.g., Rosenstone & Hansen, 1993). We also include an index (0–1) measuring non-voting political participation⁴ and another index measuring civic engagement (0–1).⁵ Indices measuring both constructs are common in the literature on political participation (e.g., Verba et al., 1995).

In our analyses, our primary analytic method is unweighted controlled regressions, a common method used to analyze participation. We include only whites in our analyses. We specify logistic models when predicting voter turnout and OLS models when predicting political and civic engagement. We introduce control variables to demonstrate the stability of identified effects. We include the question wording used to generate all variables in Appendix A. Given the difficulty of substantively interpreting logistic regression coefficients, we predict the probability of voter turnout at set levels of (a) racial resentment, (b) relative discrimination, and (c) racial resentment across different levels of relative discrimination. We use a $p < 0.05$ significance threshold. We report findings from basic descriptive analyses in Appendix B.

Analysis

Here, we examine how racial resentment and relative discrimination predict turnout, non-voting electoral participation, and civic engagement. We regress relative discrimination, racial resentment, an interaction of those same terms, and a set of control variables that predict participation on each dependent variable. First, we include controls for age, income, education, gender, residential mobility, and homeownership. Next, we add controls for frequency of church attendance, union membership in respondents’ households, and contact by a political campaign. We repeat this process for the 2012, 2016, and 2020 ANES.

As can be seen in Table 1, racial resentment and relative discrimination are each negatively correlated with turnout. When racial resentment and relative discrimination are interacted, however, they are positively associated with the probability of

⁴ The sub-items for the non-voting political participation vary by year, but typically includes: persuading someone to vote for a candidate, going to political meetings, wearing campaign apparel, doing campaign work, donating to a campaign, donating to a political party, or donating to another political organization. From 2012 to 2020, the Cronbach’s α s for these variables are .66, .58, and .72, respectively. We discuss and interpret these results in Online Appendix B.

⁵ The sub-items for the civic engagement index vary, but includes: participation in a demonstration, attending a municipal board meeting, signing a petition, donating to a religious organization, donating to another organization, posting about politics, writing an op-ed, or contacting a politician. From 2012 to 2020, the Cronbach’s α s for these variables are .67, .72, and .73, respectively. We discuss and interpret these results in Online Appendix B.

Table 1 Racial resentment and relative discrimination interactively predict turnout

	(2012)	(2012)	(2016)	(2016)	(2020)	(2020)
Interaction term	3.457*	2.817*	3.434*	3.322*	5.020*	4.627*
	(0.831)	(0.867)	(1.063)	(1.081)	(1.143)	(1.162)
Rel. discrimination	-5.172*	-5.334*	-1.637*	-1.727*	-2.608*	-2.532*
	(0.310)	(0.326)	(0.448)	(0.458)	(0.519)	(0.529)
Racial resentment	-3.786*	-3.322*	-2.447*	-2.447*	-3.637*	-3.428*
	(0.623)	(0.653)	(0.786)	(0.804)	(0.828)	(0.845)
Age	-0.006	0.001	0.025*	0.022*	0.031*	0.028*
	(0.003)	(0.004)	(0.005)	(0.005)	(0.004)	(0.004)
Income	-0.037	-0.079	0.122*	0.111*	0.062*	0.057*
	(0.043)	(0.046)	(0.040)	(0.041)	(0.010)	(0.011)
Education	-0.111*	-0.039	0.390*	0.363*	0.214*	0.206*
	(0.049)	(0.052)	(0.073)	(0.075)	(0.036)	(0.037)
Female	0.111	0.250*	0.086	0.048	-0.020	-0.044
	(0.097)	(0.103)	(0.138)	(0.140)	(0.128)	(0.130)
Resid. mobil	0.004	-0.004	0.077*	0.069*	-0.276*	-0.278*
	(0.041)	(0.043)	(0.030)	(0.030)	(0.086)	(0.087)
Homeowner	-0.358*	-0.299*	0.282	0.243	0.127	0.130
	(0.142)	(0.150)	(0.164)	(0.165)	(0.136)	(0.137)
PID strength		0.040		0.258*		0.301*
		(0.081)		(0.067)		(0.060)
Freq. Church Att		-0.455*		0.128*		0.148
		(0.034)		(0.037)		(0.097)
Union Mem. in HH		0.570*		0.171		0.066
		(0.132)		(0.217)		(0.195)
Contacted		-0.151		0.263		0.507*
		(0.109)		(0.164)		(0.214)
Constant	4.368*	4.544*	-0.318	-0.695	1.593*	1.105*
	(0.355)	(0.386)	(0.444)	(0.464)	(0.484)	(0.500)
Observations	2,428	2,423	2,345	2,338	4,506	4,494
Log likelihood	-1,295.993	-1,181.877	-745.104	-723.144	-949.218	-925.991
Akaike Inf. Crit	2,611.986	2,391.754	1,510.209	1,474.289	1,918.435	1,879.983

Source: Logistic regressions, ANES 2012, 2016, and 2020

* $p < 0.05$

respondents turning out to vote. This pattern holds in each year considered. Moreover, the coefficients on the interaction between relative discrimination and racial resentment with and without the various controls remain essentially unchanged. In our models from 2016 and 2020, the coefficients for our controls are in their expected direction (age, income, education, residential mobility, and party ID strength). We note that in 2012 many coefficients point significantly in unexpected directions (church attendance), or are unexpectedly insignificant (income, party ID

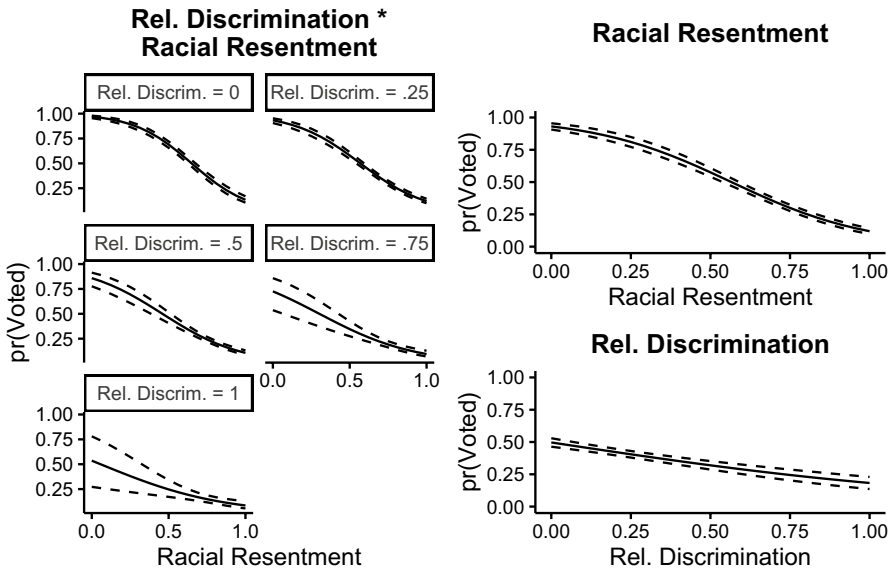


Fig. 1 Predicted probabilities of turnout, based on 2012 regression model in Table 1; column 2. Predicted probabilities with 95% confidence intervals. Source: 2012 ANES

strength). While other coefficients pointing in “wrong” directions are unstable from model to model (education), we note these findings with caution in 2012. Despite this, our 2016 and 2020 findings bolster our confidence in drawing conclusions from our regression analysis.

Given that logistic regression coefficients are difficult to substantively interpret, we produce plots (generated from the fully-controlled models included in Table 1—columns 2, 4, and 6) depicting the predicted probability of a respondent voting at a given level of racial resentment and relative discrimination (holding all other coefficients at their means). These figures also depict the probability of a respondent voting at different levels of racial resentment and specific levels of relative discrimination, thus modeling the substantive effect of the interaction term. We display the results from our 2012 analyses (Table 1, column 2) here.

We find substantively large negative effects of racial resentment, smaller negative effects of relative discrimination, and similarly small interaction effects (relative to the constituent terms) on the probability of turnout in 2012. All effects are statistically significant. As can be seen in Fig. 1, a standard deviation increase in racial resentment, *ceteris paribus*, corresponds to about a 25% lower predicted probability of voting. The effects of relative discrimination are less pronounced: a standard deviation increase in that variable corresponds to about a 10% decrease in the predicted probability of voting. The interaction effect is illustrated on the left panels of Fig. 1. At the lowest levels of perceived relative discrimination, racial resentment has a strong negative effect on the probability of turnout: a min–max change in racial resentment corresponds to about an 82% reduction in the probability of voting. On the other hand, at the highest level of relative discrimination, the negative effect of

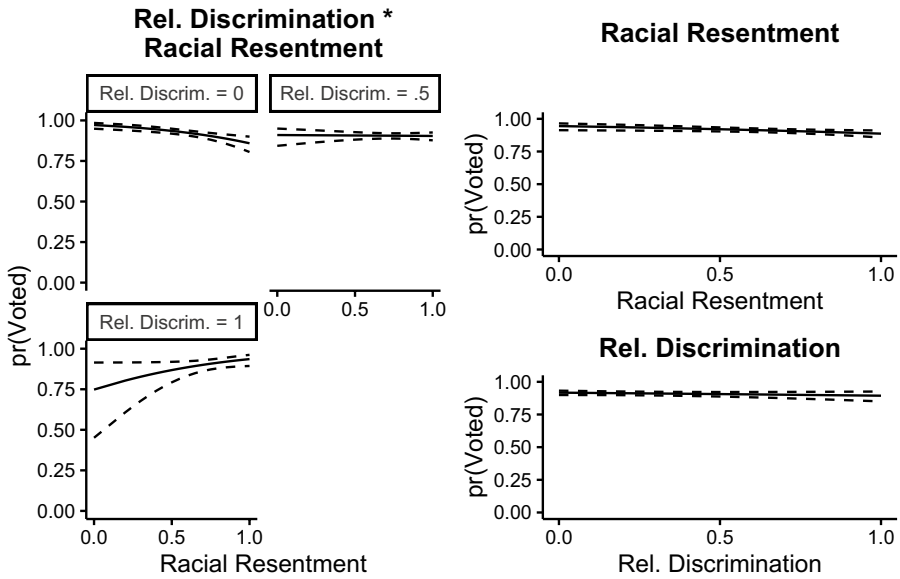


Fig. 2 Predicted probabilities of turnout, based on 2016 regression model in Table 1; column 4. Predicted probabilities with 95% confidence intervals. Source: 2016 ANES

racial resentment on the probability of voting is about halved: a min–max change in racial resentment reduces the probability of voting by 46%. Thus, while the coefficient on the interaction between racial resentment and relative discrimination is positive, the interactive effect is not substantively positive, per se. Instead, the interactive effect of racial resentment and relative discrimination moderates the negative independent effects of each constituent term. These analyses present evidence that, in 2012, whites high in relative discrimination and racial resentment were slightly less active than the least prejudiced whites, but were far more active than whites who espoused either racial resentment or relative deprivation.

The interaction effects on turnout are far stronger in 2016 compared to the effects on the constituent terms, as can be seen in Fig. 2. Increases in racial resentment and relative discrimination have statistically significant but substantively small effects on the probability of turnout. At the lowest levels of relative discrimination, a min–max change in racial resentment reduces the probability of turnout by about 15%. At the highest levels of relative discrimination, a min–max change in racial resentment increases the probability of turnout by a little above 20%, about two-thirds of a standard deviation. Our findings are similar in 2020, as can be seen in Fig. 3. In 2020, a min–max change in racial resentment at the lowest levels of relative discrimination reduces the probability of turnout by about 10%. At the highest levels of relative discrimination, a min–max change in racial resentment increases the probability of turnout by about 20%, about two-thirds of a standard deviation. Thus, in both 2016 and 2020, the interactive effects of our primary independent variables *overcome* the negative effects of their constituent terms. Whites who perceived both status threat and racial resentment were demobilized, yet in 2016 and

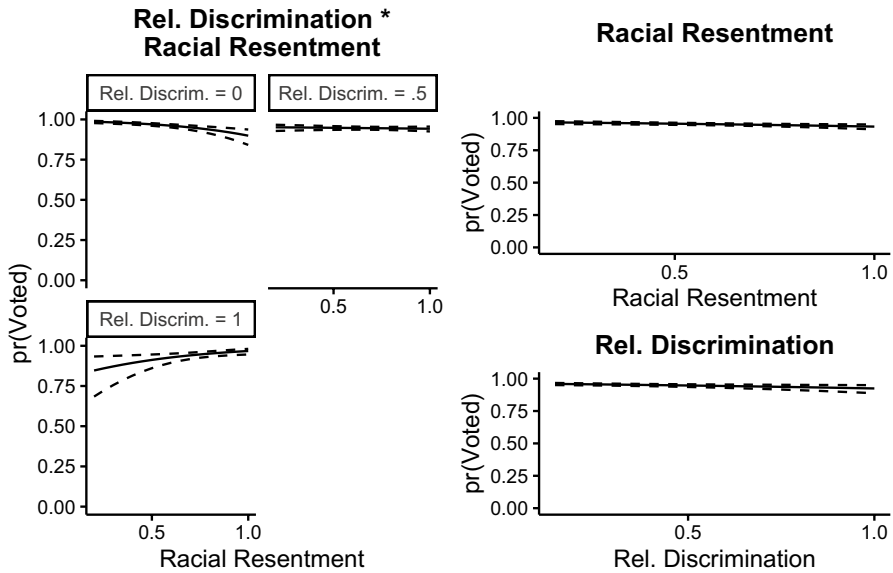


Fig. 3 Predicted probabilities of turnout, based on 2020 regression model in Table 1; column 6. Predicted probabilities with 95% confidence intervals. Source: 2020 ANES

2020 they were mobilized. We discuss the increased predictive power over time of our independent variables below.

To further test the interactive effects of racial resentment and relative discrimination on political behavior, we include political participation and civic engagement indices. In each model where we predict non-voting political participation, racial resentment and relative discrimination are independently and negatively correlated with participation, but positively correlated when interacted (see Table 2). In nearly all models, important coefficients like age, income, and education are significant and display the expected signs. In 2012, at the lowest level of perceived relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted non-voting acts from almost 1.5 (above average) to 0.8 (below average). In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted non-voting acts from less than 1 (below average) to about 1.4 (above average). In 2016, at the lowest level of perceived relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted non-voting acts from 3.4 (above average) to 2.2 (below average). This change is about a one standard deviation *decrease* in participation. In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted non-voting acts from 1.6 (below average) to about 3 (above average), which is more than a one standard deviation increase in participation. In 2020, at the lowest level of perceived relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted non-voting acts from 2.6 (well above average) to almost 0 (well below average). This change is almost a two standard deviation decrease in participation.

Table 2 Racial resentment and relative discrimination interactively predict non-voting political engagement

	(2012)	(2012)	(2016)	(2016)	(2020)	(2020)
Interaction term	0.246*	0.197*	0.376*	0.321*	0.549*	0.479*
	(0.051)	(0.049)	(0.048)	(0.047)	(0.054)	(0.053)
Rel. discrimination	-0.129*	-0.095*	-0.176*	-0.154*	-0.324*	-0.294*
	(0.018)	(0.018)	(0.019)	(0.018)	(0.022)	(0.022)
Racial resentment	-0.141*	-0.113*	-0.261*	-0.220*	-0.331*	-0.291*
	(0.041)	(0.038)	(0.036)	(0.036)	(0.041)	(0.040)
Age	0.002*	0.001*	0.002*	0.001*	0.002*	0.002*
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Income	0.008*	0.004	0.004*	0.003*	0.002*	0.002*
	(0.003)	(0.003)	(0.002)	(0.002)	(0.0005)	(0.0005)
Education	0.029*	0.021*	0.017*	0.016*	0.009*	0.009*
	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)
Female	-0.027*	-0.029*	-0.018*	-0.019*	-0.015*	-0.020*
	(0.007)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Resid. Mobil	-0.0004	-0.004	-0.001	-0.002	0.005	0.007
	(0.003)	(0.003)	(0.001)	(0.001)	(0.004)	(0.004)
Homeowner	0.009	0.001	-0.004	-0.007	0.002	0.003
	(0.010)	(0.009)	(0.008)	(0.008)	(0.006)	(0.006)
PID strength		0.051*		0.021*		0.041*
		(0.005)		(0.003)		(0.003)
Freq. Church Att		0.003		-0.002		0.001
		(0.002)		(0.002)		(0.004)
Union Mem. in HH		0.014		0.020*		0.008
		(0.009)		(0.008)		(0.008)
Contacted		0.107*		0.049*		0.048*
		(0.007)		(0.006)		(0.008)
Constant	0.029	-0.002	0.304*	0.261*	0.154*	0.068*
	(0.021)	(0.021)	(0.020)	(0.020)	(0.023)	(0.023)
Observations	3,119	3,108	2,345	2,338	4,507	4,495
R2	0.093	0.195	0.098	0.145	0.093	0.140
Adjusted R2	0.090	0.192	0.095	0.140	0.091	0.138

Source: OLS regressions, ANES 2012, 2016, and 2020

*p < 0.05

In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted non-voting acts from about 0 (below average) to about 2 (above average). That is about a one standard deviation increase in participation (see Table B2). Thus, while the least prejudiced are still the most active non-voting participants, the interactive effects of prejudice and relative discrimination almost entirely “make up” for the demobilizing effects of the constituent terms.

In each model where we predict non-electoral civic engagement, racial resentment and relative discrimination are independently and negatively correlated with civic engagement. Yet, when interacted they are positively correlated (see Table 3). Some other coefficients on control variables are unexpectedly insignificant, but have the expected sign (i.e., income in 2016). In 2012, at the lowest level of relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted acts of civic engagement from almost 3 (above the 75th percentile) to about 1.5 (below average). In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted engagement from 1 act (the 25th percentile) to more than 2 (above average). In 2016, at the lowest level of perceived relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted civic engagement from 2 acts (above average) to 1 (below average). That is almost a one standard deviation decrease in participation. In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted civic acts from -1 acts to more than 2 (75th percentile). This change is more than a two standard deviation *increase* in participation. In 2020, at the lowest level of perceived relative discrimination, a min–max change in racial resentment corresponds with a decrease in predicted civic engagement from 6 acts (one standard deviation above the 75th percentile) to about 1 (the 25th percentile), which is about a two standard deviation decrease in civic engagement. In the same year, at the highest level of relative discrimination, a min–max change in racial resentment corresponds with an increase in predicted civic engagement from just above 1 act (the 25th percentile) to about 4 acts. This change is about a one standard deviation increase in participation. The least prejudiced whites are the most civically engaged; yet, whites high in racial resentment and status threats are nearly as active, dwarfing the engagement of those who only espouse prejudice *or* status threat.

While the difference between the overall participation of the most and least prejudiced whites is often modest, it is significant across three measures. Moreover, the *increase* in political participation amongst prejudiced whites is significant, even if their level of activism is only above average. Increased activism among prejudiced whites makes them a potentially important constituency and changes political elites' incentives. We discuss this below.

Discussion

Our findings support the expectations we detailed above. We find evidence that both racial resentment and relative discrimination are independently and negatively associated with voting, non-voting participation, and civic engagement among whites. Yet, when interacted, they are positively correlated with each form of participation. In 2012, the interaction between racial resentment and relative discrimination *moderated* the negative effects of each. In 2016 and 2020, that same coefficient was substantively positive: racial resentment was positively associated with the level of participation among whites who reported the highest perceptions of relative discrimination. In short, whites' racial attitudes condition their decisions to participate,

Table 3 Racial resentment and relative discrimination interactively predict civic engagement

	(2012)	(2012)	(2016)	(2016)	(2020)	(2020)
Interaction term	0.255*	0.244*	0.676*	0.639*	0.497*	0.511*
	(0.050)	(0.048)	(0.055)	(0.054)	(0.047)	(0.046)
Rel. discrimination	-0.137*	-0.121*	-0.152*	-0.163*	-0.317*	-0.333*
	(0.018)	(0.017)	(0.026)	(0.026)	(0.019)	(0.019)
Racial resentment	-0.163*	-0.165*	-0.512*	-0.491*	-0.313*	-0.335*
	(0.039)	(0.038)	(0.038)	(0.038)	(0.035)	(0.035)
Age	0.001*	-0.0001	0.001*	0.001*	0.0003	0.00001
	(0.0002)	(0.0002)	(0.0003)	(0.0003)	(0.0002)	(0.0002)
Income	0.012*	0.009*	0.003	0.002	0.001*	0.001*
	(0.003)	(0.003)	(0.002)	(0.002)	(0.0004)	(0.0004)
Education	0.043*	0.034*	0.048*	0.044*	0.017*	0.016*
	(0.003)	(0.003)	(0.004)	(0.004)	(0.001)	(0.001)
Female	0.004	-0.003	0.008	0.003	0.001	-0.003
	(0.007)	(0.006)	(0.008)	(0.008)	(0.005)	(0.005)
Resid. Mobil	0.002	-0.001	0.0003	-0.002	0.001	0.003
	(0.003)	(0.003)	(0.002)	(0.002)	(0.004)	(0.004)
Homeowner	0.0001	-0.010	0.014	0.009	-0.001	-0.002
	(0.009)	(0.009)	(0.010)	(0.010)	(0.005)	(0.005)
PID strength		0.022*		0.0004		0.009*
		(0.005)		(0.004)		(0.002)
Freq. Church Att		0.019*		0.015*		0.042*
		(0.002)		(0.002)		(0.003)
Union Mem. in HH		0.021*		0.032*		0.013
		(0.008)		(0.011)		(0.007)
Contacted		0.099*		0.066*		0.041*
		(0.007)		(0.009)		(0.007)
Constant	0.070*	0.072*	0.082*	0.080*	0.307*	0.289*
	(0.021)	(0.020)	(0.026)	(0.026)	(0.020)	(0.020)
Observations	3,107	3,096	2,928	2,913	4,507	4,495
R2	0.125	0.217	0.162	0.196	0.137	0.174
Adjusted R2	0.122	0.214	0.159	0.193	0.135	0.172

Source: OLS regressions, ANES 2012, 2016, and 2020

* $p < 0.05$

not just vote choice. However, the relationship between prejudice and participation varies over time.

To account for the possibility that our results are driven by another attitudinal measure, rural consciousness (Cramer, 2016), which is often observationally equivalent to our primary independent variables (Nelsen & Petsko, 2021), we replicate our 2020 analyses including rural consciousness as an additional control. We include the results of these models in Appendix C. Controlling for rural

consciousness does not change the predictive power of our independent variables and the coefficient estimates on rural consciousness are either insignificant or pointing in unexpected directions. These results suggest that our findings are not merely capturing rural consciousness.

While we did not predict the differences between our findings in 2012, 2016, and 2020, they echo findings by other scholars, allowing us to speculate on how the mobilizing power of white racial attitudes changes over time. In 2012, whites who espoused *both* prejudice *and* status threat were more likely to participate than those who were either prejudiced or perceived anti-white discrimination. Yet they were still significantly less likely to participate than the least prejudiced whites. In elections contested by Trump, however, whites who reported high racial resentment and relative discrimination were as active as the least prejudiced whites—sometimes participating more. Other scholars (Jardina, 2020) have noted that Trump successfully intertwined disparate strands of white identity politics in 2016. We corroborate that finding here, and present evidence that he replicated this feat in 2020. This finding echoes arguments that the racialization of American politics accelerated during the 2016 election (Jardina, 2019; Sides, et al., 2019) and differentiates us from scholars who frame 2012 as the critical year (Tesler, 2016).

Our examination of three elections thus enables us to show that the mobilizing power of white racial attitudes may be contingent on explicit, racialized campaign rhetoric. Trump's success in mobilizing prejudiced voters will remain synonymous with the 2016 election, but we suggest it should also be connected to the 2018 election, when Trump's 2016 coalition stayed home. We discuss this below. Finally, our finding that high-prejudice, high-status threat whites are, at most, only moderately more likely to vote than the least-prejudiced whites points to the limitations of relying on these voters. Variables like racial resentment are negatively correlated with predictors of political participation like education and trust in government. This could help explain why, even in a highly racialized campaign, the most prejudiced whites are not drastically more engaged than the least prejudiced. We discuss the implications of this possibility below.

While our findings clarify the relationship between racial attitudes and participation among whites, there are limitations. First, these results are produced from an examination of cross-sectional data. While cross-sectional data is still the norm in studies on white racial attitudes and voting behavior (e.g., Sides et al., 2019), some compelling studies have begun to move to more acute methods, like analysis of panel data (Mutz, 2018). Moreover, we cannot determine whether this relationship between racial resentment, relative discrimination, and white political participation crystallized around 2012 or has existed longer. Unfortunately, as noted above, the lack of ANES items assessing perceived discrimination precludes us from making such a determination. Given the paucity of research examining how whites' attitudes about race shape their political participation, however, we argue that our findings advance understanding of white identity politics.

Conclusion

We present three novel contributions to the study of white political behavior. First, we clarify the relationship between whites' racial attitudes and political participation: whites' racial prejudice and perceptions of inequality inform their decision to turn out, not just their vote choice. Specifically, these results present evidence that whites are not mobilized by their prejudice or senses of status loss alone. Indeed, those attitudes, independently, can reduce whites' probabilities of participating in politics. In 2012, the interaction between racial resentment and relative discrimination attenuated these negative effects. In elections contested by Trump, the interaction between these constructs *overcame* the negative effects of racial resentment and relative discrimination. Scholars have previously pointed to whites' racial attitudes being understudied as predictors of political engagement (Krupnikov & Piston, 2015). Our findings advance research into white political behavior by indicating how multiple measures of white racial attitudes shape their political participation. Moreover, we demonstrate a similar relationship across different measures of participation: voting, non-voting electoral participation, and civic engagement. However, we speculate that the mobilizing power of white racial attitudes may be contingent upon elites making racial campaign appeals sufficiently explicit.

Second, our findings clarify the relationship between various constructs in the research program on white racial attitudes. More precisely, we find evidence that white out-group animus and in-group solidarity are conceptually related but analytically distinct influences on American politics (Jardina, 2020). Berry et al. (2019) studies how racial resentment (negatively) and white linked fate (positively) predict whites' political participation. Our work shows that different white racial attitudes have independent influence on whites' politics. More specifically, these results present evidence that different white racial attitudes may have interactive, rather than additive, effects.

Finally, these results clarify that the relationship between whites' prejudice and politics is *conditional*. Some whites may believe that Blacks contravene social norms around work and individualism without believing that whites, collectively, suffer due to that perceived deviance. When whites perceive status threat *and* espouse racial resentment, they may perceive greater urgency in seeking political redress. Moreover, extensive research has framed racial resentment as the dominant frame conservative white elites use to talk about race in America (Kinder & Sanders, 1996; Mendelberg, 2001; Sides et al., 2019). Whites who perceive substantial relative discrimination but not racial resentment may not have internalized this rhetoric, and thus may be unable to connect their misperceptions of marginality to political solutions. Whites who have internalized beliefs about Blacks' responsibility for racial inequality *and* believe that whites are deprived relative to other racial groups necessarily perceive inequality, and, due to their internalization of dominant racial ideologies, are more likely to seek political redress.

Together, these contributions have implications for the political meaning of whites' prejudice in contemporary politics. We offer two ways to interpret the

implications of our findings. The first interpretation emphasizes the limits of white identity politics. Research largely written after the 2016 election implies (sometimes forecasts) an increasingly self-conscious white identity politics channeled through the Republican Party (Craig & Richeson, 2014; Jardina, 2019; Weller & Junn, 2018; Wong, 2018). Such researchers primarily study vote choice or partisanship. By examining political participation, we qualify their findings: while various facets of white identity strongly predict Republican affiliation, they have a more complicated relationship with participation. Inconsistent participation among high-prejudice, high-status threat whites could constrain their political influence. We note that the “Obama coalition” heralded in 2008 proved more fragile and situational than many observers expected. Likewise, the “Trump coalition” could erode.

The second interpretation is less sanguine. As we showed above, whites may need to express prejudice *and* perceptions of anti-white discrimination to be politically mobilized to defend their group. Racial resentment and perceptions of white group status are part of broader racial ideologies that Americans develop in conversation with political and intellectual elites. As American parties increasingly polarize around competing perceptions of the American racial hierarchy, the Republican Party will stand to benefit from further reinforcing (and appealing to) these ideologies, ensuring that these voters remain mobilized. If that happens, the whites highest in prejudice and perceived status threat could move from unreliable voters to an exceptionally powerful constituency in American politics.

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