

# Identity as Dependent Variable: How Americans Shift Their Identities to Align With Their Politics

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Political science generally treats identities such as ethnicity, religion, and sexuality as “unmoved movers” in the chain of causality. I hypothesize that the emergence of partisanship and ideology as social identities in the U.S., combined with the increasing demographic distinctiveness of the nation’s two political coalitions, is leading some Americans to engage in a self-categorization and depersonalization process in which they shift their identities toward the demographic prototypes of their political groups. Analyses of a representative panel dataset that tracks identities and political affiliations over a four-year span confirm that small but significant shares of Americans engage in identity switching regarding ethnicity, religion, sexual orientation, and class that is predicted by partisanship and ideology in their pasts, bringing their identities into alignment with their politics. These findings enrich and complicate our understanding of the relationship between identity and political behavior and suggest caution in treating identities as unchanging political phenomena.

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In recent years, identity has again emerged as a key explanatory variable in both academic and popular accounts of U.S. politics. The shift was reinforced by the 2016 presidential election, which revealed an American electorate deeply divided along the lines of identity dimensions like race, Latino origin, religion, and sexual orientation. Many political scientists now point to group identity as a key independent variable predicting political behavior (e.g. Abdelal et al 2006; Achen and Bartels 2016; Huddy 2003, 2018; Kinder and Kalmoe 2017), returning full circle to the focus placed on identity by some of the earliest academic accounts of voting and attitudes (Berelson, Lazarsfeld and McPhee 1954; Campbell et al 1960).

In political science, it is often implicitly assumed that identities are stable and therefore can be confidently considered to be antecedent to political attitudes and behavior. But recent empirical work has challenged the premise that identities are exogenous to politics. These findings have been most pronounced in the religion and politics literature, where analyses of U.S. panel data demonstrate short-term change in Americans' religious identification and participation in response to their partisanship and ideology (Campbell et al 2018; Djupe, Neiheisel, and Sokhey 2018; Hout and Fischer 2014; Margolis 2018a, 2018b; Patrikios 2008; Putnam and Campbell 2010). Other research has found that politically salient factors in Americans' backgrounds and upbringing are long-term predictors of their present identities with regard to race (Davenport 2016) and sexual orientation (Egan 2012), indicating that these identities too can be endogenous to politics.

Previous work has looked at identities one category at a time, and thus the mechanisms of politicized identity change provided by researchers have largely been grounded in aspects of particular identity categories being studied. Here I use social identity theory, offered by some scholars to explain the effect of politics on religion (Campbell et al 2018; Patrikios 2008), to hypothesize more broadly about how politics can lead Americans to shift their identities across multiple categories we typically think of fixed. I begin with the well-documented fact that for more and more Americans, politics has become key to the self-concept, leading "Democrat" and Republican" as well as "liberal" and "conservative" to become identities in themselves that have meanings far beyond shared policy

preferences (Devine 2014; Huddy, Mason and Aarøe 2015; Iyengar, Sood and Lelkes 2012; Iyengar and Westwood 2015; Klofstad, McDermott and Hatemi 2012; Malka and Lelkes 2010; Mason 2018a; 2018b). Social identity theory tells us that highly salient identities such as these can provide a definition of the self in terms of the defining characteristics of the identity group. Through a process called self-categorization, these characteristics are woven together into prototypes which become stylized representations of the kinds of persons who belong to the identity group (Turner 1985; Turner et al 1987). When the identity becomes salient, identifiers can engage in a depersonalization process in which beliefs and actions converge toward those of prototypical group members (Hogg et al 1995).

A concurrent development in U.S. politics is supplying a readily accessible source of content for these prototypes. In a “social sorting” process as documented by Lilliana Mason (2016; 2018b), the nation’s two political coalitions are now quite distinct demographic groups with regard to characteristics like race, ethnicity, religion, and sexuality. Liberal Democrats are increasingly more likely to be people of color, sexual minorities, and non-religious; conservative Republicans are likely to be non-Hispanic white, heterosexual, and religious. Social sorting is evident at both mass and elite levels, making it easy for Americans to call these prototypes to mind (Ahler and Sood 2018).

These two developments in U.S. politics—political groups becoming the source of increasingly salient identities, and political groups becoming increasingly demographically distinct—provide ideal conditions for a self-categorization and depersonalization process. I hypothesize these conditions are leading some Americans to adjust their demographic identities to better align with partisan and ideological prototypes.

I explore this hypothesis using a nationally representative panel survey dataset in which questions about a range of identities were asked of empaneled respondents three times over a four-year period. I find that during this span substantial numbers of Americans shifted in and out of identities associated with ethnicity, religion, sexual orientation, and class. Furthermore, I show that small but significant shares of liberal Democrats and conservative Republicans shifted these identities in ways that conform with political

group prototypes. Conservative Republicans were more likely than liberal Democrats to shift into identification as born-again Christian, Protestant, and national origins associated with being non-Hispanic white. Liberal Democrats were more likely than conservative Republicans to shift into identification as lesbian, gay or bisexual, having no religion, and being of Latino origin. Each of these shifts brought liberal Democrats' and conservative Republicans' identities into better alignment with the identity groups that make up the two U.S. political coalitions. As expected, these shifts were more pronounced among liberal Democrats and conservative Republicans for whom party identification and ideology did not change over the four-year period. Additional analyses indicate that politics plays a particularly important role in identification with two groups for which identity is typically acquired later in life, rather than transmitted across generations—lesbians, gays and bisexuals, and those identifying as having no religious affiliation—in that the impact of politics on identity is large for these groups relative to their current prevalence in the population.

These findings demonstrate that in the United States the range of identity categories subject to override by partisanship and ideology is much wider than shown in previous research, where the most compelling findings thus far have focused on the relationship between politics and religious identity. The scope of political identity shifting includes ethnicity, sexuality, and class, which until now have not been shown to be subject to short-term change that aligns with individual political characteristics. The analysis suggests that circumstances are favorable for a similar process to take place in other contexts when the two conditions currently the case in the U.S. hold elsewhere: political groups take on the qualities of identities and political groups become demographically distinct. Because the mechanism demonstrated here applies to multiple identity categories, it serves as a complement to other valuable theories of politicized identity change documented by researchers in fields such as religion and politics, sexuality and politics, and race, ethnicity and politics. All told, these discoveries have important consequences for our understanding of the formation and maintenance of identity groups and how political scientists model the relationship between identity and other political phenomena.

## Identities and Identity Change

A straightforward definition of *identity* is a social category into which people are placed based upon one or more individual attributes. Attributes are mapped to identities according to membership rules that say which attributes are necessary for membership in the identity. Many important attributes are impossible to change (such as place of birth, ancestry of parents, and sexual attraction) or very hard to change (such as sex, skin tone, and other physical attributes), and for most purposes can be considered fixed. Another set of attributes can change but typically do so slowly, such as language, religion and socio-economic status; they are “sticky” (Chandra 2012). Fixed and sticky attributes are necessary for membership in many of the most highly salient identity categories in American politics, such as race, ethnicity, national origin, sexual orientation, religion, and class, which leads to the implicit assumptions that these identities are unchanging and can be considered unmoved movers in models of political behavior.

But for two reasons, caution is called for in assuming identities like these do not change over time. The first concern arises from the distinction between objective group membership and subjective group identification. In contrast to mere group members, group identifiers have a “subjective, or internalized sense of belonging to the group” (Huddy 2003: 513-514). While group membership is in many cases straightforward, group identification by contrast can be up to a fair amount of individual discretion. Because for some the strength of subjective identification can shift over time, claimed identities can change as well. The second concern is that many people have attributes that place them near the boundaries that demarcate one identity from another, and these boundaries tend to be fuzzy (Chandra 2006, 2012). Examples include those born to parents of different races or religions; those attracted to both sexes; or those who are subject to a change in economic status. The ethnic, religious, sexuality and class identities claimed at any given time by those located near these boundaries reflects the extent they identify more strongly with one side of the divide than the other. This too can change over time as people have new

experiences and encounter new contexts. Taken together, these two observations suggest that shifts in identities typically considered fixed may not be an entirely rare phenomenon.

### **When Politics Makes Identity a Dependent Variable**

Many over-time identity shifts have political causes, as demonstrated by a wealth of evidence from the United States and around the world. For example, scholars of comparative politics have shown that individuals' subjective ethnic identities and their salience can be shaped by political institutions such as electoral rules (Posner 2005) and government census categories (Lieberman and Singh 2017; Nobles 2000). Analyses of U.S. survey data confirm that many contemporary American identities are endogenous to politics. Work on this topic has been particularly advanced in the the religion and politics field, where scholars have employed surveys—and particularly, panel designs—to document how politics affects religious identity and participation. This research finds Democrats and liberals switching into identification as non-religious (Campbell et al 2018; Hout and Fischer 2002, 2014; Margolis 2018a, 2018b; Putnam and Campbell 2010) and claiming an affirmatively secular identity (Campbell et al 2018). It also shows how political disagreement is causing Americans to attend religious services less frequently (Margolis 2018a, 2018b; Patrikios 2008) and leave their houses of worship (Djupe, Neiheisel, and Sokhey 2018). Other survey research has documented how racial and sexual identities are affected by politics, although here the focus thus far has been on longer-term processes. Davenport (2016) finds that when Americans are of mixed-raced parentage, their own racial and ethnic identification is strongly shaped by politically salient, causally prior characteristics like gender, religion of upbringing, and parents' socio-economic status. Egan (2012) shows that the likelihood of coming out as LGB is significantly affected by characteristics of one's upbringing that are causes of political attitudes, meaning that a substantial proportion of LGBs' distinctively liberal political attitudes can be attributed to selection effects.

The mechanisms offered for how religious, racial and sexual identities are shaped by politics tend to be grounded in particular aspects of the identity being studied. Mechanisms developed to explain why politics affects religious identity include the timing of reli-

gious and political socialization processes (Margolis 2018a, 2018b) and disagreements with religion's moral traditionalism and the dogmatic politics of the Christian Right (Djupe, Neiheisel, and Sokhey 2018; Hout and Fischer 2002; 2014). Davenport's explanation for the identity choices made by biracial Americans rests on how race and ethnicity is reinforced and challenged depending on one's gender, socio-economic status, and other characteristics. Egan's theory of selection effects and LGB identity is based on the strong correlation between liberal political views and acceptance of sexual minorities in families of origin.

The mechanisms provided in previous research that are most similar to the explanation developed here focus on partisanship and ideology as social identities. Patrikios (2008) argues that self-categorization processes associated with partisanship and ideology result in group norms that affect evangelicals' levels of church attendance. Campbell et al (2018) show that the dissonance between group norms and identity for Democrats becomes pronounced when the association between the GOP and religion is made salient in an experimental context, leading them to make immediate shifts toward non-religious identification. Here I extend these social-identity theory explanations to present the first comprehensive explanation for why shifts across multiple identity categories can be predicted by political and ideological orientations.

### **Partisanship and Ideology as Social Identities**

U.S. politics in our current era is characterized by historically high levels of partisan and ideological polarization among elites and masses (e.g. Abramowitz and Saunders 2008; Bafumi and Shapiro 2009; McCarty, Poole and Rosenthal 2016). Amid this wave of polarization, "Republican" and "Democrat," as well as "liberal" and "conservative," have become more than just bundles of policy preferences. They have also taken on the qualities of social identities, a hallmark characteristic of which is that in-group members make favorable comparisons between themselves and out-group members (Tajfel and Turner 1979). This is most markedly shown by the social distancing that liberals and conservatives exhibit toward one another. They rate one another negatively on survey questions; and they prefer to be friends with, date, marry, work and do business with, and be neigh-

bors with their own ideological group (Devine 2014; Huber and Malhotra 2017; Huddy, Mason and Aarøe 2015; Iyengar, Sood and Lelkes 2012; Iyengar and Westwood 2015; Klostad, McDermott and Hatemi 2012; Malka and Lelkes 2010; Mason 2018a; McConnell et al 2018).

One of the ways social identities become integrated into the self-concept is through self-categorization and depersonalization, a process by which group identifiers come to perceive themselves as similar to the prototypical identity group member and adopt beliefs and behaviors that conform to the prototype (Turner 1985; Turner et al 1987). This process can develop among members of groups both large and small. When group identity is salient, conforming to the prototype makes individuals better liked and more popular with other in-group members (Hogg and Terry 2000). Self-categorization and depersonalization thus play an important role in generating and sustaining in-group cohesion and loyalty.

Rich material is provided for the construction of political prototypes by the fact that the demographics of Democrats and Republicans, and liberals and conservatives, now differ substantially on these identity categories, a process Lilliana Mason calls “social sorting.” Mason’s *Uncivil Agreement* documents how party identifiers have become increasingly distinct with regard to characteristics such as race and church attendance, as well as the extent to which they feel “close to” people of different demographic groups in their party coalitions (2018b). Social sorting is reflected in Table 1a, where data from the 2016 American National Election Studies (ANES) show that non-Hispanic whites and born-again Christians make up substantially greater shares of Republicans and conservatives than Democrats and liberals. By contrast, people of color (in particular, blacks and Latinos), LGBs, Jews, and those who claim no religious affiliation make up greater shares of Democrats and liberals than Republicans and conservatives.

The final column of Table 1a reports each identity’s dissimilarity index score. A widely used measure of residential segregation, here the absolute value of the index measures the proportion of Americans claiming the identity who would have to switch political affiliations in order for equal numbers of the identity group’s members to call themselves liberal



**Table 1: Sources of Partisan and Ideological Prototypes in U.S. Politics****a. Demographic characteristics of  
partisan and ideological groups, 2016 ANES**

Identity	Dem- ocrats	Repub- licans	liberals	conserv- atives	liberal Dems	conserv Reps	dissimilarity index
White, not Hispanic	57.0%	84.3%	66.2%	75.6%	65.3%	85.4%	0.28
Black, not Hispanic	19.0%	1.8%	12.3%	7.1%	13.4%	1.2%	-0.49
Asian/Pacific, not Hispanic	3.2%	2.6%	3.6%	2.8%	3.6%	2.5%	-0.10
Native, Not Hispanic	0.7%	0.5%	0.7%	0.4%	0.6%	0.6%	0.02
Other, not Hispanic	4.2%	3.4%	3.7%	3.8%	3.4%	3.6%	0.02
Hispanic	15.3%	6.7%	13.0%	9.7%	13.1%	6.3%	-0.20
Protestant	18.6%	37.3%	19.4%	34.4%	20.2%	41.7%	0.25
Catholic	22.2%	22.4%	19.4%	23.1%	20.5%	22.7%	0.03
Jewish	2.7%	1.2%	3.9%	0.8%	4.4%	0.9%	-0.36
Born-again Christian	27.7%	44.6%	19.9%	45.4%	18.3%	47.1%	0.32
Agnostic, atheist, none	28.3%	13.9%	31.9%	15.2%	33.0%	13.0%	-0.29
Lesbian, gay, bisexual	8.5%	2.1%	11.9%	2.3%	11.4%	1.4%	-0.45
Lower class	9.6%	5.4%	6.3%	7.8%	5.6%	4.5%	-0.06
Working class	38.7%	38.2%	34.8%	38.6%	31.2%	35.8%	0.05
Middle class	47.4%	52.4%	52.7%	50.4%	56.5%	55.4%	-0.01
Upper class	4.3%	3.9%	6.2%	3.2%	6.7%	4.3%	-0.12

**b. Demographic characteristics of U.S. House of Representatives, 2015**

Identity	Democrats	Republicans	Progressive Caucus	Freedom Caucus
White, not Hispanic	59.3%	94.4%	46.2%	97.2%
Black, not Hispanic	22.8%	0.8%	36.9%	0.0%
Asian/Pacific, not Hispanic	5.3%	0.0%	6.2%	0.0%
Native American, not Hispanic	0.0%	0.8%	0.0%	0.0%
Hispanic	12.7%	4.0%	10.8%	2.8%
Protestant	45.0%	65.9%	46.2%	63.9%
Catholic	36.5%	29.4%	30.8%	25.0%
Jewish	9.5%	0.4%	12.3%	0.0%
Religion: did not state	5.3%	0.0%	6.2%	0.0%
Openly lesbian, gay or bisexual	3.7%	0.0%	6.2%	0.0%
Ever had working-class job*	11.6%	4.4%		
From working-class background*	29.4%	13.8%		
Net worth <\$100,000*	16.1%	13.5%		

Sources: Membership lists: DeSilver 2015, Congressional Progressive Caucus 2018;  
Biographical information: CQ 2016. Class indicators for members of Congress: Carnes 2016.

\*Data from 2007.

Democrats and conservative Republicans.<sup>1</sup> (I have signed the index negative for identities overrepresented among liberal Democrats and positive for those overrepresented among conservative Republicans.) Elected officials further substantiate these prototypes, as shown by the demographic differences between the parties' Members of Congress displayed in Table 1b. Differences are most pronounced between the Freedom Caucus and Progressive Caucus, composed of respectively the most conservative Republicans and most liberal Democrats in the House.

Thus at both elite and mass levels, liberal Democrats and conservative Republicans present Americans with highly divergent prototypes along the lines of race, ethnicity, religion, and sexual orientation. Recent research indicates that Americans are not only aware of these demographic differences between the two political camps, but that they exaggerate them in their minds (Ahler and Sood 2018). Here I investigate the claim that as partisanship and ideology have become important social identities in U.S. politics, the process of self-categorization leads some people to adopt identities that conform with these prototypes and shed identities that do not.

## Data

Data come from the General Social Survey (GSS), the biennial study of Americans' attitudes and behavior conducted by the National Opinion Research Center (Smith et al 2017). The GSS uses cluster-based sampling to obtain nationally representative samples of non-institutionalized adults in the contiguous United States. In the 2006, 2008 and 2010 GSS, respondents were empaneled to be reinterviewed twice over the next four years, creating three three-wave panels. Interviews of the first panel took place in 2006, 2008, and 2010; the second panel in 2008, 2010, and 2012, and the third panel in 2010, 2012, and 2014. In the analyses in this paper, I pool the three panels by wave of interview. Most interviews were conducted face-to-face; successful recontact required that a greater proportion of inter-

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<sup>1</sup>For any identity group  $j$ , the magnitude of the index is calculated as  $\frac{1}{2} \left( \left| \frac{LD \cap J}{J} - \frac{LD \cap \sim J}{\sim J} \right| + \left| \frac{CR \cap J}{J} - \frac{CR \cap \sim J}{\sim J} \right| \right)$ , where  $J$  and  $\sim J$  are the shares of the population who do and do not identify as a  $j$ , and  $LD$  and  $CR$  are the shares of the population identifying respectively as liberal Democrats and conservative Republicans.

views in later waves were done by phone. Four-year recontact rates averaged 60.5 percent over the two waves across the three panels.

At the center of the analyses in this paper are data derived from questions asked in multiple waves about respondents' identities with regard to race, ethnicity and national origin, religion, class, sexuality, and partisanship and ideology. Each identity category was scored dichotomously, taking on the value one if the identity is claimed by the GSS respondent and zero if not. The GSS assessed some identities on a dichotomous basis with simple yes-or-no questions, while for other identities respondents were given several responses from which to choose or were asked to provide an open-ended response. As discussed below, all responses were recoded into dichotomous variables. For every identity, "don't know" and "refuse to answer" responses were coded as zeroes rather than as missing data. Panelists who the GSS failed to contact for a reinterview in later waves of the panel were dropped from analyses; all analyses incorporate the panel non-response and post-stratification survey weights supplied by the GSS. Because question wording can have substantial impact on the measurement of identity, here I briefly discuss the survey items and recoding choices made for each identity category.

***Race.*** Respondents were asked "What is your race? Indicate one or more races that you consider yourself to be." Respondents were then presented with a card featuring a list of choices. Respondents' first reported race was coded into three dichotomous variables: "white," "black," or "Asian/Pacific."<sup>2</sup> This last category was created by collapsing several Asian and Pacific Islander identity categories from which respondents could choose, including "Asian Indian," "Chinese," "Japanese," and "Native Hawaiian."<sup>3</sup>

***Hispanic/Latino origin.*** A separate GSS question asked, "Are you Spanish, Hispanic,

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<sup>2</sup>"American Indian or Alaska Native" was also offered as a response by the GSS, but too few people chose this category ( $N = 37$  across valid cases in the first wave) for it to be included in analyses.

<sup>3</sup>In response to the race question, some participants volunteered that they were "Hispanic," despite the fact that this category was not presented as a choice to respondents. (The GSS measured Hispanic/Latino identity separately as discussed below.) Because volunteered responses are likely to be subject to a substantial degree of fluctuation from wave to wave of a panel survey, these responses were coded as zero for all racial categories.

or Latino [Latina if female]?" "Yes" and "no" responses were scored on a dichotomous basis. This question was asked of all respondents; Latinos could therefore be of any race.

**National origin.** All GSS respondents were asked the open-ended question "From what countries or part of the world did your ancestors come?" Multiple responses were permitted; here I analyze the country or place named first by respondents. The 11 most frequent responses to this item were used to create dichotomous variables, including two frequent responses—"American Indian" as well as "American only"—that do not refer to places outside the U.S.

**Sexuality.** The computer-administered self-interview (CASI) part of the GSS included the question "Which of the following best describes you?" followed by the choices (presented on the computer screen) of "gay, lesbian, or homosexual," "bisexual," and "heterosexual or straight." The first two responses were collapsed to create the dichotomous variable "lesbian, gay or bisexual." Administration of this question did not begin until 2008, reducing the sample size for analyses of LGB identity.<sup>4</sup>

**Religion.** GSS respondents were asked "What is your religious preference? Is it Protestant, Catholic, Jewish, some other religion, or no religion?" Responses of "Protestant," "Catholic," "Jewish," and "no religion" were each coded as dichotomous variables. A separate question later in the survey asked every respondent, "Would you say you have been 'born again' or have had a 'born again' experience - that is, a turning point in your life when you committed to Christ?" "Yes" and "no" responses to this question were scored as a separate dichotomous variable; born-again Christians could thus be of any religion.

**Class.** The GSS assessed self-described class by asking "If you were asked to use one of four names for your social class, which would you say you belong in: the lower class, the working class, the middle class, or the upper class?" Each of these responses was coded as a dichotomous variable.

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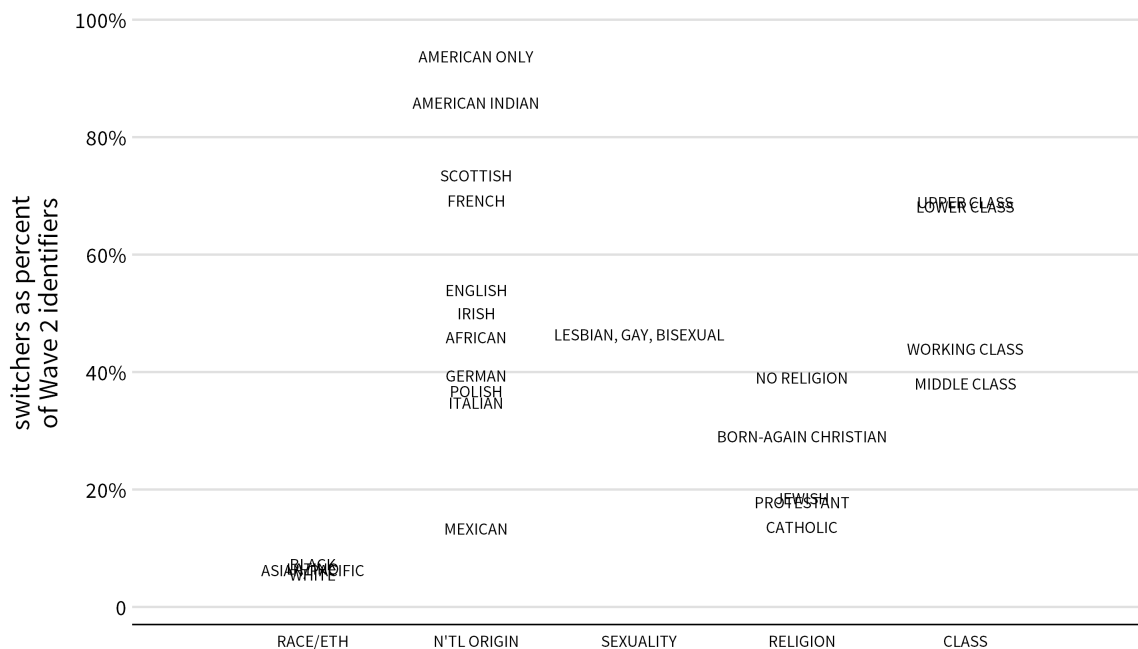
<sup>4</sup>The GSS did not include a question about transgender identity.

## The Instability of Identities Over Time

I first analyze the extent to which each identity is claimed by American adults and the extent to which identities are stable over time by focusing on identity claimed in Wave 2 of the three-wave panel. The multi-wave panel design permits expressing the total number of respondents claiming the identity in Wave 2 as those who (a) claimed the identity in both Waves 1 and 3; (b) claimed the identity in Wave 1 but not Wave 3; (c) claimed the identity in Wave 3 but not Wave 1; or (d) claimed the identity in neither Wave 1 nor Wave 3. Thus for each identity the quantity  $\frac{b+c+d}{a+b+c+d}$  is particularly revealing: ranging from zero to 100 percent, it is an estimate of the share of those claiming the identity at any given time who either did not hold the identity two years beforehand, abandoned the identity two years later, or both.

**Figure 1: Identity Switching Over Four Years  
in the Three-Wave GSS Panel Survey**

**Identity Switchers as Percent of Wave 2 Identifiers**



A graphical display of each identity's switch rate is found in Figure 1; the statistics plotted on this graph are shown in tabular form in this paper's Supporting Information (SI page 1). Identities associated with race and Latino origin exhibit the highest rates of stability, with the share of identifiers who are switchers falling below ten percent for each. By contrast, national origin identities exhibit higher rates of switching, which is likely in part due to the fact that the GSS question about national origin is open-ended. The range of switch rates among national origins is substantial. "Mexican" stands out as the most stable national origin, with more than 85% of Wave 2 identifiers consistently providing this response across all three waves. By contrast, two remarkably unstable national origin categories are "American Indian" and "American Only;" nearly all Americans claiming these identities at any given time are new to them, will decline to claim them later, or both.

Generally, religious identification is less stable than race and Latino origin, but more stable than most national origins. Those switching into and out of religious identities make up less than 20% each of Protestants, Catholics and Jews.<sup>5</sup> By contrast, nearly four-in-ten people who do not identify with a religious denomination have switched in, or will switch out of, this category over a four-year period. The identity of born-again Christian also exhibits a fair degree of instability; 29% of those saying at any given time they have had a born-again Christian experience either did not say so two years prior or did not say so two years later, or both. Two identity categories exhibit much higher rates of instability: sexuality and economic class. Nearly half (47%) of those identifying as lesbian, gay or bisexual at any given time have recently switched into or will soon switch out of the identity, or both.<sup>6</sup> Finally, as a society where class is not a particularly salient characteristic or organizing identity, it is not surprising that class identification in the United States exhibits relatively high rates of instability.

In total, these data indicate that the identities survey respondents claim can be quite fluid over time, even with regard to identities requiring fixed or sticky individual attributes for membership. The stability of identity categories in the contemporary U.S. ranges

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<sup>5</sup>Roughly half of these switches were into and out of identifying as non-religious.

<sup>6</sup>Those identifying as bisexual in Wave 2 were significantly ( $p < .05$ ) more likely to switch in or out of LGB identity than those identifying as lesbian or gay in Wave 2.

roughly from race and ethnicity as most stable, religion and sexual identity as less stable, and economic class as the least stable, while the stability of national origins is highly variable across different groups.

### **Measurement concerns**

One concern that emerges in these analyses is the extent to which instability in identities is a function of response instability among those who select the identity rather than being attributable to the salience and centrality of the identity itself. The tendency to provide unstable responses may be correlated with the claiming of particular identities, confounding stability of the identity with the response stability of identifiers. To address this concern, I estimated a model (discussed in the SI, page 2) in which each respondent essentially served as her own control. This analysis generated estimates of switching rates that controlled for any time-invariant individual characteristics, including the extent to which individuals tend to provide inconsistent responses in surveys. Results show that the relative stabilities among identities estimated by these two approaches are broadly comparable.

A second concern to consider when analyzing any over-time change in panel surveys is the extent to which observed change is due to measurement error rather than change in true values over time. Unfortunately, the standard measurement model used to assess the reliability of measures in three-wave panel surveys (developed in Heise 1969 and Wiley and Wiley 1970) rests upon a crucial but untestable assumption about how true scores change that on its face is inappropriate for measurements of identity. This assumption is that the true scores change via a lag-1 (or Markovian) process, which is to say that after accounting for one lagged measure of the true score, no additional past values of the true score are meaningful predictors of the true score's present value. This "memoryless" process is inappropriate for modeling change that unfolds over long periods of time, such as shifts in identity. In the SI (pp. 3-11), I show that when the lag-1 assumption is violated, estimates of reliability via the Heise/Wiley-Wiley model are biased, and under conditions we would expect to be common in the measurement of identity this bias is in a negative direction. To help rule out the concern that measurement error explains these results, in

the next section I conduct a series of placebo tests with measures of variables whose true values are known to be unchanging over time as well as sensitivity analyses assessing how vulnerable results are to errors in measures of identity. These tests show that it is unlikely that the key findings in this paper are attributable to measurement error.

Another reasonable concern about the measures of identity used here is their validity: to what extent can changes in survey response be equated with changes in underlying identity? As is common in most surveys, in the GSS just one question is used to measure each identity. This is an inevitably blunt way to assess identity, which political psychologists recognize as a continuous, rather than discrete concept that has several different dimensions, including salience, sense of belonging, and valence (Huddy 2003). Recalling that subjective identification can shift over time, and that many people find themselves near the boundaries between identities, over-time shifts in the identity items analyzed here are thus best interpreted as indications of movement on these unobserved dimensions of identity that is substantial enough to lead to a change in how one identifies on a survey. Previous research on identity has interpreted changes in responses to survey items like these as meaningful indicators of identity shifts, and I do so here.

### **How Americans Shift Their Identities to Align with Their Politics**

Having shown that identity switching is more commonplace than conventional wisdom suggests, I turn to an assessment of the social categorization hypothesis that liberal Democrats and conservative Republicans switch their identities to better conform with the partisan and ideological prototypes shown in Table 1. To detect change over the longest span of time possible, the main analyses of this paper focus on identity change between Wave 1 to Wave 3, or a four-year interval.<sup>7</sup> As is common in studies with panel data, I employ a lagged dependent variable specification in which each individual  $i$ 's identity at Wave 3 is modeled as a function of identity claimed four years earlier at Wave 1, partisanship and

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<sup>7</sup>Additional analyses discussed below confirm similar patterns of identity shifts between Wave 1 and Wave 2 and between Wave 2 and Wave 3.



ideology at Wave 1, and controls for each identity  $j$  in Equation 1.

$$\begin{aligned}
\text{logit}(\text{identify}_{ij, \text{wave}=3}) = & \alpha + \beta_1 \text{identify}_{ij, \text{wave}=1} \\
& + \beta_2 \text{conservatism}_{i, \text{wave}=1} + \beta_3 \text{Republican ID strength}_{i, \text{wave}=1} \\
& + \beta_4 \text{liberalism}_{i, \text{wave}=1} + \beta_5 \text{Democratic ID strength}_{i, \text{wave}=1} \\
& + \text{controls} + \epsilon_i,
\end{aligned} \tag{1}$$

where  $\text{identify}_{ijt}$  takes on the value one if  $i$  identifies as a  $j$  at time  $t$ , and zero if not.

As shown in Equation 1, ideology and partisanship were entered into the model in a way that avoided constraining their impact on identity claiming to be monotonic. The GSS assesses ideology by asking respondents to place themselves on a seven-point scale anchored by “extremely liberal” on one side, “extremely conservative” on the other, and “moderate, middle of the road” at the center. I recoded this variable as the interval-level variable *liberalism*, scored 1 if the respondent identified as “extremely liberal,” .67 if “liberal,” .33 if “slightly liberal,” and zero if the respondent chose moderate, “don’t know,” or any of the conservative responses. The variable *conservatism* was analogously constructed from the conservative responses. Similarly, I recoded the GSS’s seven-point party identification variable into the interval-level variables *Democratic ID strength* and *Republican ID strength*.

Models also included controls for two variables correlated with political affiliations that could potentially confound the politics-identity relationship: age (which can be associated with shifts in identity over time due to life-cycle effects) and Wave 1 educational attainment (which is associated with response stability). Additional controls included respondents’ sex and an indicator variable for the GSS panel in which the respondent participated. The estimation incorporated survey weights for panel non-response supplied by the GSS; robust standard errors were clustered on the GSS’s primary sampling units.

Models were estimated for each of the identities shown in Figure 1, with one substantial adjustment. Many national origins (like English, German, and Irish) are associated with non-Hispanic white identity; others (like Chinese, Mexican, and West Indian) are not. To

test the hypothesis that individuals switch in and out of national origins in ways that align with racial and ethnic political group prototypes, I collapsed the national origin measures into a single variable for which all origins associated with African, Asian, or Hispanic descent were scored one and the remainder scored zero.<sup>8</sup> Self-categorization theory leads to the expectation that liberal Democrats shift toward identification with African, Asian, or Hispanic national origins while conservative Republicans shift away from them.

Model parameter estimates of Equation 1 were used to calculate predictive margins for each identity at Wave 3 for Wave 1 conservative Republicans, Wave 1 liberal Democrats, and (as a baseline) all respondents in the GSS panels.<sup>9</sup> For each identity, these predictive margins were calculated for each observation holding all other variables constant at their actual values and then averaged over the entire dataset using the GSS's sampling weights. Figure 2 displays these estimates, with the baseline Wave 3 mean for each identity set to zero and the predictive margins for liberal Democrats and conservative Republicans plotted as departures from the baseline.<sup>10</sup> These predictions can thus be interpreted as the net probability of each political group shifting into (if the prediction is positive) or out of (if negative) each identity over a four-year period compared to the general population. The left-hand side of the figure reports the differences between the predicted shifts of conservative Republicans and liberal Democrats for each identity; differences statistically significant at  $p < .05$  are displayed in bold type. Where these differences are negative, liberal Democrats are more likely to switch into the identity than conservative Republicans. Where they are positive, the opposite is the case.

The figure confirms that for many identities, the probability of claiming the identity in the present is endogenous to political affiliations in the past. Compared to conservative

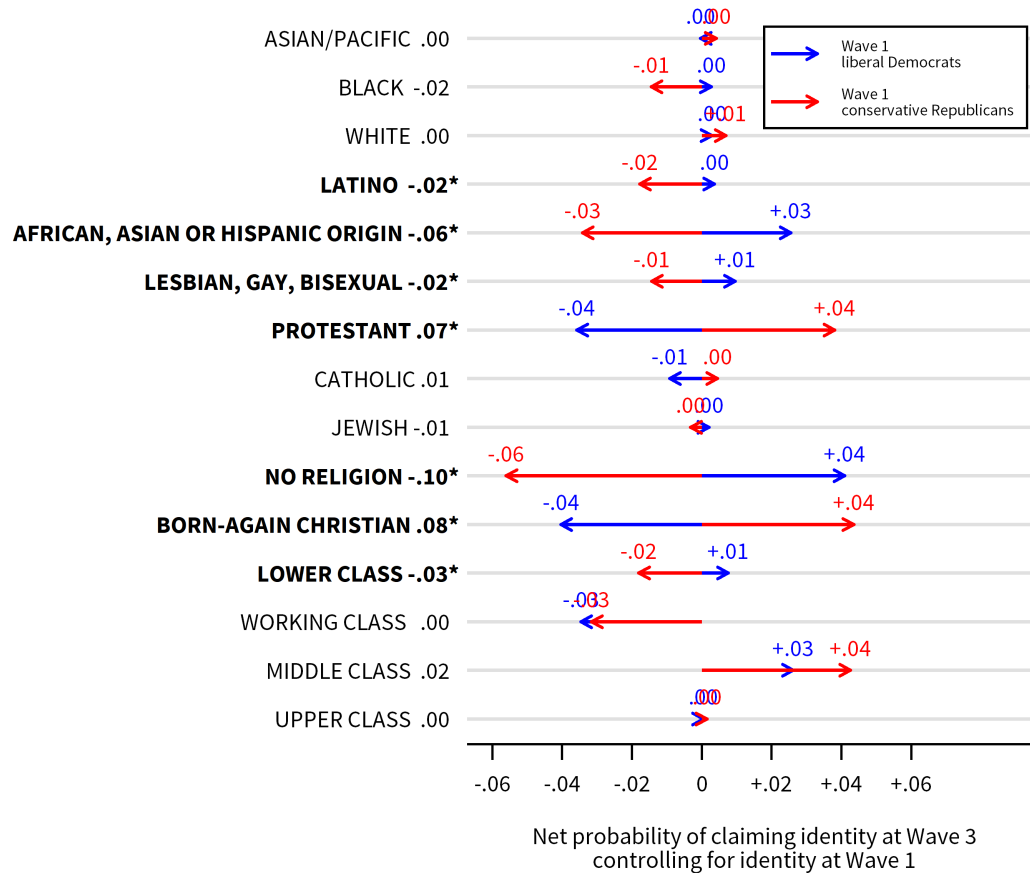
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<sup>8</sup>The nations and origins scored one (with values taken directly from the GSS's nomenclature) were Africa, Arabic, China, India, Japan, Mexico, Philippines, Puerto Rico, the West Indies, and "other Asian" and "other Spanish." The GSS's "other Spanish" category does not include Spain, which was categorized separately and scored zero. This recoded variable includes all national origins, including those with too few identifiers to be included in Figure 1.

<sup>9</sup>For each identity  $j$ , these quantities are respectively  
 $Pr(identify_{j,wave=3} = 1 | RepublicanID\ strength_{wave=1} = .67, conservatism_{wave=1} = .67);$   
 $Pr(identify_{j,wave=3} = 1 | Democratic\ ID\ strength_{wave=1} = .67, liberalism_{wave=1} = .67);$  and  
 $Pr(identify_{j,wave=3} = 1, \text{holding all other individual Wave 1 characteristics constant (including whether one identified as a } j \text{ in Wave 1)}).$  Calculations were performed using the `margins` command in Stata.

<sup>10</sup>Regression output is reported in the SI, pp. 12-13.

**Figure 2. How Partisanship and Ideology Predict Shifts in Group Identities**



*Source: predictive margins from estimated Equation 1.  
Differences between liberal Democrats and conservative Republicans  
that are statistically significant at  $p < .05$  (two-tailed test) displayed in **bold**.*

Republicans, liberal Democrats in Wave 1 were significantly more likely four years later in Wave 3 to switch into claiming identities as Latino, LGB, nonreligious, lower class, and being of African, Asian or Hispanic national origin. In a similar fashion, after controlling for claimed identity in Wave 1, conservative Republicans in Wave 1 were significantly more likely than liberal Democrats four years later to identify as Protestant and as a born-again Christian. It should be noted that the sizes of these shifts are reasonably small, with the

share of either political group's members estimated to shift in or out of any identity in the low single digits. Nevertheless a substantial number of identity shifts are significantly predicted by partisanship and ideology in directions that conform with existing political prototypes.

Additional analyses confirm that the findings in Figure 2 are robust to alternative model specifications and coding decisions.<sup>11</sup> Substantively similar and statistically significant results are obtained when ideology is dropped from Equation 1, leaving Democratic and Republican identity strength remaining as predictors; when controls for respondents' background characteristics (such as religion of upbringing, place of nativity of parents, and number of siblings) are added to Equation 1; and when "don't knows" and other non-responses to identity questions are coded as missing data. The results also hold if Equation 1 is estimated to assess two-year, rather than four-year, shifts in identity (that is, identity shifts from Wave 1 to Wave 2 and shifts from Wave 2 to Wave 3). Finally, estimating Equation 1 separately for those consistently identifying over the three survey waves as either non-Hispanic white or as a person of color finds that political prototypes significantly predict identity shifts among both groups, and yields suggestive evidence that politicized identity shifting with regard to religion and sexuality is more common among whites, while shifting with regard to national origin is more prevalent among people of color.

To address the concern that these results may be attributable to measurement error, I conducted placebo tests using repeated GSS measures of variables whose true values are facts that do not change over time. Some of these variables are correlated with partisanship and ideology (including respondents' recalled region of residence at age 16, year of birth, and parents' educational attainment) and some are not (such as respondents' recalled astrological sign). Wave 1 and Wave 3 measures of these variables were substituted for the Wave 1 and Wave 3 identities in a series of regressions similar in every other respect to the estimated Equation 1. The same predictive margins were calculated for Wave 1 liberal Democrats and Wave 1 conservative Republicans, with the placebo expectation that no significant differences should be found in the differences between the predicted shifts of the

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<sup>11</sup>Results and detailed descriptions of these analyses can be found in the SI, pp. 14-15.

two political groups for any of these variables. As shown in the SI (p. 16), this was the case: the share of the placebo tests (1 out of 27, or 3.7%) finding significant differences between liberal Democrats and conservative Republicans at the .05 level was less than would be expected by chance.

I also conducted sensitivity analyses exploring how vulnerable the results in Figure 2 are to measurement error. As explained above, the standard Heise/Wiley-Wiley measurement model used to assess reliability of measures in three-wave panel surveys is inappropriate for use in generating reliability estimates of identity measures. I therefore conducted sensitivity analyses to determine the minimum level of reliability ( $\rho_{min}$ ) for each identity measure that was required for the significant results reported in Figure 2 to hold in OLS errors-in-variables estimations of Equation 1. As shown in the SI (p. 17), these minimum levels of reliability were generally less than or equal to reliabilities calculated by researchers for other GSS measures of slowly changing demographic characteristics (e.g., Hout and Hastings 2016, 984).<sup>12</sup> The sensitivity analysis also reveals that several non-significant identity shifts in theoretically expected directions in Figure 2 become statistically significant in the errors-in-variables estimations at levels of reliability just slightly less than 1. Taken together, all of these tests help to rule out the concern that the key conclusions of this study are jeopardized by error in the measurement of panelists' identities.

### **Political Prototypes and Politicized Identity Change**

A comparison of the identity-switching patterns shown in Figure 2 with the dissimilarity index scores calculated for each identity in Table 1 provides strong support for the hypothesis that these identity shifts comport with political prototypes. Figure 3a displays this relationship, with the dissimilarity index scores again signed in the negative direction for identities that are over-represented among liberal Democrats and in the positive direction for those over-represented among conservative Republicans. What emerges is a remarkably strong correspondence between prototypes and identity-switching: the more that an identity group's members are concentrated in one of the two political groups, the greater

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<sup>12</sup>For the seven significant results in Figure 2, the range of  $\rho_{min} = \{.69, .70, .83, .86, .90, .93, 1\}$ .

the differences in rates at which political group members tend to switch their identities to align with political group prototypes. The slope of the best-fit regression line in this figure is .12 (with a standard error of .03). This can be interpreted to mean that for an identity with a dissimilarity index score of .3 (and thus in which conservative Republicans are strongly overrepresented), the net probability of conservative Republicans shifting into the identity over time is predicted to be  $.3 \times .12 = .036$ , or 3.6 percentage points higher than for liberal Democrats. As shown in the SI (pp. 14-15), this relationship is consistently statistically significant and of similar magnitude across models employing the alternative specifications and coding decisions discussed earlier.

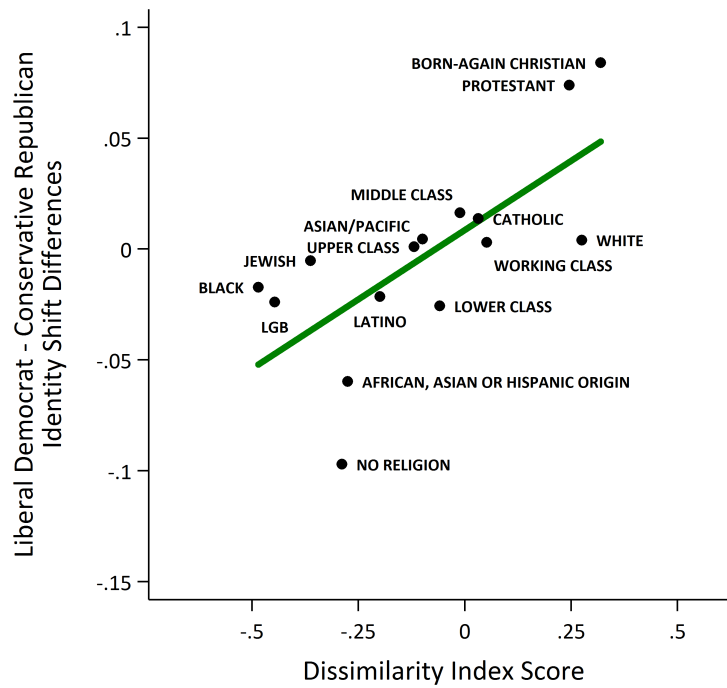
A more rigorous way to assess the relationship between prototype distinctiveness and identity switching is accomplished by pooling all observations across identities for each empaneled respondent and then interacting the dissimilarity index score for each identity with respondents' partisanship and ideology variables as follows, where individuals are indexed  $i$  and identities are indexed  $j$ :

$$\begin{aligned}
 identify_{ij3} = & \alpha + \sum_{j=1}^J \beta_j identify_{ij1} \\
 & + \gamma_1 conservatism_{i1} + \gamma_2 Republican\ ID\ strength_{i1} \\
 & + \gamma_3 liberalism_{i1} + \gamma_4 Democratic\ ID\ strength_{i1} \\
 & + \gamma_1 dissim_j + \delta_2(dissim_j \times conservatism_{i1}) + \delta_3(dissim_j \times Republican\ ID\ strength_{i1}) \\
 & + \delta_4(dissim_j \times liberalism_{i1}) + \delta_5(dissim_j \times Democratic\ ID\ strength_{i1}) \\
 & + controls + \zeta_i + \xi_j + \epsilon_{ij}.
 \end{aligned} \tag{2}$$

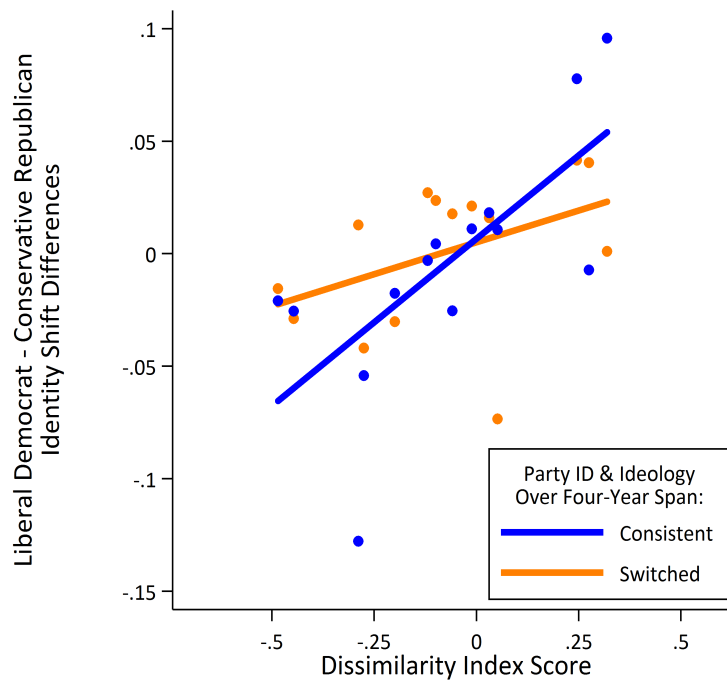
In this model, the dependent variable is again individual  $i$ 's decision to identify as a  $j$  at Wave 3, controlling for identity claimed with regard to  $j$  in Wave 1. Because the data are now pooled across identities, random intercepts  $\zeta_i$  are estimated for each individual and estimated standard errors are clustered at the individual level; the model now also includes fixed intercepts  $\xi_j$  for each identity  $j$ . Here the key coefficients of interest are those on the terms interacting the identities' dissimilarity index scores with individuals' political variables ( $\delta_2, \delta_3, \delta_4, \delta_5$ ). Because the dissimilarity index scores are positively signed for identities overrepresented among conservative Republicans,  $\delta_2$  and  $\delta_3$  are expected to

**Figure 3. The Relationship between Political Prototypes and Politicized Identity Change**

**a. All panelists**



**b. Panelists subsetting by consistency of partisanship and ideology**



**Table 2: Political Prototypes Predict Identity Shifts**

	Wave 1 to Wave 3 identity shifts (1)	$t - 1$ to $t$ identity shifts (2)	model accounting for selection (3)
identity <sub>j</sub> dissimilarity score...			
... $\times$ conservatism ( $\delta_2$ )	0.06* [0.02]	0.04* [0.01]	0.11* [0.02]
... $\times$ Republican ID strength ( $\delta_3$ )	0.07* [0.02]	0.11* [0.01]	0.21* [0.03]
... $\times$ liberalism ( $\delta_4$ )	-0.06* [0.02]	-0.03* [0.01]	-0.01 [0.03]
... $\times$ Democratic ID strength ( $\delta_5$ )	-0.05* [0.02]	-0.03* [0.01]	-0.04 [0.03]
$p$ -value, joint significance test of $\delta_2, \delta_3, \delta_4, \delta_5$	<.001	<.001	<.001
panelist $N$	3,856	4,637	3,872
predicted net identity shift difference between conservative Republicans and liberal Democrats for identity with dissimilarity score = .3 = $.3 \times .67 \times [(\delta_2 + \delta_3) - (\delta_4 + \delta_5)]$	.048	.042	.074

*Dependent variable: claiming identity  $j$  at wave  $t$ . Displayed are OLS regression coefficients of interest from estimated Equation 2 and related models described in text.*

*Standard errors in brackets (clustered on respondent).*

*\*Coefficients statistically different from zero at  $p < .05$  (two-tailed test).*

be positive, reflecting conservatives' and Republicans' tendencies to switch identities to better comport with the conservative Republican prototype. For similar reasons,  $\delta_4$  and  $\delta_5$  are expected to be negatively signed, reflecting liberals' and Democrats' propensities to switch identities to align with the liberal Democratic prototype.

I test this hypothesis with several different specifications, as shown in Table 2, which reports the coefficients on the interaction terms of interest. For ease of interpretation of the coefficients, all models are estimated via OLS; analogous logit regression models yielded substantively similar results.<sup>13</sup> Column 1 displays coefficients from estimated Equation 2. Column 2 displays estimated coefficients from a model similar to Equation 2, except identity at Wave  $t$  was predicted by identity at Wave  $t - 1$ , political variables at  $t - 1$ , and controls. This yielded two observations per identity for each GSS panelist with complete data across all three waves (that is, Wave 3 identity predicted by Wave 2 variables, and

<sup>13</sup>Full regression output is shown in the SI, pp. 18-19.



Wave 2 identity predicted by Wave 1 variables). This specification also includes one observation for each panelist who did not complete Wave 3 of the panel.

The third approach (column 3) is a cross-sectional counterfactual model for three-wave panel data developed by Stephen L. Morgan and Christopher Winship (2007) as adapted by Michael Hout and Claude S. Fischer in their study of shifts in religious identity (2014). It accounts for the concern that unobserved characteristics might be leading people to first identify as liberal Democrats or conservative Republicans and then undertake an identify shift. For example, the unmeasured personality trait of “conscientiousness” might be responsible for someone both coming to identify as a conservative Republican and subsequently identifying as a born-again Christian. If these selection effects are unaccounted for, identity shifts will be inaccurately attributed to political affiliations. The model controls for selection effects with indicator variables scored zero or one depending on whether respondents ever identified as a liberal, conservative, Democrat, or Republican over any of the three waves of the panel:

$$\begin{aligned}
identify_{ijt} = & \alpha + \beta_1 conservat_{it} + \beta_2 Republican\ ID\ strength_{it} \\
& + \beta_3 liberal_{it} + \beta_4 Democratic\ ID\ strength_{it} \\
& + \delta_1 dissim_j + \delta_2 (dissim_j \times conservat_{it}) + \delta_3 (dissim_j \times Republican\ ID\ strength_{it}) \\
& + \delta_4 (dissim_j \times liberal_{it}) + \delta_5 (dissim_j \times Democratic\ ID\ strength_{it}) \\
& + \gamma_1 year_t + \gamma_2 ever\ conservative_i + \gamma_3 ever\ Republican_i + \gamma_4 ever\ liberal_i + \gamma_5 ever\ Democrat_i \\
& + \gamma_6 (year_t \times ever\ conservative_i) + \gamma_7 (year_t \times ever\ Republican_i) + \gamma_8 (year_t \times ever\ liberal_i) \\
& + \gamma_9 (year_t \times ever\ Democrat_i) + \gamma_{10} (dissim_j \times year_t) + \gamma_{11} (dissim_j \times ever\ conservative_i) \\
& + \gamma_{12} (dissim_j \times ever\ Republican_i) + \gamma_{13} (dissim_j \times ever\ liberal_i) + \gamma_{14} (dissim_j \times ever\ Democrat_i) \\
& + \gamma_{15} (dissim_j \times year_t \times ever\ conservative_i) + \gamma_{16} (dissim_j \times year_t \times ever\ Republican_i) \\
& + \gamma_{17} (dissim_j \times year_t \times ever\ liberal_i) + \gamma_{18} (dissim_j \times year_t \times ever\ Democrat_i) \\
& + controls + \zeta_i + \xi_j + \epsilon_{ijt},
\end{aligned}$$

where random intercepts  $\zeta_i$  are estimated for each individual  $i$  and fixed intercepts  $\xi_j$  are estimated for each identity  $j$ , with standard errors clustered on  $i$ . The four political affiliation selection indicator terms (*ever liberal*, *ever conservative*, *ever Democrat*, and *ever Republican*) are interacted with year of survey, creating separate intercepts and time-trends for

each. The model therefore controls for the possibility that selection into any of the four political affiliations may be associated with either different levels of the dependent variable or with different trends in values of the dependent variable over time. All of these selection variables are interacted with the dissimilarity score; the coefficients of interest remain those on terms interacting the identities' dissimilarity index scores with individuals' political variables ( $\delta_2, \delta_3, \delta_4, \delta_5$ ). Only respondents who completed all three waves of the survey are included in this analysis.

Table 2 shows that across all of these specifications, all coefficients are signed in the theoretically expected direction. Positive coefficients on the interaction terms between the dissimilarity index score and conservatism and Republican identification strength confirm that conservative Republicans shift their identities to align with dissimilarity index scores signed in a positive direction. Negative coefficients on the other two interaction terms demonstrate that the opposite is true for liberal Democrats. Most of the interaction term coefficients are statistically significant at the .05 level, but because the four ideology and partisanship variables are highly multicollinear an assessment of their joint statistical significance is the more appropriate test. As shown in the table, the coefficients are highly jointly significant (at  $p < .001$ ) across all specifications. The final row of the table reports predicted probabilities analogous to that used previously to interpret Figure 3a: the net probability of conservative Republicans compared to liberal Democrats shifting into an identity with a dissimilarity index score of .3 (and thus in which conservative Republicans are strongly overrepresented).<sup>14</sup> These predicted probabilities range from 4.2 to 7.4 percentage points, estimates of a larger magnitude than the simple bivariate estimate reported earlier of 3.6 points derived from the slope of the best-fit regression line in Figure 3a.

Self-categorization and depersonalization theory leads us to expect that panelists whose party identification and ideology are consistent over time to be more likely to engage in politicized identity change compared to those whose shift their party identification or ide-

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<sup>14</sup>This is calculated as:

$$.3 \times [Pr(identify_{j,wave=3} = 1 | Republican\ ID\ strength_{wave=1} = .67, conservatism_{wave=1} = .67) - Pr(identify_{j,wave=3} = 1 | Democratic\ ID\ strength_{wave=1} = .67, liberalism_{wave=1} = .67)]$$

**Table 3: Politicized Identity Shifts are More Prevalent Among Americans Whose Partisanship and Ideology Remain Consistent Over Time**

panelist party ID & ideology:	Wave 1 to Wave 3 identity shifts		$t - 1$ to $t$ identity shifts		model accounting for selection	
	consistent (1)	switched (2)	consistent (3)	switched (4)	consistent (5)	switched (6)
identity <sub>j</sub> dissimilarity score...						
... × conservatism ( $\delta_2$ )	0.08* [0.02]	-0.04 [0.05]	0.07* [0.01]	-0.06 [0.04]	0.15* [0.03]	-0.11 [0.06]
... × Republican ID strength ( $\delta_3$ )	0.07* [0.02]	0.00 [0.05]	0.09* [0.01]	0.09* [0.04]	0.18* [0.03]	0.17* [0.07]
... × liberalism ( $\delta_4$ )	-0.04 [0.02]	-0.14* [0.05]	-0.04* [0.01]	-0.10* [0.04]	-0.04 [0.03]	-0.15* [0.06]
... × Democratic ID strength ( $\delta_5$ )	-0.04* [0.02]	-0.02 [0.04]	-0.03* [0.01]	-0.03 [0.03]	-0.08* [0.03]	0.02 [0.06]
panelist $N$	2,920	659	2,920	659	2,929	661
$p$ -value, test that $\delta_2, \delta_3, \delta_4, \delta_5$ are jointly significantly different between the two groups	<.001		.066		<.001	
predicted net identity shift difference between conservative Republicans and liberal Democrats for identity with dissimilarity score = .3 = $.3 \times .67 \times [(\delta_2 + \delta_3) - (\delta_4 + \delta_5)]$	.046	.024	.046	.032	.090	.038

*Dependent variable: claiming identity  $j$  at wave  $t$ . Displayed are OLS regression coefficients of interest from estimated Equation 2 and related models described in text. Data are subset by whether panelists' party ID or ideology completely switched at some point during the four-year panel.*

*Standard errors in brackets (clustered on respondent).*

*\*Coefficients statistically different from zero at  $p < .05$  (two-tailed test).*

ology. As a face validity check of the main results, I subset panelists who completed all three waves of the panel into two groups. The first group consisted of panelists who completely changed either their party identification, their ideology, or both at some point during the four-year span of the panel. To qualify for this group, a panelist had to switch from identifying as a liberal to a conservative (or vice versa), or from identifying as a Democrat to a Republican (or vice versa). The second group was made up of the remainder of panelists, whose party identification and ideology were more consistent. Members of this group either maintained the same party identification and ideology across all three interviews or registered relatively minor changes in that they shifted in or out of identification as an ideological moderate or an independent partisan.

I re-estimated the models giving rise to Figure 3a—that is, estimates of Equation 1 for each identity—separately on the subsetted data. The estimated relationships between politicized identity shifts and identity dissimilarity scores for the two groups are plotted in Figure 3b. The flatter slope of the best-fit regression line for those who switched party identification or ideology (shown in orange) compared to the steeper slope for more consistent panelists (shown in blue) suggests that politicized identity change is more common among individuals for whom partisan and ideological identities are steady over time.<sup>15</sup>

To again conduct more rigorous tests, I re-estimated the models giving rise to Table 2 separately on the subsetted data. Table 3 displays the analogous regression coefficients from these estimated models.<sup>16</sup> The table shows that for consistent panelists, all coefficients are signed in theoretically expected directions (that is,  $\delta_2, \delta_3 > 0$  and  $\delta_4, \delta_5 < 0$ ). By contrast, among panelists who completely shifted their partisanship or ideology, many coefficients are signed in directions opposite than congruent with politicized identity change. Test statistics indicate that these sets of coefficients are significantly different between the two groups.<sup>17</sup> The final row of Table 3 quantifies these differences, again reporting the net probability of conservative Republicans compared to liberal Democrats shifting into an identity with a dissimilarity index score of .3. Averaged across the three specifications, Americans with consistent partisanship and ideology are roughly twice as likely to engage in politicized identity shifting than those whose political affiliations switch over time.

## When Identities Are Infused with Politics

While thus far the analysis has been focused on the implications of politicized identity change for individuals, these shifts can also have consequences for identity groups them-

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<sup>15</sup>The slope for consistent panelists (.15, with a standard error of .04) is twice the magnitude of the slope for switching panelists (.06, with a standard error of .04). The slopes are not statistically significantly different from one another ( $p = .20$ ); the power of this significance test is reduced by the small effective number of observations (as the number of identities analyzed here  $J = 14$ ).

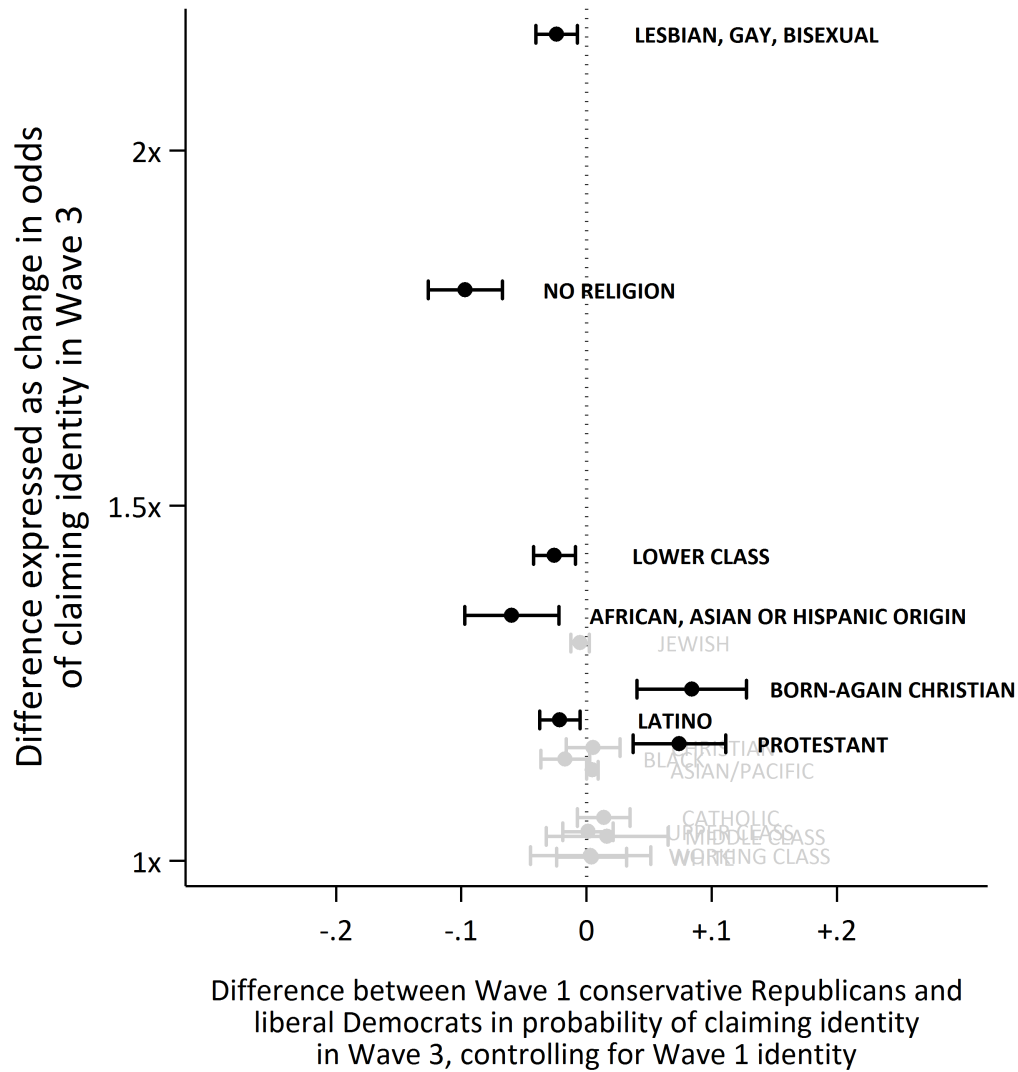
<sup>16</sup>Results were substantively similar when the models were estimated with logit. Full regression output is shown in the SI, pp. 20-22.

<sup>17</sup>These were calculated by re-combining the data, assigning a variable indicating if panelists were consistent partisans and ideologues, and running the same models with this indicator variable interacted with  $\delta_2, \delta_3, \delta_4$ , and  $\delta_5$ . The test statistic reported in the table is that of the joint significance test of these four interaction terms.

selves if it is the case that partisanship and ideology predict shifts in identification with the group that are large relative to the size of the group in the population. As an example, consider lesbians, gays and bisexuals, a group that makes up about three percent of the U.S. adult population according to the GSS panel data. While Figure 2 shows that the association between political variables and identity shifts with regard to LGB identity appears to be small (the difference between liberal Democrats and conservative Republicans in net probability of change is two percentage points), this is actually quite a substantial change given the relatively small size of this group in the U.S. population. This is seen by considering another way to assess the magnitude of the shift, which is to transform it into the the change in odds of claiming the identity at Wave 3 controlling for Wave 1 identity between the two political groups. Controlling for Wave 1 identity, the probability that conservative Republicans claim LGB identity at Wave 3 is .021 compared to .045 for liberal Democrats—a more than a doubling of the odds of claiming LGB identity.

Figure 4 displays this alternative measure of the influence of politics on identity group membership. On the  $x$ -axis, this graph displays the differences between conservative Republicans and liberal Democrats in the net probability of claiming identities in Wave 3, controlling for Wave 1 identity. (These are the same values displayed on the left-hand side of Figure 2.) Plotted on the  $y$ -axis are these same predictive margins now expressed as the absolute value of the net difference in odds of claiming the identity between the two political groups. Here we see that of all the identities in this study, by this measure lesbian, gay and bisexual identity is by far the identity most endogenous to politics. In line with previous research (Hout and Fischer 2014), identification as a religious “none” is also heavily infused with politics. By contrast, the memberships of other identity groups whose membership is significantly predicted by political variables—including born-again Christian, Latino, and Protestant—are not substantially infused with politics, as political variables are associated with relatively small changes in the odds of identification. One characteristic that LGBs and the areligious have in common is that identification with them is generally acquired later in life rather than shared between parent to offspring. It thus intuitively makes sense that selection into these two groups would be unusually shaped by politics,

Figure 4. Identity Claiming that is Particularly Infused with Politics



Source: Predictive margins from estimated Equation 1, with statistically significant differences at  $p < .05$  in **bold**.

given that intergenerational transmission is unlikely to be a channel for identification with them.

## Conclusion

This paper makes two main contributions. First, in contrast to previous work that has looked at identities one at a time, here I demonstrate that the range of identity categories subject to override in the short term by partisanship and ideology is wider than shown

in previous research. Second, in contrast to previous work that has largely provided explanations of identity change grounded in the particular identity being studied, I offer a generalized explanation, grounded in social identity theory, for the circumstances that are favorable for politicized identity shifting.

These findings yield new insight on the nature of politically salient American identities and how they can be shaped by the liberal-conservative, Democrat-Republican divide. Inter-temporal stability varies highly among identities, and some of this over-time instability in identification is associated with politics. Liberalism and Democratic party identification predicts shifts toward identification as Latino, lesbian, gay, or bisexual, as nonreligious, lower class, and claiming non-European national origin; conservatism and Republican party ID yields movement toward identification as being a member of Protestant faith, and having had an experience as a born-again Christian. This is no small discovery: many of these identities are at the center of important American policy debates, and those who claim these identities are key blocs of voters, party activists and political donors. Furthermore, in many theoretical and empirical models of politics, these identities are treated as exogenous.

Under what circumstances might we expect to see this phenomenon arise in other contexts? Two conditions would seem to be necessary for such a development. First, political groups must be so highly salient and central to the self-concept that they become identities in themselves. And second, political groups must become distinct enough with regard to ethnicity, religion, sexuality or other characteristics typically considered fixed that these characteristics are readily called to mind as components of political group prototypes. The presence of both of these conditions create fertile ground for the activation of self-categorization and depersonalization processes that result in politicized identity shifting. The period studied here—2006 through 2014—is a context when conditions became particularly favorable for such a process to unfold in the United States. The pace of social sorting accelerated with regard to Asian/Pacific identity and religiosity (Mason 2018b), while the election of the nation's first African-American president and nationwide battles over marriage equality reaffirmed political lines previously drawn on the basis of race, ethnicity

and sexuality, creating circumstances where politicized identity shifting was particularly likely.

Nearly sixty years ago, the “Michigan school” authors of *The American Voter* noted that the influence of group membership on political behavior might be overstated, as members of many identity groups often “come to identify with the group on the basis of pre-existing beliefs and sympathies” (Campbell et al 1960, 323). While these researchers did not have categories like ethnicity, religion, or sexuality in mind, the findings presented here join mounting evidence that this concern applies even to identities generally considered fixed, and that more rich discoveries await those who continue to make use of powerful tools and data to understand the origins of important identities in politics.



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**Identity as Dependent Variable:  
How Americans Shift Their Identities to Better Align with Their Politics**

**SUPPORTING INFORMATION**

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**SI Table 1. Identity Claiming and Switching in the U.S. Adult Population**

Identity	% of population claiming identity in Wave 2		% of Wave 2 identity claimers switching identity (did not hold identity in Wave 1, abandoned identity in Wave 3, or both)	
	mean	se	mean	se
<b>Race/ethnicity</b>				
ASIAN/PACIFIC	3.6	(1.0)	6.3	(2.1)
BLACK	13.6	(1.7)	7.3	(1.4)
WHITE	76.6	(2.3)	5.6	(0.9)
LATINO*	12.2	(2.2)	6.6	(2.3)
<b>National origin</b>				
AFRICAN	7.7	(1.0)	46.0	(3.7)
AMERICAN INDIAN	2.6	(0.4)	85.9	(4.1)
AMERICAN ONLY	2.3	(0.4)	93.8	(2.3)
ENGLISH	11.7	(0.9)	54.0	(2.8)
FRENCH	2.4	(0.3)	69.2	(5.1)
GERMAN	16.8	(1.5)	39.5	(2.4)
IRISH	10.3	(0.7)	50.1	(3.1)
ITALIAN	4.8	(0.7)	34.8	(5.5)
MEXICAN	6.6	(1.6)	13.4	(2.3)
POLISH	2.5	(0.3)	36.8	(5.2)
SCOTTISH	3.1	(0.4)	73.5	(5.3)
<b>Sexual orientation</b>				
LESBIAN, GAY, BISEXUAL (LGB)	3.3	(0.5)	46.5	(9.2)
<b>Religion</b>				
PROTESTANT	49.5	(2.6)	17.9	(1.2)
CATHOLIC	24.2	(1.9)	13.6	(1.3)
JEWISH	2.2	(0.6)	18.5	(5.2)
NO RELIGION	17.2	(1.1)	39.1	(2.3)
BORN-AGAIN CHRISTIAN**	38.4	(2.3)	29.1	(1.9)
<b>Class</b>				
LOWER	7.0	(0.5)	68.2	(3.5)
WORKING	46.7	(1.4)	44.0	(1.6)
MIDDLE	43.2	(1.4)	38.1	(1.7)
UPPER	2.8	(0.3)	69.0	(4.2)

\*Latinos can be of any race. \*\*Born-again Christians can be of any religion.

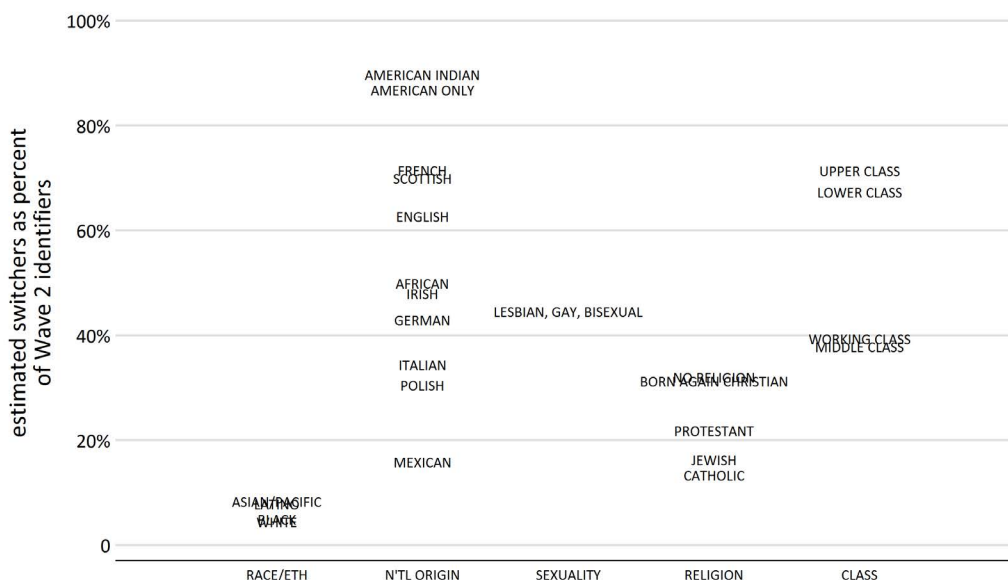
Valid N = 3,875 for all identities except LGB, where N = 1,099.

## SI Figure 1. Description of fixed-effects estimation of identity switching

The estimates displayed below come from analysis conducted by restacking the dataset so that the unit of observation was at the individual ( $i$ )  $\times$  identity ( $j$ ) level, generating 26 observations—one for each of the  $J=26$  identities displayed in Figure 1a—for each individual  $i$ . Each  $ij$  observation included three indicator variables ( $identify_{ij1}$ ,  $identify_{ij2}$ ,  $identify_{ij3}$ ) scored one if  $i$  identified as a  $j$  at wave  $t$  and zero if not. A final indicator variable,  $switch_{ij}$ , was scored one if  $i$ 's responses regarding identity  $j$  switched at all during the panel; that is if  $identify_{ij2} \neq identify_{ij1}$ , or  $identify_{ij2} \neq identify_{ij3}$ . I then estimated the model

$$switch_{ij} = \sum_{j=1}^{26} \alpha_j + \beta_j identify_{ij2} + \gamma_j (identify_{ij2} \times J) + \sum_{i=1}^N v_i + \varepsilon_{ij},$$

where  $J$  is an indicator variable assigned to each of the  $j = 1 \dots 26$  identity groups; and  $v_i$  is a fixed-effects estimate of  $i$ 's tendency to switch answers to survey questions about identity. The interaction  $identify_{ij2} \times J$  is included in the specification to allow for different rates of switching among those who identify as a  $j$  in wave 2 and those who do not. Taken together, the parameter estimates  $\hat{\alpha}_j$ ,  $\hat{\beta}_j$ , and  $\hat{\gamma}_j$  were used to estimate  $\Pr(switch_j | identify_{j2} = 1)$ , a measure of the average consistency with which people identify with group  $j$ , controlling for any correlation in individual response stability with the tendency to identify with  $j$ . I estimated this equation as a linear probability model, using survey weights for panel non-response and clustering standard errors by individual  $i$ . I then generated the predicted probabilities of an individual's switching in or out of an identity  $j$  given that she identified as a  $j$  in Wave 2, or  $\Pr(switch_j | identify_{j2} = 1)$ . These estimates are displayed below, a comparison of which with Figure 1a suggests that the relative stability of most identities is substantively similar across both measures.



# Implications of a Violation of the Lag-1 Assumption in the Heise/Wiley-Wiley Model of Reliability in Three-Wave Panels

## The Heise/Wiley-Wiley Model

Estimates of the reliability of survey items in three-wave panel designs are derived from a model introduced by Heise (1969) and further developed by Wiley and Wiley (1970). The Heise/Wiley-Wiley (HWW) model begins with a measurement model of the form

$$Y_{i1} = T_{i1} + E_{i1}$$

$$Y_{i2} = T_{i2} + E_{i2}$$

$$Y_{i3} = T_{i3} + E_{i3},$$

where for any individual  $i$ ,  $T_{it}$  is the (mean-deviated) unobserved *true score* at time  $t$ ,  $Y_{it}$  is the *observed score* at time  $t$ , and  $E_{it}$  is random measurement error at time  $t$ .<sup>1</sup> Rather than being constant over time, the true scores are assumed to vary according to the structural model

$$T_{it} = \beta_{t,t-1}T_{i,t-1} + Z_{it},$$

where  $\beta_{t,t-1}$  is a parameter specifying how stable  $T$  is from one period to the next, and  $Z_{it}$  is a random disturbance associated with individual  $i$ . Crucially, true scores are assumed to change via a lag-1, or Markovian, process in which the value of  $T$  at time  $t$  is dependent only on its lagged value—and not on the value of  $T$  at any previous time.

Applied to a three-wave panel, this model is written (dropping  $i$  for simplicity):

$$T_1 = Z_1$$

$$T_2 = \beta_{21}T_1 + Z_2$$

$$T_3 = \beta_{32}T_2 + Z_3$$

$$= \beta_{32}(\beta_{21}T_1 + Z_2) + Z_3,$$

Plugging this into the measurement model gives

$$Y_1 = Z_1 + E_1$$

$$Y_2 = \beta_{21}Z_1 + Z_2 + E_2$$

$$Y_3 = \beta_{32}(\beta_{21}Z_1 + Z_2) + Z_3 + E_3.$$

All disturbances  $Z$  are assumed to have an expected value of zero, to be uncorrelated over time, and to be uncorrelated with the true scores and with measurement error. Measurement errors  $E$  are also assumed to have an expected value of zero, to be uncorrelated over time and to be uncorrelated with true scores. We can therefore simply write the variances of the observed scores as

$$\text{VAR}(Y_1) = \text{VAR}(Z_1) + \text{VAR}(E_1)$$

$$\text{VAR}(Y_2) = \beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)$$

$$\text{VAR}(Y_3) = \beta_{32}^2 [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \text{VAR}(Z_3) + \text{VAR}(E_3),$$

---

<sup>1</sup>Here I largely follow Alwin's (2007, pp 104-107) notation of the model.

and the covariances of the observed scores as

$$\begin{aligned} \text{COV}(Y_1, Y_2) &= \beta_{21} \text{VAR}(Z_1) \\ \text{COV}(Y_1, Y_3) &= \beta_{32} \beta_{21} \text{VAR}(Z_1) \\ \text{COV}(Y_2, Y_3) &= \beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)]. \end{aligned}$$

### Reliability in the HWW Model

*Reliability*, denoted  $\rho^2$ , is defined as the ratio of the variance of true scores to that of observed scores.  $\rho^2$  approaches one as observed scores approach perfect alignment with true scores and the variance of  $E$  approaches zero:

$$\rho^2 = \frac{\text{VAR}(T)}{\text{VAR}(Y)} = \frac{\text{VAR}(Z)}{\text{VAR}(Y)} = \frac{\text{VAR}(Y) - \text{VAR}(E)}{\text{VAR}(Y)} = 1 - \frac{\text{VAR}(E)}{\text{VAR}(Y)}.$$

Here Heise and Wiley and Wiley part ways: Heise assumes that reliability is constant across time periods; Wiley and Wiley do not. This distinction is not relevant here, as I focus on reliability at  $t = 2$ , which both approaches agree is

$$\begin{aligned} \rho_2^2 &= 1 - \frac{\text{VAR}(E_2)}{\text{VAR}(Y_2)} = 1 - \frac{\text{VAR}(Y_2) - \beta_{21}^2 \text{VAR}(Z_1) - \text{VAR}(Z_2)}{\text{VAR}(Y_2)} \\ &= \frac{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)}{\text{VAR}(Y_2)} \\ &= \frac{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)}{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)}. \end{aligned} \tag{1}$$

Here,  $\rho_2^2$  (which for simplicity I denote going forward as  $\rho^2$ ) is written as a function of primitive, unobservable elements of the model. Heise and Wiley and Wiley show that the same quantity can be calculated from the following ratio of the variances and covariances of observable scores  $Y$ :

$$\begin{aligned} \frac{\text{COV}(Y_1, Y_2) \text{COV}(Y_2, Y_3)}{\text{COV}(Y_1, Y_3) \text{VAR}(Y_2)} &= \frac{\beta_{21} \text{VAR}(Z_1) \beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)]}{\beta_{32} \beta_{21} \text{VAR}(Z_1) [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)]} \\ &= \frac{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)}{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)} \\ &= \rho^2. \end{aligned}$$

### When the Lag-1 Assumption Fails to Hold

As noted above, a key assumption of the HWW model is that true scores  $T$  change via a lag-1 autocorrelation process. It is a “short memory” model, in that after accounting for one lagged value of the true score  $T$ , it is assumed that no previous information about  $T$  meaningfully predicts its current value. This is an untestable assumption that on its face would seem difficult to justify in analyses of long-term processes that unfold over many years, such as the formation of identity.



What happens to estimates of  $\rho^2$  when the lag-1 assumption fails to hold? I investigate this question by developing a model that is similar the HWW model in every respect except that true scores follow a lag-2 process. In this revised model, all assumptions remain the same with one exception: the lag-2 nature of the process is reflected by adding the term  $\beta_{31}T_1$  to the equation for  $T_3$ , where  $\beta_{31}$  represents the lagged direct effect of  $T_1$  on  $T_3$ , controlling for  $T_2$ :

$$\begin{aligned}T_1 &= Z_1 \\T_2 &= \beta_{21}T_1 + Z_2 \\T_3 &= \beta_{32}T_2 + \beta_{31}T_1 + Z_3 \\&= \beta_{32}(\beta_{21}T_1 + Z_2) + \beta_{31}Z_1 + Z_3.\end{aligned}$$

The observed scores are now

$$\begin{aligned}Y_1 &= Z_1 + E_1 \\Y_2 &= \beta_{21}Z_1 + Z_2 + E_2 \\Y_3 &= \beta_{32}(\beta_{21}Z_1 + Z_2) + \beta_{31}Z_1 + Z_3 + E_3,\end{aligned}$$

and the variances and covariances of the observed scores become

$$\begin{aligned}\text{VAR}(Y_1) &= \text{VAR}(Z_1) + \text{VAR}(E_1) \\ \text{VAR}(Y_2) &= \beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2) \\ \text{VAR}(Y_3) &= \beta_{32}^2 [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \beta_{31}^2 \text{VAR}(Z_1) + \text{VAR}(Z_3) + \text{VAR}(E_3) \\ \\ \text{COV}(Y_1, Y_2) &= \beta_{21} \text{VAR}(Z_1) \\ \text{COV}(Y_1, Y_3) &= (\beta_{32}\beta_{21} + \beta_{31}) \text{VAR}(Z_1) \\ \text{COV}(Y_2, Y_3) &= \beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \beta_{21}\beta_{31} \text{VAR}(Z_1).\end{aligned}$$

Calculations of  $\rho^2$  using the standard HWW formula for reliability now return a biased estimate of  $\rho^2$ , denoted here as  $\rho_B^2$ :

$$\begin{aligned}\rho_B^2 &= \frac{\text{COV}(Y_1, Y_2)\text{COV}(Y_2, Y_3)}{\text{COV}(Y_1, Y_3)\text{VAR}(Y_2)} \\ &= \frac{\beta_{21} \text{VAR}(Z_1) [\beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \beta_{21}\beta_{31} \text{VAR}(Z_1)]}{(\beta_{32}\beta_{21} + \beta_{31}) \text{VAR}(Z_1) [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)]} \\ &= \frac{\beta_{21} [\beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \beta_{21}\beta_{31} \text{VAR}(Z_1)]}{(\beta_{32}\beta_{21} + \beta_{31}) [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)]}.\end{aligned}$$

Substituting the expression for  $\rho^2$  from (1) above, write the bias of  $\rho_B^2$  for  $\rho^2$  as

$$\begin{aligned}\text{BIAS}[\rho_B^2] &= \rho_B^2 - \rho^2 \\ &= \frac{\beta_{21} [\beta_{32} [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)] + \beta_{21}\beta_{31} \text{VAR}(Z_1)]}{(\beta_{32}\beta_{21} + \beta_{31}) [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)]} - \frac{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2)}{\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)} \\ &= -\frac{\beta_{31}(\beta_{21}(\beta_{21} - 1)\text{VAR}(Z_1) + \text{VAR}(Z_2))}{(\beta_{21}\beta_{32} + \beta_{31}) [\beta_{21}^2 \text{VAR}(Z_1) + \text{VAR}(Z_2) + \text{VAR}(E_2)]}.\end{aligned}\tag{2}$$

Four properties of  $\rho_B^2$  provide cause for concern that the HWW formula produces biased estimates of reliability for the identity measures in this paper.

**1.  $\rho_B^2$  is biased for  $\rho^2$  whenever the lag-1 assumption is violated (i.e.,  $\beta_{31}$  is nonzero).**

By inspection, expression (2) indicates that  $\rho_B^2$  simplifies to  $\rho^2$  whenever  $\beta_{31} = 0$ , but  $\beta_{31} \neq 0 \implies \rho_B^2 \neq \rho^2$  in all but trivial, knife-edge cases. This confirms that a violation of the assumption of a lag-1 model—which on its face would seem likely in a “long memory” process like identity change—leads the HWW formula to return a biased calculation of reliability.

Additional analysis indicates that the direction of this bias is likely negative—that is, that  $\rho_B^2$  will underestimate the true  $\rho^2$ —for many of the identity measures in this paper. As auto-regressive parameters,  $\beta_{21}$ ,  $\beta_{31}$  and  $\beta_{32}$  each fall in the range  $(0, 1)$ , making the denominator of the ratio in expression (2) always positive. Thus  $\rho_B^2$  is negatively biased when:

$$-\beta_{31}(\beta_{21}(\beta_{21} - 1)\text{VAR}(Z_1) + \text{VAR}(Z_2)) < 0$$

and thus when

$$(\beta_{21} - \beta_{21}^2)\text{VAR}(Z_1) - \text{VAR}(Z_2) < 0. \quad (3)$$

This expression gives rise to two additional concerns about  $\rho_B^2$  with relevance to the reliability of the measurement of identities in this paper:

**2. When the lag-1 assumption is violated, reliability is more likely to be underestimated for measures of true scores, such as identities, that exhibit a high degree of stability between Wave 1 and Wave 2.**

Inequality (3) is more likely to be satisfied as the quantity  $\beta_{21} - \beta_{21}^2$  grows smaller.  $\beta_{21}$  is the auto-regressive parameter representing the stability of identity between Wave 1 and Wave 2, and as shown below,  $\beta_{21} \in (.5, 1)$  for nearly all of the identities in this paper. In this range,  $\beta_{21} - \beta_{21}^2$  declines monotonically in  $\beta_{21}$ , approaching zero as  $\beta_{21}$  approaches one. Inequality (3) therefore tells that when the lag-1 assumption is violated, the HWW formula is more likely to underestimate the reliability of measures of identities that are highly stable between Wave 1 and Wave 2.

**3. When the lag-1 assumption is violated, reliability is more likely to be underestimated for measures of true scores with a low variance in the population, and thus measures of dichotomous identities with low prevalence in the population.**

Inequality (3) shows that  $\rho_B^2$  underestimates  $\rho^2$  when  $\text{VAR}(Z_1)$  is small relative to  $\text{VAR}(Z_2)$ . Recall that  $\text{VAR}(Z_1) = \text{VAR}(T_1)$ , the variance of the true scores in the population. By contrast,  $\text{VAR}(Z_2)$  is the variance of individual random disturbances to true scores at Wave 2. Thus *ceteris paribus*, any measure of true scores that exhibits a low variance about its mean in the population is more likely to be inaccurately determined unreliable by the HWW method. An important implication of this result relevant to this paper is that when true scores are dichotomous—as are the identities measured in this paper, where  $T_1 \in \{0, 1\}$

and thus  $VAR(T_1) = \mu_T(1 - \mu_T)$ , then the calculated reliabilities of measures of variables with a low prevalence in the population (and thus low variance) will be more likely to underestimate the measure's true reliability when the lag-1 assumption is violated.

A final result identifies additional circumstances under which the underestimation of reliability by  $\rho_B^2$  is aggravated:

**4. As the lag-2 effect of  $T_1$  on  $T_3$ , grows,  $\rho_B^2$  becomes more negatively biased for  $\rho^2$ .**

This is shown by taking the derivative of (2) with respect to  $\beta_{31}$ , which yields

$$\frac{\partial BIAS[\rho_B^2]}{\partial \beta_{31}} = - \frac{\beta_{21}\beta_{32} (VAR(Z_2) + (\beta_{21}^2 - \beta_{21}) VAR(Z_1))}{(VAR(Z_2) + \beta_{21}^2 VAR(Z_1) + VAR(E_2)) (\beta_{21}\beta_{32} + \beta_{31})^2},$$

the sign of which is the sign of

$$(\beta_{21} - \beta_{21}^2) VAR(Z_1) - VAR(Z_2),$$

which is negative whenever inequality (3) is satisfied. Thus if  $\rho_B^2$  is negatively biased for  $\rho^2$ , it is more so to the extent that  $\beta_{31}$  is large.

**Evidence that the HWW Formula May Underestimate Reliability for Identities Measured in the GSS Panel Surveys**

To what extent do these concerns apply to the measures of identities in the GSS? While three-wave panels do not provide enough data to assess the degree of bias directly, here I report estimates that strongly suggest that for the measures of identities in this paper, calculations of reliability using the HWW model are vulnerable to concerns about negative bias.

For each of the 27 identities in the paper, I produced estimates of  $\beta_{21}$  and  $VAR(Z_2)$  with a bivariate regression that generated the estimated model

$$y_{i2} = \hat{\alpha}_2 + \hat{\beta}_{21}y_{i1} + e_{i2}$$

which produces  $\hat{\beta}_{21}$  as an estimate of  $\beta_{21}$  and the variance of the residuals  $VAR(e_{i2})$  as an estimate of  $VAR(Z_2)$ . I generated estimates of  $\beta_{31}$  with a trivariate regression yielding the estimated model

$$y_{i3} = \hat{\alpha}_3 + \hat{\beta}_{32}y_{i2} + \hat{\beta}_{31}y_{i1} + e_{i3}.$$

Finally, as noted above,  $VAR(Z_1)$  is simply the variance of the true scores in the population at Wave 1; I estimated this quantity using  $VAR(Y_1)$ .

Appendix Table 5 displays all of these quantities for each of the identities analyzed in this paper as well as the calculated values of the left-hand side of inequality (3). It shows that for every identity,  $\beta_{31}$  is estimated to be significantly greater than zero, leading us to be concerned about biased estimates of reliability for these identities. Furthermore, for all but one identity—white racial identity—inequality (3) is satisfied in that  $(\beta_{21} - \beta_{21}^2) VAR(Z_1) -$

$VAR(Z_2)$  is estimated to be negative, which means that the HWW formula for reliability is expected to underestimate the reliability of the GSS measures of these identities. In addition, for several identities this negative bias is further aggravated because  $\beta_{31}$  is estimated to be relatively high. This can be seen in Appendix Figure 1, which plots estimated values of  $\beta_{31}$  against the left-hand side of inequality (3). We expect  $\rho_B^2$  to be negatively biased for  $\rho^2$  for measures of identities located to the left of zero on the  $x$ -axis; the negative bias is greater for identities located toward the top of the figure, which include Latino, lesbian, gay or bisexual, Mexican, and African, Asian or Hispanic national origin.

A final piece of suggestive evidence that the MWW model underestimates the reliability of the GSS measures is found when the MWW estimates of  $\rho^2$  are used in errors-in-variables (EIV) regression models to estimate the lag-1 stability parameter  $\beta_{21}$  for each of the identities in this paper. As shown in Appendix Table 6, the EIV models generate estimates of  $\beta_{21}$  that are near—and in some cases, greater than—one, which unrealistically suggest nearly perfect continuity in most identities over the four-year period due to over-correction for measurement error. Tellingly, among the measures there is a strong correspondence between unrealistically high estimates of  $\beta_{21}$  and the measures' vulnerability to underestimated  $\rho^2$  as indicated by their locations toward the top of the plot in Appendix Figure 1.

## Conclusion

This analysis provides reason for general concern that the HWW formula for calculating the reliability of survey measures administered in a three-wave panel—the standard way reliability is calculated for purposes of correcting for measurement error—is vulnerable to concerns about bias when the HWW model's lag-1 assumption fails to hold. Specifically, this analysis suggests that the HWW model underestimates the reliability of the GSS measures of identity employed in this paper, particularly for identities that are (1) relatively stable from one period to another ( $\beta_{21}$  is high); (2) relatively rare in the population ( $VAR(Z_1)$  is low); and (3) are “long memory” in that current identity is explained by more than one lagged value of identity ( $\beta_{31}$  is high). Nearly all—26 out of 27—of the identities analyzed in this paper appear to meet these criteria, making it difficult to have confidence in the MWW estimates of their measures' reliability.

## References

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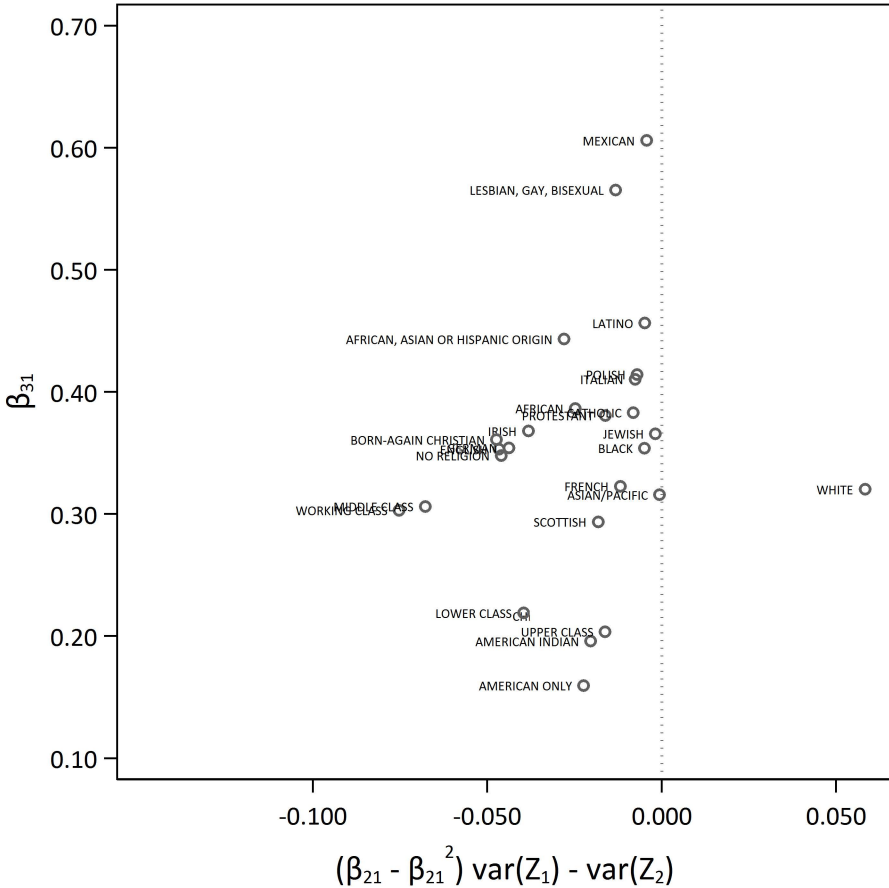
**Table SI 3. Parameter estimates for evaluation of the HWW formula for reliability**

<b>Identity</b>	$\beta_{21}$	se	$VAR(Z_1)$	$VAR(Z_2)$	LHS of ineq (3)	$\beta_{31}$	se
<b>Race/ethnicity</b>							
ASIAN/PACIFIC	0.89	0.04	0.04	0.00	-0.001	0.32	0.10
BLACK	0.95	0.01	0.14	0.01	-0.005	0.35	0.10
WHITE	0.83	0.03	0.76	0.05	0.058	0.32	0.04
LATINO	0.94	0.02	0.12	0.01	-0.005	0.46	0.10
<b>National origin</b>							
AFRICAN	0.63	0.03	0.08	0.04	-0.025	0.39	0.05
AMERICAN INDIAN	0.28	0.05	0.02	0.03	-0.020	0.20	0.05
AMERICAN ONLY	0.12	0.03	0.02	0.02	-0.022	0.16	0.04
ENGLISH	0.55	0.03	0.11	0.07	-0.047	0.35	0.03
FRENCH	0.57	0.05	0.02	0.02	-0.012	0.32	0.07
GERMAN	0.65	0.02	0.17	0.08	-0.044	0.35	0.03
IRISH	0.57	0.03	0.11	0.06	-0.038	0.37	0.03
ITALIAN	0.77	0.04	0.05	0.02	-0.008	0.41	0.06
MEXICAN	0.93	0.02	0.06	0.01	-0.004	0.61	0.08
POLISH	0.67	0.06	0.03	0.01	-0.007	0.41	0.08
SCOTTISH	0.49	0.05	0.03	0.02	-0.018	0.29	0.05
AFRICAN, ASIAN OR HISPANIC ORIGIN	0.77	0.02	0.21	0.07	-0.028	0.44	0.04
<b>Sexuality</b>							
LESBIAN, GAY, BISEXUAL	0.73	0.07	0.03	0.02	-0.013	0.57	0.12
<b>Religion</b>							
PROTESTANT	0.78	0.01	0.49	0.10	-0.016	0.38	0.03
CATHOLIC	0.86	0.01	0.25	0.04	-0.008	0.38	0.04
JEWISH	0.95	0.02	0.02	0.00	-0.002	0.37	0.24
NO RELIGION	0.67	0.02	0.16	0.08	-0.046	0.35	0.03
CHRISTIAN	0.29	0.06	0.03	0.03	-0.023	0.22	0.04
BORN-AGAIN CHRISTIAN	0.70	0.02	0.36	0.12	-0.047	0.36	0.02
<b>Class</b>							
LOWER CLASS	0.45	0.03	0.07	0.06	-0.040	0.22	0.03
WORKING CLASS	0.47	0.02	0.45	0.19	-0.075	0.30	0.02
MIDDLE CLASS	0.52	0.02	0.45	0.18	-0.068	0.31	0.02
UPPER CLASS	0.48	0.05	0.03	0.02	-0.016	0.20	0.05

Table SI 4. HWW reliability and lag-1 stability estimates of GSS identity measures

Identity	$\rho^2$	$\beta_{21}$	se
<b>Race/ethnicity</b>			
ASIAN/PACIFIC	0.93	0.97	0.00
BLACK	0.96	0.99	0.00
WHITE	0.89	0.94	0.01
LATINO	0.92	1.01	0.00
AFRICAN	0.68	0.96	0.01
<b>National origin</b>			
AMERICAN INDIAN	0.29	0.94	0.04
AMERICAN ONLY	0.19	0.83	0.07
ENGLISH	0.54	1.04	0.02
FRENCH	0.60	0.92	0.02
GERMAN	0.67	0.93	0.01
IRISH	0.60	0.95	0.02
ITALIAN	0.80	0.96	0.01
MEXICAN	0.91	1.02	0.01
POLISH	0.73	0.93	0.01
SCOTTISH	0.46	1.04	0.02
AFRICAN, ASIAN OR HISPANIC ORIGIN	0.77	0.99	0.01
<b>Sexuality</b>			
LESBIAN, GAY, BISEXUAL	0.59	1.12	0.03
<b>Religion</b>			
PROTESTANT	0.80	0.97	0.01
CATHOLIC	0.90	0.96	0.01
JEWISH	0.95	0.98	0.01
NO RELIGION	0.70	0.95	0.01
CHRISTIAN	0.39	0.70	0.03
BORN-AGAIN CHRISTIAN	0.73	0.96	0.01
<b>Class</b>			
LOWER CLASS	0.58	0.84	0.02
WORKING CLASS	0.54	0.91	0.02
MIDDLE CLASS	0.53	0.96	0.02
UPPER CLASS	0.53	0.88	0.02

Figure SI 1. Parameter estimates for evaluation of the HWW formula for reliability



SI Table 4. Logistic Regression Estimates of Equation 1 (Output Giving Rise to Figure 2)

	ASIAN/ PACIFIC	BLACK	WHITE	LATINO	AFRICAN, ASIAN, HISPANIC ORIGIN	LGB	PROTEST- ANT	CATH- OLIC
Identity, Wave 1	9.974*** [0.851]	7.593*** [0.310]	5.240*** [0.276]	8.633*** [0.552]	4.423*** [0.214]	6.222*** [0.480]	3.954*** [0.120]	5.715*** [0.193]
Democrat ID strength, Wave 1	0.304 [1.167]	0.966 [0.522]	-0.566* [0.237]	1.013 [0.867]	0.753** [0.244]	-0.128 [0.508]	0.094 [0.204]	-0.094 [0.322]
Republican ID strength, Wave 1	0.796 [0.966]	-1.429* [0.546]	0.200 [0.385]	-0.936 [0.683]	-0.513 [0.304]	-1.192 [0.746]	0.290 [0.214]	0.288 [0.359]
Liberalism, Wave 1	0.222 [0.926]	-0.198 [0.638]	0.379 [0.350]	-0.723 [0.739]	0.2 [0.338]	0.746 [0.431]	-0.426 [0.225]	-0.275 [0.319]
Conservatism, Wave 1	1.442 [1.324]	0.618 [0.789]	-0.279 [0.413]	-1.799** [0.657]	0.096 [0.461]	-0.728 [0.926]	0.429 [0.225]	-0.207 [0.391]
Education, Wave 1	0.202** [0.076]	-0.061 [0.050]	-0.033 [0.025]	-0.162*** [0.035]	-0.075** [0.024]	-0.175* [0.082]	-0.037 [0.020]	-0.027 [0.036]
Age, Wave 1	0.009 [0.016]	-0.017* [0.008]	0.006 [0.005]	-0.016 [0.011]	-0.018** [0.006]	-0.009 [0.013]	0.015*** [0.004]	0.010 [0.007]
Sex	0.677 [0.367]	0.190 [0.291]	0.041 [0.220]	0.356 [0.350]	-0.345** [0.128]	0.201 [0.364]	0.174 [0.121]	0.037 [0.186]
Wave 1 conducted in 2008	1.738* [0.675]	0.940* [0.413]	-0.667* [0.255]	-0.793 [0.449]	-0.062 [0.210]		0.145 [0.138]	0.037 [0.224]
Wave 1 conducted in 2010	1.608* [0.660]	1.101* [0.493]	-0.802** [0.238]	-1.717** [0.545]	-0.100 [0.193]	0.176 [0.384]	0.052 [0.126]	-0.248 [0.186]
Intercept	-13.124*** [2.208]	-4.598*** [1.112]	-0.885 [0.622]	-1.732 [1.035]	-0.903 [0.539]	-2.057 [1.311]	-2.720*** [0.429]	-3.878*** [0.676]

N = 3,856 except for LGB, where N = 2,235. \*  $p < 0.05$ , \*\*  $p < 0.01$ , \*\*\*  $p < 0.001$  (two-tailed tests)



**SI Table 4. Logistic Regression Estimates of Equation 1 (Output Giving Rise to Figure 2) (continued)**

	JEWISH	NONE	BORN- AGAIN CHRIS- TIAN	LOWER CLASS	WORK- ING CLASS	MIDDLE CLASS	UPPER CLASS
Identity, Wave 1	8.871*** [0.873]	3.333*** [0.129]	3.236*** [0.102]	2.528*** [0.159]	1.924*** [0.084]	1.975*** [0.111]	3.490*** [0.303]
Democrat ID strength, Wave 1	1.427 [0.895]	-0.759*** [0.150]	0.185 [0.163]	-0.377 [0.221]	0.281* [0.132]	-0.173 [0.142]	0.008 [0.413]
Republican ID strength, Wave 1	1.476 [0.827]	-1.337*** [0.280]	-0.161 [0.174]	-0.829** [0.311]	-0.138 [0.153]	0.336* [0.162]	0.193 [0.476]
Liberalism, Wave 1	0.573 [0.624]	0.907*** [0.222]	-0.483* [0.205]	0.201 [0.255]	-0.699*** [0.187]	0.667*** [0.178]	0.168 [0.587]
Conservatism, Wave 1	-1.535 [2.097]	-0.239 [0.266]	0.818*** [0.179]	-0.032 [0.346]	-0.256 [0.198]	0.294 [0.178]	0.057 [0.579]
Education, Wave 1	0.134** [0.045]	0.077*** [0.021]	-0.105*** [0.021]	-0.161*** [0.027]	-0.132*** [0.018]	0.179*** [0.021]	0.146*** [0.041]
Age, Wave 1	-0.035* [0.017]	-0.020*** [0.004]	0.007* [0.003]	-0.008* [0.004]	-0.018*** [0.003]	0.018*** [0.003]	0.017* [0.007]
Sex	0.752 [1.028]	-0.616*** [0.116]	0.387*** [0.101]	0.223 [0.188]	-0.047 [0.104]	-0.028 [0.093]	0.073 [0.268]
Wave 1 conducted in 2008	1.053 [0.645]	0.031 [0.184]	0.174 [0.133]	0.240 [0.186]	-0.014 [0.104]	0.004 [0.108]	-0.413 [0.290]
Wave 1 conducted in 2010	0.426 [0.518]	0.448* [0.176]	-0.014 [0.129]	0.188 [0.193]	-0.073 [0.112]	0.044 [0.116]	0.113 [0.334]
Intercept	-9.325*** [1.935]	-1.634*** [0.397]	-1.451*** [0.418]	-0.705 [0.533]	1.647*** [0.306]	-4.560*** [0.368]	-7.175*** [0.860]

$N = 3,856$  except for LGB, where  $N = 2,235$ . \*  $p < 0.05$ , \*\*  $p < 0.01$ , \*\*\*  $p < 0.001$  (two-tailed tests)

**SI Table 5. Robustness of results to alternative specifications and coding decisions**

Identity	(1) Baseline specification (Estimates in Fig 2)	(2) PID only as predictor	(3) Controls for add'l background characteristics	(4) Code DKs as missing	(5) Two- year shifts (stacked dataset)	(6) Limit to non- Hispanic whites	(7) Limit to people of color
ASIAN/PACIFIC	.00 [.00]	.00 [.00]	.01 [.00]	.00* [.00]	.00 [.00]	. .	. .
BLACK	-.02 [.01]	-.02* [.01]	-.02 [.01]	-.01 [.01]	-.01* [.00]	. .	. .
WHITE	.00 [.01]	.02 [.01]	.00 [.01]	.00 [.01]	.05* [.01]	. .	. .
LATINO	-.02* [.01]	-.02* [.01]	-.02* [.01]	-.02* [.01]	.00 [.00]	. .	. .
AFRICAN, ASIAN, HISPANIC ORIGIN	-.06* [.02]	-.06* [.02]	-.05* [.02]	-.04* [.02]	-.05* [.01]	.00 [.00]	-.07 [.06]
LGB	-.02* [.01]	-.01* [.01]	-.03* [.01]	-.02* [.01]	-.03* [.01]	-.02* [.01]	.00 [.02]
PROTESTANT	.07* [.02]	.04* [.01]	.09* [.02]	.07* [.02]	.10* [.01]	.09* [.02]	-.01 [.04]
CATHOLIC	.01 [.01]	.01 [.01]	.02 [.01]	.01 [.01]	.00 [.01]	.02 [.01]	.04 [.03]
JEWISH	-.01 [.00]	.00 [.00]	.00 [.00]	-.01* [.00]	.00 [.00]	.00 [.01]	-.01 [.01]
NONE	-.10* [.02]	-.05* [.01]	-.10* [.02]	-.10* [.02]	-.11* [.01]	-.15* [.02]	-.02 [.04]
BORN-AGAIN CHRISTIAN	.08* [.02]	.01 [.01]	.07* [.02]	.08* [.02]	.10* [.01]	.10* [.02]	.09 [.06]
LOWER CLASS	-.03* [.01]	-.02* [.01]	-.02* [.01]	-.02* [.01]	-.03* [.01]	-.02 [.01]	-.02 [.04]
WORKING CLASS	.00 [.02]	-.04* [.02]	.00 [.03]	.00 [.02]	-.01 [.01]	.00 [.03]	.00 [.07]
MIDDLE CLASS	.02 [.02]	.05* [.02]	.03 [.03]	.02 [.02]	.03* [.01]	.01 [.03]	.03 [.06]
UPPER CLASS	.00 [.01]	.00 [.01]	.00 [.01]	.00 [.01]	.00 [.01]	.01 [.01]	-.02 [.02]
REGRESSION OF IDENTITY SHIFT DIFFS ON DISSIMILARITY INDEX SCORES	.12* [.03]	.07* [.03]	.12* [.04]	.12* [.03]	.16* [.02]	.18* [.06]	.10* [.04]

Cells display estimated differences in predicted identity shifts between conservative Republicans and liberal Democrats with robust standard errors. \* $p < .05$  (two-tailed tests). Notes to table on following page.

## Notes to SI Table 5. Robustness of results to alternative specifications and coding decisions

SI Table 6 displays analyses exploring the robustness of the main results of this paper to alternative model specifications and data coding decisions:

**Column 1** presents the baseline results as shown in the text for comparative purposes. These are the estimates for each identity as displayed on the left-hand side of Figure 2 (the differences between the predicted identity shifts of conservative Republicans and liberal Democrats as estimated by Equation 1). The final row displays the estimated slope of a bivariate regression of these shifts on the identity dissimilarity index scores (as plotted in Figure 3a and reported in the text on p. 22).

**Column 2** presents results yielded by re-estimating Equation 1 without any predictors for ideology, leaving partisanship as the only predictor of politicized identity shifts. The patterns of identity shifts and their relationship to political prototypes are very similar, although of a smaller magnitude, compared to those presented in Column 1.

**Column 3** presents results yielded by re-estimating Equation 1 with additional controls for the following background characteristics (as reported by panelists in Wave 1) that are causally prior to the process of politicized identity shifting: mother's education (GSS mnemonic: MADEG), type of place of residence at age 16 (RES16), region of residence at age 16 (REG16), fundamentalism of religion at age 16 (FUND16), number of siblings (SIBS), place of nativity of parents (PARBORN), and whether the panelist with living with his or her parents at age 16 (FAMILY16). The patterns of identity shifts and their relationship to political prototypes are very similar compared to Column 1.

**Column 4** presents results yielded by re-estimating Equation 1 with a dataset where responses of "don't know" or "refused" to identity questions are coded as missing data rather than as zeroes. The patterns of identity shifts and their relationship to political prototypes are very similar compared to Column 1.

**Column 5** presents results yielded by estimating Equation 1 to assess two-year, rather than four-year shifts in identity. The dataset was stacked so that there were two observations for each panelist: Wave 1-to-Wave 2 identity shifts and Wave 2-to-Wave 3 shifts. Equation 1 was then estimated with two-year lagged values of all controls as predictors using random-effects GLS with standard errors clustered by respondent. The patterns of identity shifts and their relationship to political prototypes are very similar compared to Column 1.

**Column 6** and **Column 7** respectively present results yielded by estimating Equation 1 separately for those consistently identifying as either non-Hispanic white or as a person of color (that is, those not identifying as white plus those identifying as Hispanic regardless of race) in all three waves. With the caveat that these are smaller sample sizes and thus less precise estimates, a comparison of the two columns suggests that politicized identity shifting with regard to religion and sexuality is more prevalent among whites, while shifting with regard to national origin is more prevalent among people of color. The final row of both columns confirms that identity shifts are significantly predicted by political prototypes among both whites and people of color.

SI Table 6a. Placebo Tests

	predictive margins for...			
	...Wave 1	...Wave 1		
	liberal	conserv		
	Democrats	Republicans	difference	<i>p</i> -value of difference
<b>Astrological sign</b>				
Aries	0.082	0.081	-0.001	0.781
Taurus	0.080	0.080	0.000	0.958
Gemini	0.087	0.085	-0.001	0.712
Cancer	0.085	0.087	0.002	0.689
Leo	0.077	0.083	0.006	0.224
Virgo	0.097	0.094	-0.002	0.515
Libra	0.085	0.090	0.005	0.088
Scorpio	0.075	0.073	-0.002	0.533
Sagittarius	0.077	0.073	-0.004	0.327
Capricorn	0.082	0.079	-0.003	0.385
Aquarius	0.088	0.088	-0.001	0.859
Pisces	0.086	0.088	0.001	0.286
<b>Region of residence at age 16</b>				
Not in U.S.	0.085	0.078	-0.006	0.275
New England	0.040	0.039	0.000	0.842
Middle Atlantic	0.158	0.148	-0.010	0.142
East North Central	0.187	0.189	0.002	0.658
West North Central	0.068	0.062	-0.005	0.067
South Atlantic	0.153	0.159	0.006	0.495
East South Central	0.052	0.058	0.006	0.249
West South Central	0.082	0.084	0.002	0.650
Mountain	0.044	0.045	0.002	0.538
Pacific	0.131	0.133	0.003	0.551
<b>Additional variables</b>				
sex (1 male, 2 female)	1.55	1.55	-0.004	0.494
year of birth	1962.5	1962.2	-0.311	0.049
mother's education (years)	11.86	11.78	-0.082	0.428
father's education (years)	11.68	11.73	0.054	0.645
number of siblings	3.55	3.45	-0.106	0.176

### SI Table 6b. Sensitivity of Results in Figure 2 to Reliability of Measurement

The results reported in Figure 2 are generated from the regressions reported in SI Table 5, which assume all variables in the estimations are measured without error—that is that their reliability, or rho, equals 1. This table reports the minimum level of rho ( $\rho_{\min}$ ) for each identity measure required for significant results reported in Figure 2 to hold. These minimum levels were determined by successive runs of errors-in-variables estimates of Equation 1 in which the rho for each identity measure was iterated between .50 and .99 in steps of .01. Identities in **bold type** are those exhibiting significant politicized identity shifts as reported in Figure 2.

This process also revealed that for some identities, non-significant results in Figure 2 become significant at levels of rho just slightly lower than 1 in theoretically expected directions; the maximum level of rho ( $\rho_{\max}$ ) at which significant results are obtained are also reported here.

Identities for which significant results were not achieved at any level of reliability are left blank.

identity	$\rho_{\min}$	$\rho_{\max}$
ASIAN/PACIFIC		0.90
BLACK		0.91
WHITE		0.93
<b>LATINO</b>	<b>0.90</b>	
<b>AFRICAN, ASIAN, OR HISPANIC ORIGIN</b>	<b>0.86</b>	
<b>LESBIAN, GAY, BISEXUAL</b>	<b>0.93</b>	
<b>PROTESTANT</b>	<b>0.83</b>	
CATHOLIC		0.78
JEWISH		
<b>NO RELIGION</b>	<b>0.69</b>	
<b>BORN-AGAIN CHRISTIAN</b>	<b>0.70</b>	
<b>LOWER CLASS</b>	<b>1</b>	
WORKING CLASS		0.47
MIDDLE CLASS		
UPPER CLASS		

SI Table 7. Regression Estimates Giving Rise to Table 2

	Wave 1 to Wave 3 identity shifts		$t - 1$ to $t$ identity shifts	
	logit	OLS	logit	OLS
lagged identification as a $j$	3.801*** [0.042]	0.690*** [0.005]	3.943*** [0.033]	0.710*** [0.004]
dissimilarity index, $j$	3.797 [6.839]	0.09 [0.139]	-1.237 [4.860]	0.032 [0.081]
lagged liberalism	-0.014 [0.040]	-0.007 [0.004]	0.007 [0.025]	-0.003 [0.002]
lagged conservatism	0.077* [0.038]	0.012*** [0.003]	0.095*** [0.027]	0.012*** [0.002]
lagged Democrat	0.033 [0.032]	0.000 [0.003]	0.090*** [0.019]	0.005** [0.002]
lagged Republican	-0.179*** [0.037]	0.001 [0.003]	-0.122*** [0.025]	0.007*** [0.002]
dissimilarity index, $j$ x...				
...lagged conservatism	0.759** [0.283]	0.056** [0.019]	0.544** [0.200]	0.043** [0.013]
...lagged Republican	1.755*** [0.293]	0.073*** [0.018]	2.436*** [0.202]	0.112*** [0.011]
...lagged liberalism	-0.691** [0.236]	-0.057** [0.020]	-0.401* [0.160]	-0.035** [0.013]
...lagged Democrat	-0.764*** [0.194]	-0.050** [0.016]	-0.554*** [0.133]	-0.032** [0.011]
age	-0.001* [0.001]	0.000 [0.000]	-0.002*** [0.000]	-0.000*** [0.000]
education	-0.020*** [0.003]	-0.002*** [0.000]	-0.022*** [0.002]	-0.002*** [0.000]
sex	0.03 [0.019]	0.003* [0.001]	0.015 [0.012]	0.001 [0.001]
Wave 1 in 2008	0.005 [0.024]	0.000 [0.002]		
Wave 1 in 2010	-0.003 [0.023]	0.000 [0.002]		
year of survey: 2010			-0.002 [0.020]	0 [0.001]
year of survey: 2012			0.03 [0.020]	0.002 [0.001]
year of survey: 2014			-0.024 [0.022]	-0.002 [0.001]
Intercept	-3.423*** [0.725]	0.036* [0.015]	-3.899*** [0.527]	0.033*** [0.009]
N	3,856		4,637	
$p$ , joint significance of shaded coefficients above	<.001	<.001	<.001	<.001

\*  $p < 0.05$ , \*\*  $p < 0.01$ , \*\*\*  $p < 0.001$  (two-tailed tests).

**SI Table 7. Regression Estimates Giving Rise to Table 2 (continued)**

	model accounting for selection	
	logit	OLS
liberalism	-0.067*** [0.018]	-0.011*** [0.003]
conservatism	0.036 [0.019]	0.017*** [0.003]
Democrat	0.132*** [0.018]	0.019*** [0.003]
Republican	-0.116*** [0.023]	0.017*** [0.003]
dissimilarity index, $j$	-22.962 [123.042]	-3.61 [16.657]
dissimilarity index, $j$ x...		
...conservatism	0.926*** [0.198]	0.115*** [0.025]
...Republican	2.282*** [0.242]	0.208*** [0.027]
...liberalism	0.094 [0.171]	-0.011 [0.027]
...Democrat	-0.263 [0.166]	-0.037 [0.026]
age	-0.002*** [0.000]	-0.000*** [0.000]
education	-0.030*** [0.002]	-0.004*** [0.000]
sex	0.040*** [0.011]	0.005*** [0.002]
year of survey	0.002 [0.006]	0.000 [0.001]
ever Democrat	-4.399 [10.637]	-0.616 [1.807]
ever Republican	-0.291 [10.593]	0.930 [1.826]
ever liberal	-1.107 [8.972]	0.454 [1.525]
ever conservative	0.512 [8.690]	-0.439 [1.510]
controls for (year x "ever..." variables), (dissimilarity index x "ever..." variables) and (year x dissimilarity index x "ever..." variables)	X	X
N	3,872	
$p$ , joint significance of shaded coefficients above	<.001	<.001

\*  $p < 0.05$ , \*\*  $p < 0.01$ , \*\*\*  $p < 0.001$  (two-tailed tests).

**SI Table 8. Regression Estimates Giving Rise to Table 3**

Wave 1 to Wave 3 identity shifts (columns 1 and 2 in Table 3)

panelist party ID and ideology:	OLS		logit	
	consistent	switched	consistent	switched
lagged identification as a <i>j</i>	0.70*	0.68*	3.89*	3.64*
	[0.01]	[0.01]	[0.05]	[0.09]
dissimilarity index, <i>j</i>	0.04	0.15	1.29	0.32
	[0.16]	[0.31]	[7.82]	[17.80]
lagged liberalism	-0.01*	-0.01	-0.05	0.04
	[0.00]	[0.01]	[0.05]	[0.09]
lagged conservatism	0.02*	0.01	0.08	0.12
	[0.00]	[0.01]	[0.04]	[0.09]
lagged Democrat	0.00	-0.01	0.06	-0.08
	[0.00]	[0.01]	[0.04]	[0.08]
lagged Republican	0.00	-0.01	-0.22*	-0.21*
	[0.00]	[0.01]	[0.04]	[0.09]
dissimilarity index, <i>j</i> x...				
...lagged conservatism	0.08*	-0.04	1.33*	-0.69
	[0.02]	[0.05]	[0.34]	[0.55]
...lagged Republican	0.07*	0.00	2.14*	0.11
	[0.02]	[0.05]	[0.35]	[0.58]
...lagged liberalism	-0.04	-0.14*	-0.39	-1.71*
	[0.02]	[0.05]	[0.28]	[0.53]
...lagged Democrat	-0.04*	-0.02	-0.78*	-0.29
	[0.02]	[0.04]	[0.23]	[0.43]
age	-0.00*	0.00	-0.00*	0.00
	[0.00]	[0.00]	[0.00]	[0.00]
education	-0.00*	-0.00*	-0.01*	-0.03*
	[0.00]	[0.00]	[0.00]	[0.01]
sex	0.00*	0.00	0.04	0.02
	[0.00]	[0.00]	[0.02]	[0.05]
year of Wave 1 of survey	0.00	0.00	0.00	-0.01
	[0.00]	[0.00]	[0.01]	[0.01]
Intercept	0.18	1.91	-0.4	15.46
	[0.85]	[2.27]	[13.15]	[28.42]
panelist <i>N</i>	2,920	659	2,920	659

\*  $p < 0.05$  (two-tailed tests).



**SI Table 9. Regression Estimates Giving Rise to Table 3 (continued)**

Wave  $t-1$  to Wave  $t$  identity shifts (columns 3 and 4 in Table 3)

panelist party ID and ideology:	OLS		logit	
	consistent	switched	consistent	switched
lagged identification as a $j$	0.72*	0.68*	3.66*	4.06*
	[0.00]	[0.01]	[0.09]	[0.04]
dissimilarity index, $j$	-0.01	0.08	-1.61	-2.93
	[0.09]	[0.21]	[13.98]	[5.88]
lagged liberalism	-0.01*	0.00	0.13*	-0.08*
	[0.00]	[0.01]	[0.06]	[0.03]
lagged conservatism	0.01*	0.00	0.09	0.09*
	[0.00]	[0.01]	[0.07]	[0.03]
lagged Democrat	0.01*	0.01	0.10*	0.11*
	[0.00]	[0.01]	[0.05]	[0.02]
lagged Republican	0.00*	0.01	0.02	-0.18*
	[0.00]	[0.01]	[0.06]	[0.03]
dissimilarity index, $j$ x...				
...lagged conservatism	0.07*	-0.06	-0.93*	1.08*
	[0.01]	[0.04]	[0.44]	[0.25]
...lagged Republican	0.09*	0.09*	1.16*	2.51*
	[0.01]	[0.04]	[0.52]	[0.25]
...lagged liberalism	-0.04*	-0.10*	-1.22*	-0.32
	[0.01]	[0.04]	[0.41]	[0.19]
...lagged Democrat	-0.03*	-0.03	-0.41	-0.60*
	[0.01]	[0.03]	[0.34]	[0.17]
age	-0.00*	0.00	0.00	-0.00*
	[0.00]	[0.00]	[0.00]	[0.00]
education	-0.00*	-0.00*	-0.03*	-0.01*
	[0.00]	[0.00]	[0.01]	[0.00]
sex	0.00*	0.00	0.00	0.02
	[0.00]	[0.00]	[0.03]	[0.01]
year of survey: 2010	0.00	0.00	-0.01	0.01
	[0.00]	[0.00]	[0.05]	[0.02]
year of survey: 2012	0.00	0.00	0.03	0.03
	[0.00]	[0.00]	[0.05]	[0.02]
year of survey: 2014	0.00	0.00	-0.04	0.00
	[0.00]	[0.00]	[0.06]	[0.03]
Intercept	0.02*	0.04	-4.05*	-4.21*
	[0.01]	[0.02]	[1.53]	[0.64]
panelist $N$	2,920	659	2,920	659

\*  $p < 0.05$  (two-tailed tests).

**SI Table 10. Regression Estimates Giving Rise to Table 3 (continued)**  
Model accounting for selection (columns 5 and 6 of Table 3)

panelist party ID and ideology:	OLS		logit	
	consistent	switched	consistent	switched
liberalism	-0.02*	0.00	-0.11*	0.04
	[0.00]	[0.01]	[0.02]	[0.04]
conservatism	0.03*	0.00	0.06*	0.03
	[0.00]	[0.01]	[0.02]	[0.04]
Democrat	0.02*	0.02*	0.15*	0.13*
	[0.00]	[0.01]	[0.02]	[0.04]
Republican	0.01*	0.03*	-0.18*	0.09
	[0.00]	[0.01]	[0.03]	[0.05]
dissimilarity index, <i>j</i>	-0.02*	0.00	-0.11*	0.04
	[0.00]	[0.01]	[0.02]	[0.04]
dissimilarity index, <i>j</i> x...				
...conservatism	0.15*	-0.11	1.33*	-0.75
	[0.03]	[0.06]	[0.27]	[0.39]
...Republican	0.18*	0.17*	2.40*	1.13*
	[0.03]	[0.07]	[0.33]	[0.47]
...liberalism	-0.04	-0.15*	-0.05	-0.97*
	[0.03]	[0.06]	[0.22]	[0.36]
...Democrat	-0.08*	0.02	-0.63*	0.07
	[0.03]	[0.06]	[0.21]	[0.38]
age	-0.00*	-0.00*	-0.00*	-0.00*
	[0.00]	[0.00]	[0.00]	[0.00]
education	-0.00*	-0.01*	-0.03*	-0.04*
	[0.00]	[0.00]	[0.00]	[0.00]
sex	0.01*	0.00	0.04*	0.02
	[0.00]	[0.00]	[0.01]	[0.03]
year of survey	0.00	-0.01	-0.01	-0.03
	[0.00]	[0.01]	[0.01]	[0.03]
ever Democrat	-3.15	-7.98	-12.86	-31.24
	[2.73]	[5.06]	[16.73]	[25.46]
ever Republican	-0.76	-5.86	-1.68	-35.48
	[2.79]	[4.92]	[17.45]	[23.59]
ever liberal	0.70	-8.46	-11.93	-26.69
	[1.95]	[5.01]	[12.37]	[26.94]
ever conservative	-1.85	-6.40	-21.67	-5.90
	[1.96]	[5.21]	[12.56]	[27.06]
controls for (year x "ever..." variables), (dissimilarity index x "ever..." variables) and (year x dissimilarity index x "ever..." variables)	X	X	X	X
panelist <i>N</i>	2,929	661	2,929	661

\*  $p < 0.05$  (two-tailed tests).