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**Working Paper**

# The Surprising Stability of Asian Americans' and Latinos' Partisan Identities in the Early Trump Era

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March 4, 2022

*The authors declare no competing interests and gratefully acknowledge the support of the Russell Sage Foundation (Awards 87-14-05 and 93-17-02 to Hopkins, Kaiser, and Pérez). This research was reviewed and approved by Institutional Review Boards at Georgetown University (2014-1003), the University of Pennsylvania (823994), the University of Washington (48286), and Vanderbilt University (151979). The experimental analyses were pre-registered via EGAP (https://osf.io/ca96e). We thank Emma Arsekin, David Azizi, Nicholas Fernandez, Tori Gorton, Adelaide Lyall, Carlos Montes, Georgia Ray, Elizabeth Sanchez, and Janelle Schneider for excellent research assistance and Drew Engelhardt, Zoltan Hajnal, Jae Yeon Kim, Jennifer Lee, Michele Margolis, Matt Nelsen, and Tanika Raychaudhuri for incisive comments. Versions of this research were presented at the Annual Meeting of the Midwest Political Science Association (April 14, 2021), PRIEC (May 7, 2021), and the International Society of Political Psychology (July 12, 2021). There are no conflicts of interest or competing interests to report.

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## Abstract

Two prominent, compatible accounts contend that Asian Americans and Latinos are not strongly connected to America's political parties and that their partisanship is responsive to identity threats. Donald Trump's political ascent presents a critical test, as Trump reoriented the Republican Party by foregrounding anti-immigrant hostility. Here, we test these perspectives using one of the first-ever population-based panels of Asian Americans and Latinos fielded 2016 to 2018. Across various empirical tests, we uncover surprising strength and stability in respondents' partisan identities. In a period of pronounced anti-immigrant rhetoric, these groups remained steadfast in their party affiliation. We also show that pan-ethnic identities were stable over this period and that partisanship can predict subsequent pan-ethnic identities more consistently than the reverse. By 2016, pan-ethnic identities were already stably integrated with partisanship, with little evidence of situational shifts in response to identity threats.

*"When Mexico sends its people, they're not sending their best... They're sending people that have lots of problems, and they're bringing those problems with us. They're bringing drugs. They're bringing crime. They're rapists. And some, I assume, are good people."*

—Donald Trump, June 16, 2015

Both as a candidate and then president, Donald Trump wore his hostility to immigrants on his sleeve. Although his xenophobic rhetoric was unmistakable, its consequences for an already-polarized mass public are less certain (Levendusky, 2009; Mason, 2018). In particular, the effects of Trump's rhetoric on the partisan loyalties of Asian American and Latino adults are not yet fully understood. Nonetheless, Trump's rhetoric would seem to make his candidacy and presidency uniquely threatening to these groups, meaning that his political ascendance poses a critical test for prominent theories of inter-group politics.[1]

Two such theoretical approaches together yield the expectation that Trump's xenophobic rhetoric may have strengthened Democratic allegiances among these two fast-growing, heavily immigrant pan-ethnic groups. The first emphasizes that many Asian Americans and Latinos are born outside the U.S., or else have parents born outside the U.S. In fact, 59% of Asian Americans and 34% of Latinos were born abroad (Flores, 2017; López, Ruiz and Patten, 2017). This complicates the standard "parental socialization" model of partisanship in which U.S.-born parents transmit their partisan attachments to their U.S.-born children (Campbell et al. 1960; Green, Palmquist and Schickler 2002; Jennings, Stoker and Bowers 2009; see also Carlos 2021). These groups are also not subject to consistent partisan mobilization (Hajnal and Lee 2011; Garcia Bedolla and Michelson 2012; see also Fraga 2018). Certainly, Asian Americans and Latinos lean toward the Democratic Party on average (Cho, 1995; Alvarez and Garcia Bedolla, 2003; Lien, 2004; Abrajano and Alvarez, 2010; Barreto, 2010; Wong et al., 2011; Masuoka et al., 2018; Leung, 2021; Chan, Kim and Leung, 2021). Still, extensive scholarship contends that Asian Americans and Latinos have weaker partisan attachments than Black or White Americans (Hajnal and Lee, 2011; Wong et al., 2011; Wals, 2011, 2013; Masuoka et al., 2018; Raychaudhuri, 2020).

The second theoretical approach is often termed "identity threat." This approach, which occupies a special status in the study of minority politics, emphasizes the role of threats to valued social identities in

[1]The experiment detailed below was pre-registered at [URL]; the pre-analysis plan is in the Appendix.

reshaping partisanship. Once acquired, political partisanship is typically highly stable (Campbell et al., 1960; Green, Palmquist and Schickler, 2002). Thus, after a solid sense of partisanship forms, party identification changes significantly only under exceptional circumstances, as demonstrated by the migration of racially conservative White Democrats to the Republican party after 1970 (Valentino and Sears, 2005). To the extent that partisanship does shift, such changes are thought to be the product of evolving perceptions about the social groups associated with the two major parties (Green, Palmquist and Schickler, 2002; Achen and Bartels, 2017; Kuo, Malhotra and Mo, 2017; Ahler and Sood, 2018; White and Laird, 2020). Through his xenophobic rhetoric, Trump may have shifted the social group imagery of the two parties. Indeed, a related vein of research shows that identity threats can influence the partisanship and political behavior of Asian Americans and Latinos (see esp. Gutierrez et al. 2019; Garcia-Rios, Pedraza and Wilcox-Archuleta 2019; Wong et al. 2011; Kuo, Malhotra and Mo 2017; Chan, Kim and Leung 2021; see also Pantoja, Ramirez and Segura 2001; Hui and Sears 2018).

Together, these theoretical perspectives lay the groundwork for expecting that Trump's rhetoric moved Asian Americans and Latinos toward the Democrats. After all, sizable numbers of Asian American and Latino adults weren't firmly attached to the major parties *before* Trump's political emergence, making them available to be influenced by xenophobic rhetoric. And under the second perspective, despite partisanship's general stability, identity threats are one engine of change. In fact, Trump would seem to be a most-likely case for identity threat.

In contrast to these expectations is the stability hypothesis. In this view, Asian American and Latino adults hold partisan identities that are stronger than implied by some prior work yet not as reactive to Trump's immigration rhetoric as research on identity threats would suggest (Hopkins et al., 2020). Unlike the first two perspectives, the stability framework contends that many Asian American and Latino adults already possess underappreciated levels of partisan attachment, primarily (though not exclusively) to the Democratic Party. This perspective can be traced partly to scholarship demonstrating that high levels of clear and consistent information at key points in the life course can crystalize predispositions like partisanship, which can later affect behaviors (Sears and Valentino, 1997). In light of sustained political polarization across recent decades (e.g. Levendusky, 2009), this trend has clarified which social groups belong to which parties, fortifying the connection between people of color and the Democratic Party (Ahler and Sood, 2018). Thus, sustained polarization may affect the evolving partisanship of politically marginalized groups, such as Asian Americans and Latinos, who previously demonstrated lower levels of partisan crystallization. This would be consistent with scholarship on the cumulative impact of racialized politics on the political mobilization of communities of color (Pantoja, Ramirez and Segura, 2001; Bowler, Nicholson and Segura, 2006; Pedraza, 2014; Pérez, 2015*a*,*b*). In other words, rather than transforming Asian American and Latino adults into Democratic partisans, Trump's rhetoric may have affirmed or crystallized these groups' pre-existing Democratic loyalties, as indicated by their high and stable partisanship levels throughout his presidency (but see Valentino and Zhirkov, 2018; Chan, Kim and Leung, 2021). Instead of inducing shifts in partisanship, Trump's rhetoric may have reinforced the status quo.

The size and influence of these two growing pan-ethnic groups make it crucial to understand Asian American and Latino politics on their own terms. But studying Asian American and Latino partisanship between 2016 and 2018 also provides a high-leverage case to understand partisanship and the mechanisms that underpin its adoption and change more generally. This slice in time includes a watershed moment in which a highly visible politician (Trump) further shifted the image of the Republican party in a White nationalist direction, complete with unambiguous hostility to Asian Americans, Latinos, and other communities of color. The consistency and strength of this real-world stimulus provides a critical test of whether threatened racial/ethnic identities drive the adoption of partisan identities. This matters because there are ongoing debates about the extent to which partisan identities are rooted in social group attachments as opposed to issue positions, ideology, or other factors (Achen and Bartels, 2017; Mason, 2018; Barber and Pope, 2019; Fowler et al., 2020; White and Laird, 2020). According to theoretical accounts of partisanship foregrounding social group identities, the content of the issue may be less important than the perception of rhetorical attacks against already-meaningful in-groups. In that case, it may be those who strongly identify with the group whose partisanship is most likely to shift (Pérez, 2015*a*,*b*; Garcia-Rios, Pedraza and Wilcox-Archuleta, 2019; Hickel Jr. et al., 2021; White and Laird, 2020).

Still, individuals commonly have multiple identities within their repertoires (Chandra, 2006), and it is plausible that partisan identities and pan-ethnic attachments may not be hierarchically ordered so much as mutually constitutive. It may not always be the case that pan-ethnic attachments are adopted prior to partisan ones. For some, partisan attachments may be sufficiently strong as to influence attachment to pan-ethnic identities, a dynamic observed with other types of identities and predispositions (Margolis 2018; Engelhardt 2020; Egan 2020; Agadjanian and Lacy 2021). As a related point, identities are likely to be contextually activated, meaning that partisan identities could be potent within political contexts even while not being central to individuals' self-image elsewhere (Garcia-Rios, Pedraza and Wilcox-Archuleta, 2019; Pérez, 2020, 2021). Still, it is important to add that any of these mechanisms may operate on distinct subsets of Asian American or Latino residents—and that the explanation for partisan stability among some may differ from the explanation for partisan change among others.

Even if scholars agree that these competing perspectives are worth testing, appraising them has been stymied by methodological challenges. Asian American and Latino adults are some of the hardest-to-reach groups in the U.S. due to their geographic concentration, linguistic variation, heterogeneity in national origins, and other factors that limit scholars' ability to yield high-quality, representative samples (Barreto, Reny and Wilcox-Archuleta, 2017; Barreto et al., 2018). A related challenge involves political science's reliance on cross-sectional survey data to identify changes in partisanship among Asian American and Latino adults. While cross-sectional data can provide valuable information, testing the various theoretical perspectives requires data tracking the partisan identities of the same Asian American and Latino adults over time. Even for general populations, such panel data is rare; for Latinos, none existed prior to 2016 (see also McCann and Jones-Correa, 2020; Carlos, 2021), and we are not aware of population-based panels of Asian Americans. Certainly, concerns about causal inference loom even with panel data, justifying our inclusion of a survey experiment.

## Empirical Contribution

This paper presents a first effort to assemble data to test the theoretical perspectives and identity-oriented mechanisms previously outlined. In doing so, we provide evidence from what is among the first population-based panel surveys of Asian American and Latino respondents. This survey included 1,541 panelists recruited by GfK (later Ipsos) prior to spring 2016 using population-based sampling methods such as address-based sampling and random-digit dialing. By virtue of its sampling strategy, this panel is less prone to sample selection biases than some opt-in online surveys. The questionnaires were administered in Spanish as well as English. Panelists were initially interviewed in March/April 2016 and subsequently re-interviewed in fall 2016 (n=867) and fall 2018 (n=453). We provide extensive analyses of sample representativeness and attrition below.

By analyzing this three-wave panel, we are able to assess varied observable implications from these three theoretical perspectives. We track the same Asian Americans and Latinos throughout the Trump campaign and early presidency, finding remarkable stability in the expression of party identity and party-related attitudes and so affirming prior work on the nature of partisanship (Green, Palmquist and Schickler, 2002). (While partisanship is quite stable, measures of related variables such as candidate affect and vote choice show somewhat more variability.) The panel enables us to examine the relationship between partisan identity and pan-ethnic identities, and their shifts in the face of political events over this period. Descriptively, we find that our respondents' attachments to their pan-ethnic identities are quite stable. So, too, is the relationship between attachment to pan-ethnic identities and partisanship: Asian Americans and Latinos for whom pan-ethnic identities are important are consistently more likely to be Democrats, with only substantively small increases in that relationship between 2016 and 2018.

Our results also show that when asked to rank the importance of various identities, the respondents rank their pan-ethnic attachments as more important than their partisan ones. Yet contexts can make some identities more or less relevant, with even ostensibly weaker identities proving more influential in specific domains (Garcia-Rios, Pedraza and Wilcox-Archuleta, 2019; Pérez, 2020, 2021). Consistent with that logic, cross-lagged regression models demonstrate that partisan identities shaped subsequent attachments to pan-ethnic identities. Partisanship proves not just stable, but capable of affecting the expressions of other identities, thus extending prior work on how individuals' partisanship can influence the development of other identities (Margolis 2018; Egan 2020; Agadjanian and Lacy 2021).

To provide causal evidence, we also present a pre-registered survey experiment in which we showed some panelists randomly assigned clips of Trump making negative comments about Asian Americans or Latinos. Neither clip influenced vote choice in hypothetical general match-ups. However, the clip targeting Latinos did influence a few party-related attitudes, and thus provides some evidence that their party-related views can be moved with one-sided information flows. Still, in the context of other experimental findings (Hopkins et al., 2020), this overall pattern of results is consistent with general partisan stability among our respondents. That conclusion is reinforced through analyses of open-ended questions about perceptions of the parties included in Appendix A. We find a general pattern of stability in those responses, alongside evidence that ethnic/racial groups are not commonly mentioned when respondents describe the two parties in either 2016 or 2018. This finding further undercuts the identity threat hypothesis.

These results cover 32 months, and while that time period saw major events including Trump's nomination and election, they do not capture events before or after. It is quite possible that some of the impacts of the GOP's hostility to immigration were manifest before March 2016; it's likewise possible that some influential effects, especially the spike in anti-Asian rhetoric related to the COVID-19 pandemic, took place after November 2018 (Chan, Kim and Leung, 2021). However, as we show in Appendix Figure A2, the shifts in partisanship in the year prior to our first wave were limited, indicating that our results are not simply a product of pre-treatment effects. (Roughly speaking, the groups show generally stable macropartisanship in the decade prior to our panel.)

Overall, we interpret our results as calling for a re-appraisal of partisanship's nature among Asian Americans and Latinos, groups which are sometimes thought to be distinguished by weaker partisan attachments. Additionally, our results point to a greater need for synergy between the standard model of party identification and group-specific insights about Asian American and Latino adults, in an effort to develop a better model of the origins of partisanship among these and other minority groups (see also Pérez and Kuo, In press). In a polarized political moment, the political socialization of Asian Americans and Latinos may have more in common with that of other groups than we recognize.

## Theory and Hypotheses

We aim to test three theoretical perspectives about the nature of Asian Americans' and Latinos' partisanship. The first framework stipulates that Asian Americans and Latinos are not strongly partisan (Hajnal and Lee, 2011). The second is that both of these groups are highly responsive to threats to their pan-ethnic identities, with this response influencing their partisanship (Pantoja, Ramirez and Segura, 2001). Finally, a third claim is that partisanship is already solidified among both of these groups and that it remains resoundingly stable in the face of threats to one's pan-ethnic group (Huddy, Mason and Horwitz, 2016). We explain each perspective and yield specific hypotheses.

The first framework we test posits that Asian Americans and Latinos both possess weakly crystallized forms of partisanship, if they possess it at all. The foundation for this view is the standard parental-socialization model of party identification (Campbell et al., 1960). Here, young partisans come online and identify as either Democrats or Republicans based on their parents' partisan allegiance. This last point is crucial, for it presumes that parents have been thoroughly socialized into a specific partisan camp in the U.S. But for Asian American and Latino immigrants and their U.S.-born children, this implies negligible or incomplete socialization as partisans because they have had more limited exposure to American politics, are weakly motivated to engage in politics, and/or are sparsely contacted or recruited through partisan mobilization efforts (e.g. Abrajano and Alvarez, 2010; Wong et al., 2011; Ramirez, 2015). One hypothesis implied here is that insofar as Asian Americans and Latinos possess anemic levels of partisanship, we should observe over-time fluctuations in party identification within these groups, as they react to the latest available party-relevant information.

A second theoretical perspective suggests that in light of weak partisan identity levels among Asian

Americans and Latinos, pan-ethnic identities are the more politically relevant attachments. This contention aligns with multiple literatures that highlight the racial marginalization of these communities (Masuoka and Junn, 2013; Zou and Cheryan, 2017). According to this work, Asian Americans and Latinos should be highly responsive to identity threats–e.g., cues in elite rhetoric or public policy indicating hostility to their pan-ethnic group (Pérez, 2015*a*; Gutierrez et al., 2019; Chan, Kim and Leung, 2021). The paradigmatic example here involves Proposition 187 and its aftermath (Pantoja, Ramirez and Segura, 2001). In the wake of the campaign for that manifestly anti-Latino ballot measure in California, Latinos appear to have increased their long-run allegiance to Democrats. If this reasoning is correct, it implies an alternate hypothesis with a clear empirical signature. Inasmuch as identity threats are a main driver of Asian American and Latino politics, we should observe stable levels of pan-ethnic identity, which drive the development of stronger partisan attachments.

A final perspective anticipates steadfast levels of partisanship among Asian Americans and Latinos. Drawing on the partisan polarization literature (Levendusky, 2009; Ahler and Sood, 2018), this outlook sees the increasingly consistent positioning of Democrats and Republicans as providing Asian Americans and Latinos further clarity about which party is friendlier toward them (Huddy, Mason and Horwitz, 2016; Kuo, Malhotra and Mo, 2017; Chan, Kim and Leung, 2021). Since partisan polarization has overlapped substantially with the demographic growth of Asian Americans and Latinos, it is plausible that sufficient information about which parties are associated with various people of color has already reached them, reducing their informational deficits about both major parties (Garcia Bedolla and Michelson, 2012). Moreover, research on memory and attitude formation further suggests that individuals can encode patterns in information streams (e.g., party polarization), even if there is a delay in how well they can articulate new attitudes on the basis of that information (Pérez and Riddle, 2020). While other researchers have found low levels of partisan affiliation among Asian American and Latino adults (Hajnal and Lee, 2011), the perspective here suggests that polarization has solidified partisan preferences among many Asian Americans and Latinos. If correct, then we should observe clear and stable levels of partisan identification over time among members of these two major ethnic/racial groups.

## Data and Research Design

Our goal is to test three hypotheses—Asian Americans and Latinos are not strongly partisan, they are responsive to threats to their pan-ethnic identities, and their partisanship remains stable in the face of anti-immigrant rhetoric. Distinguishing between these perspectives demands an over-time look that allows one to evaluate whether, how much, and in what direction(s) partisanship and party-related attitudes have shifted in the Trump era. Moreover, to increase confidence in one of these perspectives, it is important to bore down into the possible psychological processes behind these shifts, which can be most directly assessed through carefully designed survey items and experiments. Here, we detail our data collection.

Partnering with the survey firm GfK (later Ipsos), we conducted a three-wave panel survey in English and Spanish using a population-based sample of Asian American and Latino adults residing in the U.S.[2] After pre-testing, the first wave was fielded between March 23 and April 11, 2016 and yielded 1,541 complete interviews (n=721 Asian American, 820 Latino). Fielded between October 20th and November 1st, 2016, the second wave re-sampled wave one respondents, yielding 867 complete interviews (n=415 Asian American, 452 Latino). Similarly, the third wave was administered between October 23 and November 5, 2018, and it re-interviewed 453 fall 2016 respondents (n=230 Asian American, 223 Latino). A wave prior to Trump's 2015 political ascendance would have been useful, but these waves enable us to track partisanship and opinions over a turbulent 2.5-year period.

Respondents were recruited offline using random-digit dialing or address-based sampling. This sampling strategy enables the recruitment of respondents who are not highly politically engaged. Appendix Tables A1 and A2 provide descriptive statistics for all three waves separately for Asian American and Latino respondents. For the Asian American sample, most summary statistics remain roughly similar across the three waves. For example, 84% of spring 2016 respondents and 85% of fall 2018 respondents are citizens. For that group, the main evidence of attrition is in voting behavior—the fraction of respondents who are not validated 2016 voters declined from 43% in the spring 2016 wave to 38% in fall 2016 and 35% in fall

[2] We were unable to field the survey to non-English-speaking Asian Americans.

2018. Put differently, low-turnout respondents are less likely to remain in the sample. However, for the Latino sample, there is evidence of systematic attrition across more variables. The fraction of respondents who were citizens in spring 2016 rises across the waves (from 74% to 79% and then 82%), as do average levels of education (from 12.1 to 13.2 years) and incomes (from $55K to $69K); the fraction replying in Spanish declines from 43% to 35% and then 26% while the fraction lacking a validated 2016 vote drops from 53% to 45% and then 39%. Given that political engagement may be correlated with attitude stability, we present robustness checks which reweight respondents to address this attrition below.

Another concern about representativeness relates to language. While Latinos could respond in Spanish or English, Asian Americans could respond only in English, meaning that we excluded from our sampling frame the roughly 28% of Asian Americans (Budiman and Ruiz, 2021) not proficient in English. To assess the impacts of this, Appendix Table A3 and the corresponding section present data from the 2016 post-election National Asian American Survey (NAAS) (Ramakrishnan et al., 2016), separating out respondents by language of survey. Appendix Table A3 shows that there are certainly differences by language, with respondents who took the survey in non-English languages much less likely to be born in the U.S. and having lower incomes and educational attainment. Notably, non-English respondents are also more likely to be in the center of the partisanship scale (3.6 vs. 2.9), meaning that the absence of non-English speaking Asian Americans may partially explain the differences between our conclusions and those of Hajnal and Lee (2011) for that pan-ethnic group. Even so, it's noteworthy that among some national-origin groups—including Indians, Filipinos, and Japanese—rates of English-language survey response are high, an observation which somewhat mitigates the impact of only interviewing Asian Americans in English.[3]

[3]By comparing these estimates with Appendix Table A1, we can also identify differences between this panel's Asian American respondents and the NAAS post-election respondents overall. While surveying in 11 languages, the NAAS has a *higher* fraction of Asian American respondents who were citizens than our panel (92% vs. 84%). Still, its respondents are less educated (13.6 years vs. 15.8), older (54 years vs. 45 years), and less likely to be born in the U.S. (25% vs. 47%).

## Panel Trends

We seek to track any shifts in Asian American or Latino partisanship during the volatile period from spring 2016 to fall 2018. At the same time, we track shifts in related attitudes, such as affect toward the political parties or Trump himself. Certainly, Trump had come to dominate headlines even prior to March 2016, so our panel won't capture the effects of his campaign's first eight months. In spring 2016, Trump had yet to clinch the GOP nomination and was far from becoming president. By fall 2018, he was the incumbent president, and in the weeks leading up to the 2018 midterm election, Trump tried to focus attention on a caravan of Central American migrants heading to the southern border. Thus, the prospects for observing identity threats in this period are real. Moreover, Appendix Figure A2 uses cross-sectional data from Pew Research Center to demonstrate that there were detectable but not large-scale shifts in Asian Americans' or Latinos' partisanship in 2015. Trump had not already induced major changes prior to our panel's first wave. These results also demonstrate that the distribution of partisanship for these groups was broadly similar in 2016 and in the preceding decade, meaning that the distribution of macropartisanship wasn't markedly different from that in the 2006-2008 period emphasized in Hajnal and Lee (2011).

Figure 1's left panel reports the mean level of partisanship, measured via the standard seven-category scale (Klar and Krupnikov, 2016). One thing to note is that respondents who initially did not identify with a major party were then asked a question about leaning which did not include an "independent" option, meaning that respondents were only classified as pure independents if they left the leaning question unanswered. As a consequence, 4% of wave one respondents were classified as pure independents.

Latinos responding in Spanish are the most heavily Democratic group, averaging between 2.38 and 2.65, which puts them between "weak Democrat" and "lean Democrat." English-responding Latinos vary between 3.07 and 3.38, while English-responding Asian Americans are similar (3.17-3.36). Still, what is most striking about Figure 1 is the aggregate over-time stability.

A related question is just how important partisan identities are to our respondents relative to other identities. The panel also asked respondents to rank different identities on the basis of their importance,

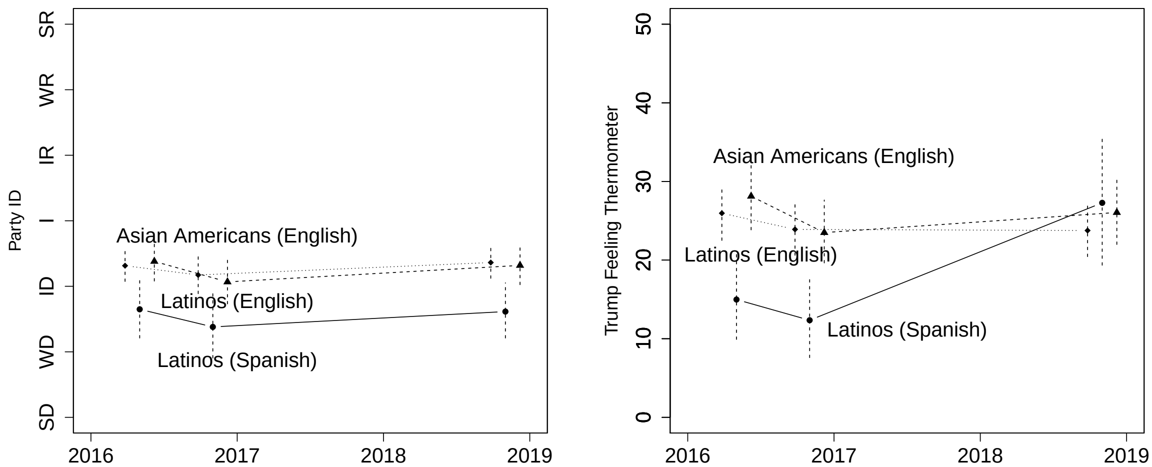


Figure 1: Left: over-time trends in partisanship for English-responding Latinos (n=163), Spanish-responding Latinos (n=57), and English-responding Asian Americans (n=229). Right: over-time trends for the same groups for their Trump feeling thermometer assessments.

including religion, job/occupation, gender, family role, political party, pan-ethnic group, national origin group, and being American. Appendix Table A4 reports the mean rankings in fall 2016 and fall 2018. While respondents typically report that their political party isn't central to their identity—its average rank out of 8 was 7.47 in fall 2016—partisan identities did become more important to respondents over those two years, reaching 6.43 ($p < 0.001$ from a two-sided t-test) in fall 2018. It's also noteworthy that pan-ethnic identities increased in rank as well, from 4.92 to 4.68 ($p = 0.017$). Over a period when partisan and pan-ethnic identities were increasingly important, aggregate partisanship was nonetheless stable.

Views of a political figure are likely to be more malleable than social identities, especially as people come to know that figure as the incumbent president. Figure 1's right side displays feeling thermometer assessments of Trump for the same groups of respondents. Here, too, there is reasonable stability, with English-responding Latinos only varying between thermometer scores of 26 and 28 while English-responding Asian Americans vary between 24 and 26. Spanish-responding Latinos (n=57) do show more variation, with a shift from 12 in fall 2016 to 27 two years later. Such variance may partly reflect the smaller sample size, although it proves robust.[4] Still, Trump assessments are somewhat more variable than partisanship, even while the broad story is of stability. We see no aggregate evidence that identity threats between 2016 and 2018 fostered Democratic partisanship.

## Partisan Transitions

One key advantage of panel data is that it enables researchers to examine patterns of individual-level change over time. Here, we consider the shifts in partisan identities among the 229 Asian American respondents who participated in the spring 2016 and fall 2018 waves. 161 (or 70%) placed themselves at precisely the same point on the seven-point scale both times, as Figure 2's top left panel illustrates. If we look at the fraction who moved by no more than one point—say, from "lean Democrat" to "pure independent"—it increases to 90%. The fraction identifying as some form of independent varies from 42-45%, but the fraction identifying as pure independents is just 2-4%. In short, such results show high stability, alongside sizable numbers of "closet partisans" who identify as independent but lean to a party. Such high levels of stability may partly reflect the English-dominant sample (see also Masuoka et al., 2018), but remain striking even for that subpopulation. (Appendix Figure A3 illustrates the results when reweighting respondents to account for attrition between spring 2016 and fall 2018, using the procedure detailed in the Appendix. The results are highly similar, with 70% of Asian American respondents estimated to have the same partisanship in both waves.)

The 220 Latino respondents include Spanish speakers as well as English speakers, and so may be less politically integrated and evince less stability. In fact, 124 respondents (56%) reported the same level of partisanship, a point illustrated in Figure 2's top right panel. The percentage does rise to 83% when we add those who shift by no more than one point; only 3% of the sample moves by four or more categories. In both waves, 40-41% of respondents identify as "independent," with just 5-6% identifying as pure independents.

One concern is that respondents whose partisanship is less stable may also be less likely to com-

[4] Specifically, we bootstrapped the median change between fall 2016 and 2018; its mean is 5.9 with a 95% confidence interval from 0 to 18.

plete future panel waves. Accordingly, Appendix Figure A3 illustrates over-time transitions after (again) reweighting observations to account for attrition between spring 2016 and fall 2018. The results are highly similar—the percentage of Latino respondents who report exactly the same partisanship remains 56%. That suggests that the stability is not an artifact of differential attrition.

In Appendix Table A5, we model changes in partisan stability for the full sample as a function of various demographics. That table reinforces the observation that the Latino respondents' partisanship is less stable than that of Asian American respondents. Beyond that, there are no other covariates that are strongly associated with partisan transitions.[5]

By way of benchmarking, we also estimated the same numbers for Black and White respondents using a separate, population-based panel also conducted by GfK (later Ipsos). Figure 2's bottom panels illustrate the results. For the 83 Black respondents participating in the January 2016 and October 2018 waves, 72% reported exactly the same partisanship, while for the 487 White respondents, the comparable figure is 71%. A remarkable 96% of Black respondents shifted no more than one level on the seven-point scale, while for White respondents, the figure is 90%.[6] Our Latino respondents do evince lower levels of stability than either English-speaking Asian American respondents or separate panels of Black and White respondents. But their levels of stability are nonetheless high in absolute terms. These findings thus depart from claims that Asian Americans and Latinos are not strongly attached to either political party.

# Vote Choice

We can also consider individual-level stability in vote choice, which is likely to be more malleable than partisan identity. Specifically, we consider whether respondents support the Democrat, the Republican, or neither in the 2016 presidential election and the 2018 Congressional election. For the 225 Asian American

[5]For example, respondents with higher levels of education are not more or less stable; respondents who identify as men are not more or less stable, either.

[6]The overall fraction of White respondents identifying as independent was 22% in 2016 and 13% in 2018—but the fractions of pure independents vary from 3-5%.

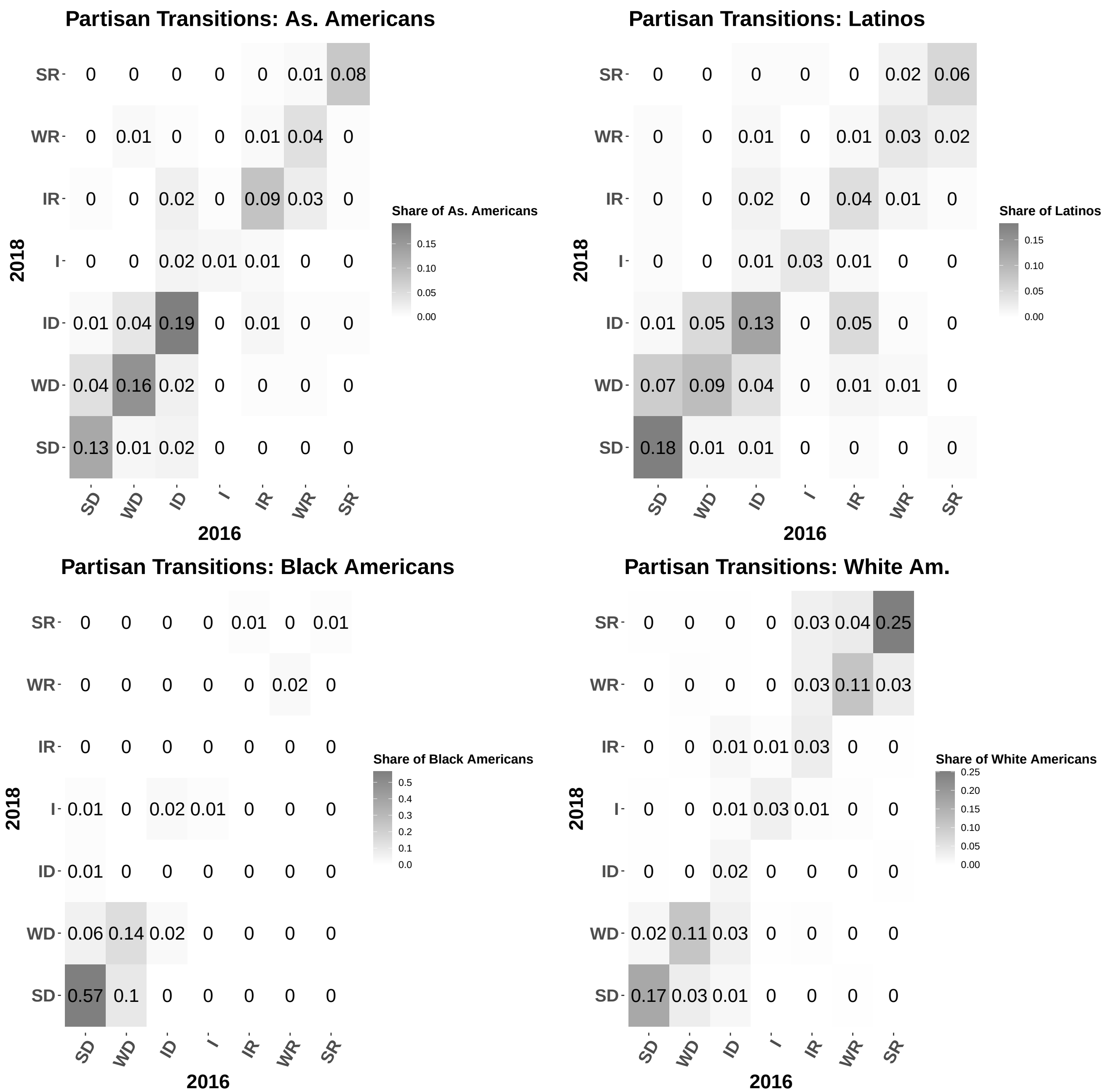


Figure 2: Transitions in partisan identification for the Asian American sample, the Latino sample, and separate samples of Black and White respondents also recruited via GfK's Knowledge Panel.

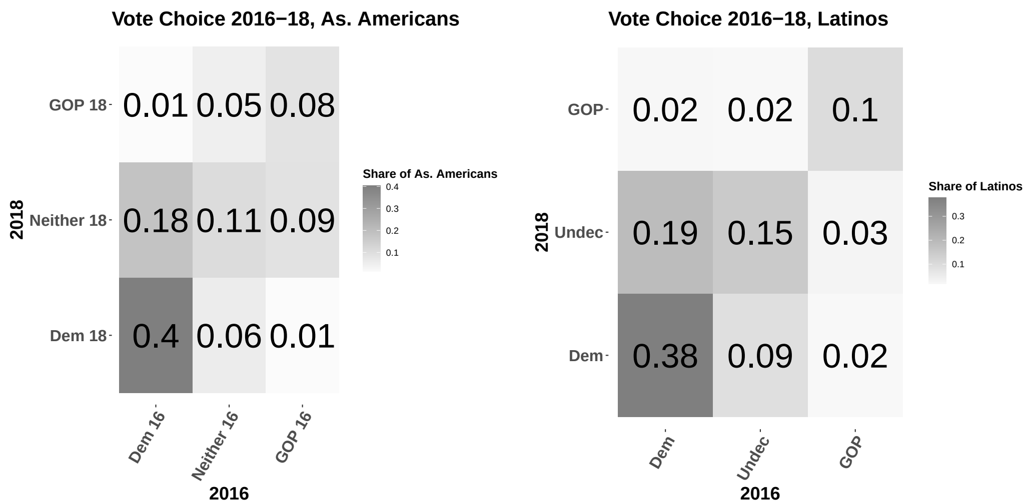


Figure 3: Transitions in vote intention between the 2016 presidential race (asked in spring 2016) and the fall 2018 Congressional race. Asian American respondents (n=225) are at left, while Latino respondents (n=220) are at right.

respondents observed in wave three, as illustrated in Figure 3's left panel, 60% gave the same reply in spring 2016 and fall 2018. Only 1% switched parties in either direction, leaving the remaining 38% to switch between being undecided and backing a major party. This low level of partisan defection is additional evidence of partisanship's relative stability. Among Asian Americans who were undecided in 2016 but not in 2018, partisanship averaged 4.0, meaning that this group was on average perfectly split; it was not simply anti-Trump Republicans. Meanwhile, those who became undecided were slightly more Democratic, with an average partisanship of 3.75.

For Latinos, the fraction who backed the same party in spring 2016 and fall 2018 is slightly higher, at 63%. Here, too, the fraction switching between parties is very low—4%—with the remainder moving into or out of the "undecided" camp. Among Latinos who were undecided only in 2016, average partisanship leaned slightly Democratic, at 3.8. That said, 19% of those who backed Democrat Hillary Clinton in spring 2016 were undecided in their fall 2018 House race, perhaps partly because the question format named the candidates in each district. On average, this group is more Democratic-leaning, with a mean partisanship of 3.2.

Here, too, we can also compare these numbers to the benchmark of Black and White respondents in the separate GfK-administered panel. Among that combined group, the percentage with the same party preference was a somewhat higher 69%, with again just 4% of people shifting from Republican to Democrat or vice versa. So while African Americans and Whites do show more stability in their partisan preferences, the difference with Asian Americans and Latinos isn't especially large—it's on the order of 6 to 9 percentage points. The stability isn't just at the aggregate level; it appears at the individual level, too.

## Predicting Over-time Shifts

Despite the general stability identified above, there are some Latinos and Asian Americans whose partisanship does change, making it imperative that we inquire about the correlates of such individual-level shifts. Following prior research (e.g. White and Laird, 2020), our focus is on social identities, and specifically on pan-ethnic and national-origin attachments. In the 2016-2018 period, were those with stronger pan-ethnic or national-origin identities more likely to change their partisanship, as the identity threat hypothesis suggests?

In these analyses, the key independent variables include indices of attachment to either pan-ethnic or national-origin identities varying from 2 to 8.[7] On average, our respondents placed themselves almost identically on these two indices in spring 2016 and fall 2018—the mean score on the pan-ethnic index in wave one was 4.90, and it remained virtually identical (4.91) in wave three.[8] The average levels are also unchanged when we look separately by pan-ethnic group (Appendix Figure A4). Asian Americans average 4.98 in spring 2016 and 5.04 in fall 2018; for Latinos, the scores are 4.80 and 4.77. The over-time stability of these identity attachments is itself evidence supporting the stability framework.[9]

[7]See Appendix F for question wording.

[8]Attachment to national origins is slightly higher, with respondents averaging 5.27 in wave one and then 5.25 in wave three. These two identity attachments are closely related, with a Pearson's correlation of 0.61 in wave three.

[9]Our subsequent analyses sometimes foreground pan-ethnic identities (e.g., Latino) rather than national origin identities (e.g., Mexican). While both forms of identity are socially relevant to Latinos and

We focus on pan-ethnic identity attachments, based partly on the insight that the elements of people's identity repertoires which are influential vary with the context. Pan-ethnic groups are to some degree explicitly political in their construction, as well as being a focus of contemporary political categorization/mobilization and the primary emphasis of recent research (Hajnal and Lee, 2011; Kuo, Malhotra and Mo, 2017; Kim, 2020).

With respect to issue attitudes, we focus on immigration, as it is a highly salient issue that is closely associated with Trump (Sides, Tesler and Vavreck, 2019) and has particular resonance among members of heavily immigrant pan-ethnic groups. An issue-based account might lead us to expect that respondents who are especially supportive of creating a pathway to citizenship are likely to become Democrats (see also Sanchez, Medeiros and Sanchez-Youngman, 2012) while those with the most restrictive immigration attitudes are likely to become Republicans.[10]

Table 1 provides descriptive assessments of these hypotheses via regression models. The dependent variable is the *difference* between spring 2016 and fall 2018 partisanship, measured via 7-category scales with 7 indicating strong Republicans. The covariates include basic variables that could confound our relationship of interest: education, gender, age, income, place of birth, citizenship, and indicators for large national-origin categories including Chinese, Filipino, Mexican, or Puerto Rican. Differences are notoriously hard to predict, as the low R-squared values illustrate. But more noteworthy is that for neither group are either wave one immigration attitudes *or* pan-ethnic attachments strong predictors of

Asian Americans—see Appendix Table A4—in the civic realm, political institutions and elites generally emphasize pan-ethnic over national origin identities, with research establishing the political influence of these broad attachments under varied conditions (Barreto, 2007; Junn and Masuoka, 2008; Pérez, 2015*a*,*b*; Kuo, Malhotra and Mo, 2017). Indeed, part of the political power of pan-ethnic identities comes from their encapsulation of narrower forms of identification (Pérez, 2020, 2021).

[10]We measure immigration attitudes via a seven-category scale ranging from 1 ("Return illegal immigrants to their native countries") to 7 ("Create a pathway to U.S. citizenship for illegal immigrants").

| | Asian Am | Asian Am | Latino | Latino |
|---|---|---|---|---|
| Intercept | $-0.579$ | $-0.295$ | 0.355 | 0.763 |
| | (0.603) | (0.665) | (0.649) | (0.837) |
| Education | 0.016 | 0.005 | $0.081^{*}$ | $0.080^{*}$ |
| | (0.032) | (0.033) | (0.031) | (0.031) |
| Female | $-0.193$ | $-0.084$ | $-0.203$ | $-0.322$ |
| | (0.132) | (0.134) | (0.174) | (0.180) |
| Age | 0.236 | $-0.003$ | $-0.212$ | $-0.027$ |
| | (0.432) | (0.436) | (0.584) | (0.602) |
| Logged Income | 0.067 | 0.018 | $-0.282^{*}$ | $-0.310^{*}$ |
| | (0.065) | (0.072) | (0.115) | (0.117) |
| Born in U.S. | 0.040 | 0.037 | $-0.201$ | $-0.444$ |
| | (0.141) | (0.140) | (0.263) | (0.274) |
| U.S. Citizen | 0.039 | 0.224 | 0.169 | 0.183 |
| | (0.204) | (0.213) | (0.280) | (0.290) |
| Chinese | $-0.005$ | $-0.165$ | | |
| | (0.153) | (0.156) | | |
| Filipino | 0.189 | 0.106 | | |
| | (0.186) | (0.187) | | |
| Support Pathway | | $-0.035$ | | $-0.021$ |
| | | (0.029) | | (0.049) |
| Pan-Ethnic | | 0.042 | | $-0.029$ |
| Attachment | | (0.040) | | (0.054) |
| Spanish | | | $-0.094$ | $-0.288$ |
| | | | (0.281) | (0.291) |
| Mexican | | | $-0.341$ | $-0.282$ |
| | | | (0.193) | (0.202) |
| Puerto Rican | | | $-0.130$ | 0.062 |
| | | | (0.291) | (0.315) |
| $R^2$ | 0.024 | 0.034 | 0.071 | 0.097 |
| Num. obs. | 229 | 204 | 220 | 194 |

$^{*}p < 0.05$

Table 1: Linear regressions predicting the *change* between fall 2018 GOP partisanship and spring 2016 GOP partisanship as a function of key covariates. Higher values of the dependent variable indicate pro-GOP shifts.

whose partisanship shifts.[11][12] This contrasts with an identity threat perspective, under which we might expect those with stronger spring 2016 pan-ethnic attachments to shift disproportionately toward the Democrats by fall 2018.[13]

## Priming Pan-Ethnic Attachments?

Table 1 probes what predicts whose partisanship shifts. Still, there is an alternative approach to examining changes over time: did the period between spring 2016 and fall 2018, when Trump went from a presidential contender to a president, strengthen the relationship between pan-ethnic attachments and partisanship? To be sure, any such changes could be driven by learning *or* priming (Lenz, 2013). But they provide a benchmark to assess the stability of partisanship and its correlates.

Table 2 enables us to answer that question by reporting models in which partisanship is regressed on variables including pan-ethnic attachments separately for Asian American and Latino respondents in spring 2016 and fall 2018. Pan-ethnic identity is always negatively related to Republican partisanship. The 2016 coefficient is slightly stronger for Latinos than for Asian Americans, and the relationships strengthen slightly for both groups by fall 2018. Still, neither change is statistically significant (one-sided p=0.09 for Asian Americans and 0.16 for Latinos), meaning that the evidence of a strengthening relationship

[11]Another prospect is that shifts may be related to prior political knowledge. While Asian American respondents who score higher on a two-item political knowledge battery are somewhat more likely to shift toward the Democrats ($\beta = -0.18, SE = 0.09$), the relationship among Latinos is nearly the opposite ($\beta = 0.25, SE = 0.13$).

[12]While beyond the scope of our paper, it is noteworthy that for Latinos, income and education predict in opposite directions, with those who are higher in education but lower in income shifting to the GOP.

[13]Appendix Table A6 probes a related but separate question by regressing partisanship or partisan feeling thermometers on spring 2016 measures including a key issue position (support for a pathway to citizenship) alongside pan-ethnic attachments. With controls for the lagged dependent variable, such models allow us to estimate whether key wave-one measures predict subsequent partisanship or affect toward the parties. The models show that issue positions have some predictive power while pan-ethnic attachments do not.

| | Asian Am Spring 2016 | Asian Am Fall 2018 | Latinos Spring 2016 | Latinos Fall 2018 |
|---|---|---|---|---|
| Intercept | 5.373* | 5.362* | 2.735* | 1.988* |
| | (1.181) | (1.196) | (0.894) | (0.816) |
| Pan-ethnic Attach, W1 | −0.099 | | −0.169* | |
| | (0.078) | | (0.076) | |
| Education, W1 | −0.090 | −0.076 | −0.050 | 0.066 |
| | (0.063) | (0.062) | (0.045) | (0.044) |
| Female, W1 | −0.338 | −0.499* | 0.471 | 0.303 |
| | (0.256) | (0.252) | (0.256) | (0.245) |
| Age, W1 | 0.709 | 0.821 | 1.464 | 1.053 |
| | (0.847) | (0.835) | (0.860) | (0.821) |
| Income, W1 | −2.538 | −2.173 | 11.182* | 7.682* |
| | (1.941) | (1.922) | (2.733) | (2.692) |
| Born in U.S., W1 | −0.240 | −0.287 | 0.948* | 0.730* |
| | (0.275) | (0.269) | (0.376) | (0.360) |
| Citizen, W1 | 0.086 | 0.296 | −0.350 | −0.067 |
| | (0.398) | (0.398) | (0.409) | (0.403) |
| Pan-ethnic Attach, W3 | | −0.162* | | −0.269* |
| | | (0.080) | | (0.068) |
| Spanish, W1 | | | 0.134 | 0.098 |
| | | | (0.398) | (0.385) |
| $R^2$ | 0.049 | 0.067 | 0.163 | 0.200 |
| N | 227 | 224 | 217 | 209 |

$^*p < 0.05$

Table 2: This table reports regressions of Republican partisanship on demographic variables including pan-ethnic attachments separately for waves 1 and 3.

itself isn't strong. It's plausible that the relationship between these identity attachments and partisanship strengthened, but this evidence is far from definitive—and any such shifts do not appear to have been substantively large.[14]

Appendix Table A8 reports similar models with measures of national-origin identity. There's no strong evidence that Asian Americans who identify more strongly with their national origins are less likely to identify as Republican in either spring 2016 or fall 2018. For Latinos, the strength of national-origin identification does grow slightly—from -0.163 (SE=0.07) to -0.200 (SE=0.06)—but that shift does not approach significance (one-sided p-value=0.35).[15] For both pan-ethnic and national-origin identities, any priming effects of the 2016-2018 period were limited.

## Cross-Lagged Regressions

As an additional probe of the interplay between respondents' attachment to their pan-ethnic identities and partisan identities, we follow Engelhardt (2020) by using cross-lagged regressions in which we regress wave three measures of each on wave one measures of both. This strategy allows us to assess whether wave one partisanship can predict wave three pan-ethnic identity attachment or vice versa. In short, it is a way of testing whether the evidence is consistent with partisanship shaping subsequent pan-ethnic identity, pan-ethnic identity shaping subsequent partisanship, or both. The identity threat hypothesis leads to the prediction that those with stronger initial pan-ethnic identities will be more likely to shift their partisanship in the face of hostile rhetoric.

As Table 3 shows, the results are asymmetric in a way that suggests the centrality of partisanship. Spring 2016 partisanship strongly predicts fall 2018 attachment to pan-ethnic identity, with more Republican respondents in 2016 less attached to their pan-ethnic identity in 2018 even after accounting for their

[14] For comparable evidence with respect to perceived discrimination, see Appendix Table A7.

[15] Appendix Table A9 shows similar results when predicting feeling thermometers toward the two major parties in each wave. There is no evidence that either immigration attitudes or national origin attachments become stronger predictors of feelings toward either party in later waves.

| | GOP Party ID | Pan-ethnic Index |
|---|---|---|
| Intercept | 0.814* | 2.945* |
| | (0.192) | (0.267) |
| Lagged Pan-ethnic | −0.027 | 0.477* |
| Attachment | (0.030) | (0.042) |
| Lagged GOP | 0.791* | −0.112* |
| Party ID | (0.026) | (0.037) |
| $R^2$ | 0.679 | 0.266 |
| Num. obs. | 444 | 434 |

$^*p < 0.05$

Table 3: Cross-lagged linear regression models in which fall 2018 GOP partisanship (left) or pan-ethnic attachment (right) are regressed on spring 2016 measures of both variables.

2016 responses. However, we do not see the same pattern with pan-ethnic identity: once we account for spring 2016 partisanship, knowing the strength of respondents' pan-ethnic attachments does not provide added predictive power for their downstream partisanship. Prior scholarship has reported the capacity of partisanship to influence other social identities (Margolis, 2018; Egan, 2020; Agadjanian and Lacy, 2021); here, we show that for Asian Americans and Latinos, partisanship can have a similar influence with respect to pan-ethnic identity attachment. Democrats come to adhere more strongly to their pan-ethnic identities than Republicans. Partisanship is not weak—for a subset of respondents, it's sufficiently strong so as to shape subsequent pan-ethnic attachments.

## Randomized Video Experiment

Panel data has key virtues, as it enables researchers to track the same individuals over time and to rule out differential survey response as an explanation for any over-time trends. Still, it is valuable to couple observational analyses of panel data with tests of identity threat that isolate the causal impact of Trump's rhetoric. In part, by doing so, we can determine whether the stability in the face of identity threats reported above persists even in the face of one-sided communication. Here, we employ a pre-registered experiment on the same population-based sample of English-speaking Asian Americans and English- and

Spanish-speaking Latinos. Partisanship can be difficult to move experimentally (Hopkins et al., 2020; Coppock, Green and Porter, 2021), so our focus is on prospective vote choice and party-related attitudes.

The experiment took place in the panel's first wave, so it includes 820 Asian American and 721 Latino respondents. Given the theoretical emphasis on group solidarity and images, we sought to identify video instances of Trump leveling verbal attacks that drew on long-standing stereotypes against the pan-ethnic groups in question. For Latinos, a straightforward choice was a 23-second clip from Trump's June 2015 announcement of a presidential bid which is this paper's epigraph, and in which he called unauthorized Mexican immigrants "rapists." This rhetoric represents a very clear verbal attack on Latinos, especially those of Mexican ancestry (see also Leung, 2021). However, its impacts may be limited by its widespread circulation before March 2016. Appendix Figures A5 and A6 show one image from each of the videos.

Since Trump's 2016 campaign focused less explicitly on immigration from Asia—and since Trump's comments blaming the coronavirus on China were years later (Chan, Kim and Leung, 2021)—we employ a fall 2015 clip in which Trump interrupts an Asian American college student who asked a question about Korea to demand, "are you from Korea?" The student replies that he was born in Texas and raised in Colorado; the exchange is an example of an identity denial that has been shown in other instances to influence attitudes and captures beliefs about perpetual foreignness central to stereotypes of Asian Americans (Cheryan and Monin, 2005; Kuo, Malhotra and Mo, 2017; Zou and Cheryan, 2017). This video, too, has the potential to move attitudes, but we are not assuming any equivalence between the salience, content, mechanisms, or effects of the two videos. The Trump announcement was clearly much more widely viewed, a familiarity which may shape its impacts. Respondents assigned to the control group instead watched similarly timed videos related to astronomy, whether in English or Spanish. We also used auxiliary data to appraise the fidelity of our manipulation and quality of responses. The inter-quartile range for the time spent watching the video was from 33 seconds to 64 seconds, suggesting that a significant majority of respondents did watch the video. In a survey question asked immediately after the video, 85% of respondents reported that they were able to see and hear the video.

Figure 4 illustrates the impacts of seeing the videos on party-related attitudes. These impacts are measured via regression models which condition on indicators for exposure to the treatment as well as basic demographics including pan-ethnic background, language of interview, citizenship, place/country of origin, education, gender, age category, citizenship, birth in the U.S., income, and prior partisanship. One prospect is that the Trump videos may influence attitudes toward Trump but not other prospective GOP candidates, measured such that -1 is backing Hillary Clinton, 0 indicates undecided respondents, and 1 indicates backing the Republican. However, the left figure shows that the videos have no statistically significant effects on respondents' assessments of any general election match-ups between Hillary Clinton and Republican contenders, whether they feature Trump or not. Even for Trump, the effects are substantively small: -0.07 (SE=0.05, p=0.11) for Asian Americans and 0.03 (SE=0.04, p=0.46) for Latinos.

We also considered whether the videos affect respondents' willingness to say that the Democrats or Republicans fight for specific groups/issues. Respondents were asked whether 4 statements better describe the Republican Party or Democratic Party via 5-category scales, with 1 indicating the Republican Party, 5 indicating the Democratic Party, and 3 indicating "Both equally." The statements asked about religious people, rich people, fighting to reduce taxes, and treating Asian Americans or Hispanics/Latinos fairly.

As Figure 4's right panel shows, the effects are generally limited. For Asian Americans, there is no strong evidence that the video induced reassessments of the Republican Party on any dimension. However, Latinos are somewhat more likely to say that the Democrats fight for Latinos/Hispanics after seeing the Trump video, with an effect of 0.13 (SE=0.07, p=0.07). They are also somewhat more likely to say that Republicans fight for the rich, with an effect of 0.16 (SE=0.08, p=0.05). Even though many Latino respondents were likely to have been familiar with the Trump video, it did influence party-related attitudes, both about the Latino pan-ethnic group and the wealthy.[16] Such shifts help us put the general stability reported throughout the paper into context, as there is a causal connection between Trump's rhetoric and

[16] Appendix Figure A7 illustrates that the results are similar (albeit with greater uncertainty) when we restrict the sample to low-knowledge respondents who may have been less familiar with Trump's candidacy or comments in March/April 2016. These respondents did not correctly answer two knowledge questions about the U.S. Congress.

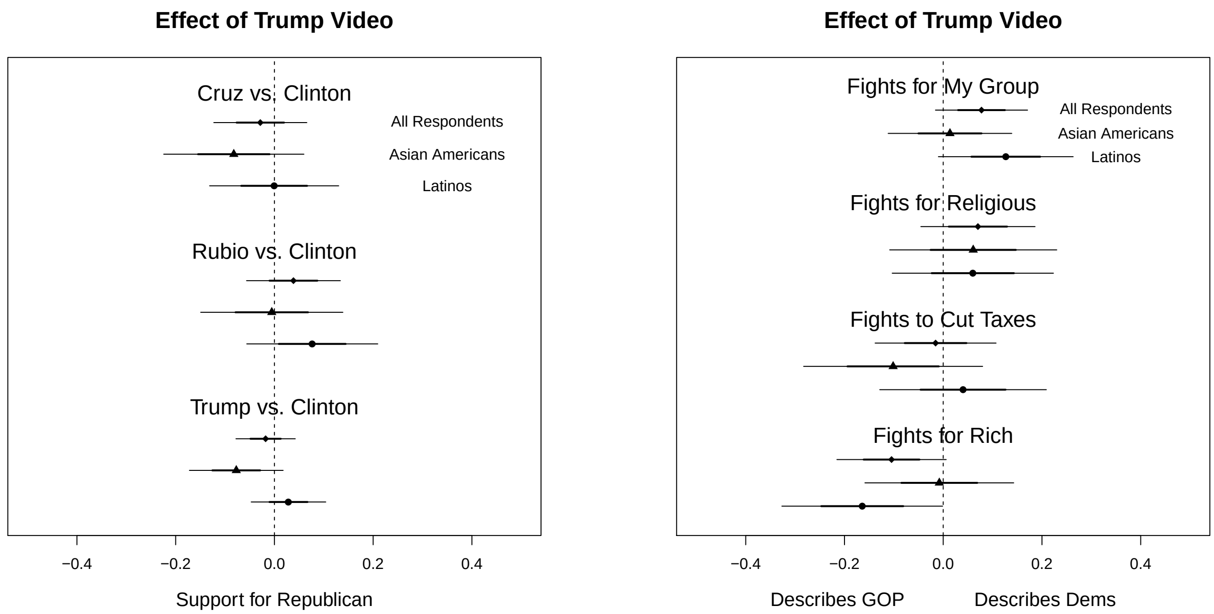


Figure 4: These figures illustrate the impacts of the two video experiments embedded in the spring 2016 survey. The left panel shows impacts on expected general election vote choice while the right panel shows impacts on party-related perceptions.

certain party-related attitudes. Still, the overall story is one of stability, with vote choice in prospective elections eight months away not influenced by either video.

# Implications

We have focused on explaining the evolving nature of partisanship among two of America's fastest-growing racial/ethnic groups—and studying it during a period when rhetoric denigrating these communities grew sharply. We identified three theoretical accounts. The first two are compatible, and suggest that Asian Americans and Latinos are not firmly anchored by partisanship but are likely to respond to identity threats. The third instead emphasizes that Asian Americans' and Latinos' standing partisan ties may remain steady in the face of Trump's invective against these groups.

To test these possibilities, we conducted among the first-ever panel studies of these populations tracking attitudes over multiple years. Employing an embedded video experiment, we find unexpectedly persistent levels of partisan attachment. Pan-ethnic and national-origin identities show stability over this tumultuous period, too. We also find stability in the relationship between pan-ethnic attachments and partisanship, alongside evidence that partisanship may even shape subsequent levels of pan-ethnic attachment, as it does with other social identities (Margolis, 2018; Egan, 2020; Agadjanian and Lacy, 2021). In analyses in Appendix A, we also present evidence from open-ended responses assessing the political parties from the three waves to further bolster these claims. The overall story is one of stability, and not one of parties that are perceived to be increasingly tied to specific ethnic/racial groups. In fact, fewer than 5% of respondents from either pan-ethnic group mention racial/ethnic groups when describing the two parties in any of the three waves. Despite the prominence of identity threat in prior accounts—and despite Trump being a critical case for such accounts—we find little evidence that identity threats shifted partisanship.

The stability we have uncovered fits with other recent work on Latinos' political participation (McCann and Jones-Correa, 2020). It suggests a sense of party identification in these communities that is more developed, coherent, and crystallized than is sometimes presumed based on cross-sectional analyses. Why do these results contrast with the lower levels of partisan attachment reported by seminal works such as Hajnal and Lee (2011)? As detailed above, the absence of Asian Americans who do not speak English from our sampling frame or differential attrition are not complete explanations. Another explanation focuses on the subtleties of question wording. Previous work finds high levels of party non-affiliation using questions that place less emphasis on challenging people's reports of "independence" from either major party. As a result, some of these individuals may have partisan ties that are stronger than what their survey responses indicate. In contrast, our measure of partisanship nudges independents and the unaffiliated to express a partisan leaning. Yet because we are able to reassess this expressed partisanship over time, we can be confident that this stability reflects authentic aspects of the self and is no mere survey artifact.

# Conclusion

The unexpected partisan stability documented here has several theoretical and conceptual implications, but we focus on two. The first concerns the sources of partisanship among Asian Americans, Latinos, and potentially other groups with substantial numbers of foreign-born members. In traditional accounts of partisanship, individuals acquire an initial sense of party identification from their native-born parents. But this pathway is less viable for the many Asian American and Latino individuals who have been raised by immigrant parents or are immigrants themselves. Alongside our findings of partisan stability, this suggests that Asian Americans and Latinos are acquiring a robust sense of partisanship from other, non-parental sources (see also Raychaudhuri, 2020; Carlos, 2021). What might those sources be?

In terms of the origins of the crystallized partisanship we uncover, our own data can only take us back to 2016—and even then, we find evidence of already-stabilized partisan identification among Asian Americans and Latinos. This crystallization is consistent with the gradual political polarization of the U.S. mass public since at least the 1990s (Levendusky, 2009; Mason, 2018), driven in part by racial politics. Indeed, we do not think it coincidental that polarization along party lines has co-occurred with division sparked by political debates about how to address racial inequalities, such as police brutality, welfare, voting rights, and immigration (Pantoja, Ramirez and Segura, 2001; Bowler, Nicholson and Segura, 2006). Based on this evidence, we find it highly plausible that the crystallization of partisanship among Asian Americans and Latinos, at least in the aggregate, is a byproduct of partisan polarization in U.S. politics. In sum, one prospect is that contemporary political polarization facilitates the acquisition of partisanship, as the political parties send consistent, high-volume signals about their values and positions. If so, the political incorporation of heavily immigrant pan-ethnic groups may differ from that of prior generations.

Another, overlapping possibility is that members of these communities may be more attentive to politics than current thinking suggests. Absent strong parental cues, Asian American and Latino individuals might compensate by encountering, processing, and storing information about parties from other sources (see also Raychaudhuri, 2020; Carlos, 2021). This is consistent with a variety of perspectives on political information-processing, which show that people can develop preferences for political objects without fully retaining details of the reasons for those preferences (Lodge, Steenbergen and Brau, 1995).[17] It also suggests the value of reconsidering differences across racial/ethnic groups in how people learn about politics over time, how they store information, and under what circumstances researchers are best able to recover and utilize party-centered information. The immigrant-native distinction in political socialization may have important but underappreciated limits.

Ongoing debates about partisanship among Asian Americans and Latinos continue to foreground theoretical approaches that are necessarily incomplete. In the traditional, parent-driven model of partisanship development, many Asian American and Latino individuals seem at a loss, and yet our findings illustrate that they have somehow acquired robust levels of partisan attachment. In contrast, in models that are more sensitive to the unique experiences of Asian American and Latino communities, alarming levels of disengagement with parties are attributed to ethnic/racial marginalization. Still, we find that even in a context of efforts to further marginalize these groups, their sense of partisanship is more crystallized and stable than typically assumed. Reconciling these accounts will be best achieved not by choosing one perspective but by creating greater synergy between them. This synergy should dig deeper into the political psychology of these groups, and it should acknowledge the prospect that partisanship's levels and changes may have different explanations.

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[17]It is consistent with work on implicit cognition, too (Pérez, 2016; Theodoridis, 2017).

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# Appendix

## Table of Contents

# A Open-Ended Responses

In all three waves, respondents were asked open-ended questions assessing the same randomly selected major political party. The question read, "If you had to describe the [Democratic/Republican] Party in one sentence, what would you say?" Replies were on average 41.8 characters describing Democrats and 42.1 describing Republicans. When looking only at responses in Spanish, the figures are 39.1 and 39.7, respectively. 52 fall 2018 respondents—11%—did not answer either question.

To quantify responses falling into key categories, research assistants annotated the responses using the codebook below. We focus here on the results for the 175 respondents who were randomly assigned to assess the Democrats in the fall 2018 wave and the 173 assigned to assess the GOP. All responses were coded by 2-3 independent annotators; we report means across annotators.

Views of the parties are thought to be influenced by their perceived connections with visible social groups (Green, Palmquist and Schickler, 2002). Given the identity threat hypothesis, we are especially interested in the extent to which Trump's rhetoric racialized the GOP (or their Democratic opponents) in the eyes of Asian American and Latino respondents. To assess that possibility, Figure A1 illustrates the fraction of responses coded into key categories such as those which mention socio-economic or ethnic/racial groups. A response can fall into multiple categories, so the individual who said in 2018 that the Democratic Party "tries to include everyone and compassionate to minorities, immigrants, and labor" is coded as mentioning both socio-economic and ethnic/racial groups.

Overall, the figures show significant stability, with relatively few respondents mentioning social/economic groups and fewer still mentioning ethnic/racial groups in assessing either party. With respect to Democrats, 15% of spring 2016 responses mentioned social/economic groups like the "working class" or "the little guy," a fraction that fell to 8% by 2018. But mentions of ethnic/racial groups in perceiving the Democrats were just 4% in spring 2016 and just 0.6% by fall 2018. By contrast, mentions of ideology (not shown) are much more frequent—for example, Republicans described their party by mentioning ideology in fewer than 12% of posts, while for Democrats the figure is 13%.

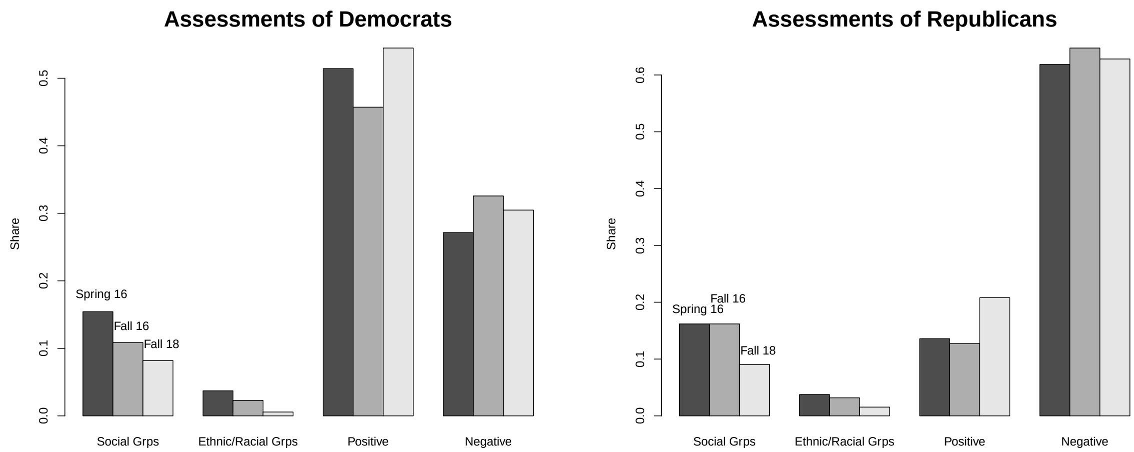


Figure A1: Distribution of open-ended assessments by Latino and Asian American respondents who provided responses in spring 2018. Respondents were randomly assigned to assess one party; N=175 for Democrats and 173 for Republicans.

The story for perceptions of the Republican Party is similar. Four percent of respondents mention ethnic/racial groups in their spring 2016 assessments, a number that falls to just 2% by fall 2018. Even in the midst of the 2016 campaign, Asian American and Latino respondents were more likely to describe the parties with respect to socioeconomic groups than racial/ethnic groups. This is inconsistent with strong claims that the 2016-2018 period increased the extent to which the two parties were viewed through the lens of racial and ethnic groups.

We do see again the heavily Democratic tilt of these respondents overall, as the Democratic Party is consistently assessed more positively while the Republican Party is assessed more negatively. That said, positive assessments of the GOP rise somewhat between 2016 and 2018, from around 14% to 21%. In one representative 2018 answer, a respondent said the GOP "has common sense and truly acts on protecting [the] USA." Still, the overall story is one of stability, and not one of parties that are perceived to be increasingly tied to specific ethnic/racial groups.

## A.1 Codebook for Open-ended responses

NEGATIVE — code as 1 if the remark's tone is negative toward the party in question, 0 otherwise

POSITIVE — code as 1 if the remark's tone is positive toward the party in question, 0 otherwise

ETHGROUPS — code as 1 if the response makes a mention of an ethnic/racial group (e.g. minorities, Hispanics, Mexicans, Chinese people, etc.), 0 otherwise

SOCGROUPS — code as 1 if the response makes mention of a socioeconomic group (e.g. the poor, the wealthy, the middle class), 0 otherwise

EQUALITY — code as 1 if there is any mention of equality or fairness (e.g. "treats groups equally"), 0 otherwise

EXCLUSION — code as 1 if the party is identified as exclusive, discriminatory, or only for some people; 0 otherwise

IDEOLOGY — code as 1 if the party's political ideology is mentioned (e.g. liberal, conservative, libertarian), 0 otherwise

ISSUES — code as 1 if a specific political issue is mentioned, 0 otherwise

IMMIGRATION — code as 1 if immigration is mentioned, 0 otherwise

APOLITICAL — code as 1 if the response indicates the respondent is apolitical, 0 otherwise

# B Partisan Identification Over Time

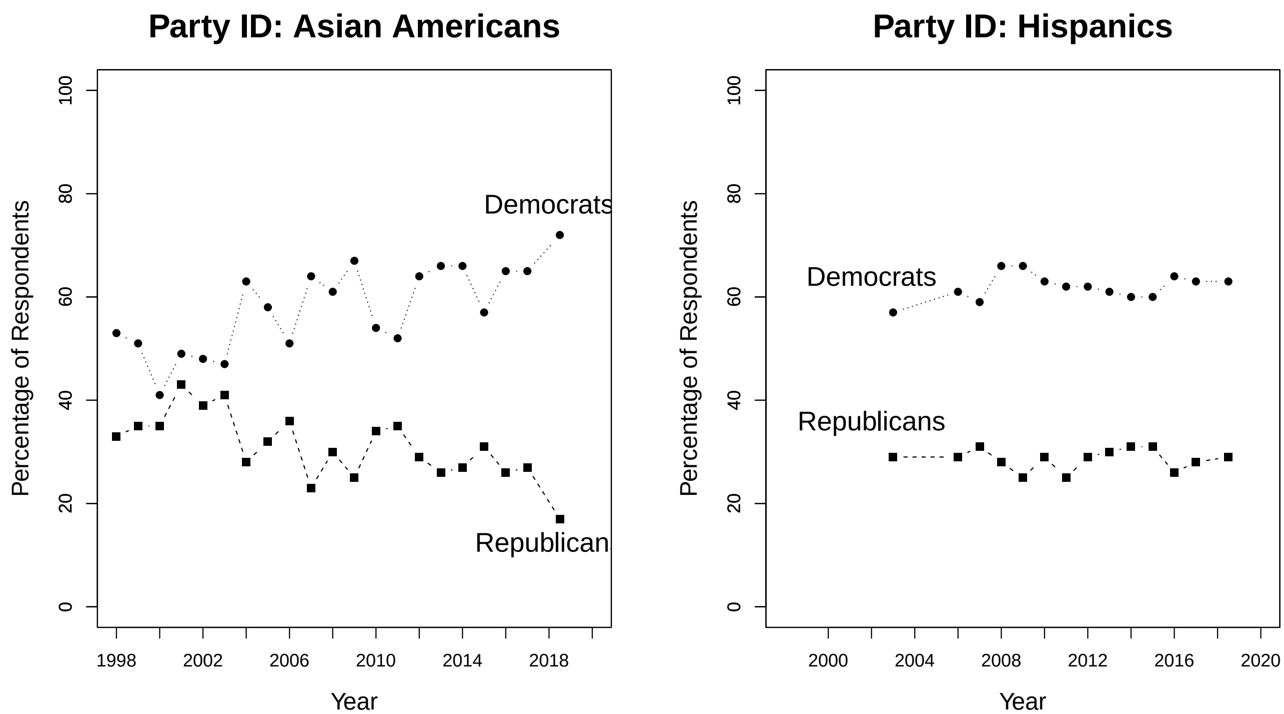


Figure A2: This figure illustrates the percentage of Asian American (left) and Latino (right) respondents to Pew Research Surveys conducted in English and Spanish who identify with or lean to either party by year.

## C Descriptive Statistics

| | Spring '16 | | Fall '16 | | Fall '18 | |
|---|---|---|---|---|---|---|
| | Mean | SD | Mean | SD | Mean | SD |
| Citizen | 0.838 | 0.369 | 0.851 | 0.357 | 0.852 | 0.356 |
| Born US | 0.469 | 0.499 | 0.460 | 0.499 | 0.457 | 0.499 |
| Japanese | 0.140 | 0.347 | 0.145 | 0.352 | 0.196 | 0.398 |
| Chinese | 0.245 | 0.431 | 0.263 | 0.441 | 0.270 | 0.445 |
| Indian | 0.147 | 0.354 | 0.152 | 0.359 | 0.135 | 0.342 |
| Filipino | 0.187 | 0.390 | 0.178 | 0.383 | 0.157 | 0.364 |
| Education | 15.838 | 2.471 | 15.900 | 2.366 | 16.030 | 2.090 |
| Female | 0.570 | 0.495 | 0.537 | 0.499 | 0.535 | 0.500 |
| Income | 87.074 | 63.823 | 86.557 | 63.595 | 93.239 | 67.060 |
| Age | 44.964 | 14.835 | 46.173 | 15.265 | 47.335 | 15.198 |
| Didn't Vote '16 | 0.431 | 0.496 | 0.378 | 0.486 | 0.348 | 0.477 |
| California | 0.323 | 0.468 | 0.328 | 0.470 | 0.361 | 0.481 |
| Texas | 0.061 | 0.240 | 0.053 | 0.224 | 0.057 | 0.231 |
| New York | 0.074 | 0.261 | 0.077 | 0.267 | 0.100 | 0.301 |
| Florida | 0.039 | 0.193 | 0.046 | 0.209 | 0.030 | 0.172 |

Table A1: This table provides descriptive statistics for Asian American respondents' wave 1 answers for those completing each of the three waves.

| | Spring '16 | | Fall '16 | | Fall '18 | |
|---|---|---|---|---|---|---|
| | Mean | SD | Mean | SD | Mean | SD |
| Citizen | 0.743 | 0.437 | 0.790 | 0.408 | 0.816 | 0.388 |
| Born US | 0.495 | 0.500 | 0.555 | 0.497 | 0.596 | 0.492 |
| Hispanic | 1.000 | 0.000 | 1.000 | 0.000 | 1.000 | 0.000 |
| Spanish | 0.429 | 0.495 | 0.345 | 0.476 | 0.256 | 0.437 |
| Mexican | 0.584 | 0.493 | 0.551 | 0.498 | 0.484 | 0.501 |
| Cuban | 0.040 | 0.197 | 0.051 | 0.220 | 0.067 | 0.251 |
| Puerto Rican | 0.101 | 0.302 | 0.113 | 0.317 | 0.130 | 0.337 |
| Education | 12.100 | 3.576 | 12.638 | 3.500 | 13.170 | 3.079 |
| Female | 0.480 | 0.500 | 0.438 | 0.497 | 0.448 | 0.498 |
| Income | 55.319 | 50.138 | 62.611 | 51.922 | 69.361 | 53.975 |
| Age | 44.017 | 15.906 | 43.695 | 15.608 | 44.489 | 15.411 |
| Didn't Vote '16 | 0.526 | 0.500 | 0.454 | 0.498 | 0.386 | 0.488 |
| California | 0.337 | 0.473 | 0.323 | 0.468 | 0.291 | 0.455 |
| Texas | 0.194 | 0.396 | 0.192 | 0.395 | 0.166 | 0.373 |
| New York | 0.062 | 0.242 | 0.066 | 0.249 | 0.085 | 0.280 |
| Florida | 0.099 | 0.299 | 0.124 | 0.330 | 0.135 | 0.342 |

Table A2: This table provides descriptive statistics for Latino respondents' wave 1 answers for those completing each of the three waves.

# D Benchmarking with the 2016 NAAS

Due to limited resources, our survey was only conducted in English and Spanish, raising questions about the impacts of excluding Asian American respondents who are not fluent in English. To assess those impacts, we present data from the 2016 NAAS post-election telephone survey (Ramakrishnan et al., 2016). The survey was conducted in 10 non-English languages, including Mandarin, Cantonese, Korean, Vietnamese, Hindi, Tagalog, Japanese, Hmong, Cambodian, and Spanish. It thus allows us to evaluate the consequences of only surveying Asian Americans in English.

Table A3 presents the same demographics as Tables A1 and A2 for 1) all NAAS respondents (n=4,362, columns 1-2); 2) only those who responded in a non-English language (n=1,878, columns 3-4); 3) and only those who responded in English (n=2,484, columns 5-6). While it would certainly have been desirable to interview Asian American respondents in non-English languages, Table A3 and the NAAS data generally indicate some mitigating factors which may limit the impact of this decision on our substantive conclusions. Even though only 25% of Asian American respondents were born in the U.S., their rate of taking the NAAS in English is 57%, and so higher than that of Latinos (50%). Relatedly, it's important to note that some groups of Asian Americans have high rates of interviewing in English, even when offered the opportunity to interview in other languages. Notably, 99% of respondents of Indian descent, 94% of respondents of Filipino descent, and 83% of respondents of Japanese descent took the survey in English.

There are certainly differences between Asian Americans who do and do not respond in English. While 41% of those who take the survey in English were born in the US, just 3% of those who take the survey in another language were. Those who took the survey in a non-English language are on average less educated (11.1 vs. 15.4 years), have lower average incomes ($48K vs. $98K), and have lower rates of not voting (14% vs. 29%). Importantly for our purposes, Asian American NAAS respondents who don't speak English are more likely to be toward the center of the seven-category partisan identification scale (3.6 versus 2.9).

| | All | | Non-English | | English | |
|---|---|---|---|---|---|---|
| | Mean | SD | Mean | SD | Mean | SD |
| Citizen | 0.929 | 0.257 | 0.894 | 0.309 | 0.955 | 0.207 |
| Born | 0.249 | 0.433 | 0.032 | 0.176 | 0.413 | 0.493 |
| Japanese | 0.115 | 0.319 | 0.046 | 0.209 | 0.167 | 0.373 |
| Chinese | 0.109 | 0.311 | 0.217 | 0.412 | 0.027 | 0.162 |
| Indian | 0.115 | 0.319 | 0.002 | 0.040 | 0.200 | 0.400 |
| Filipino | 0.114 | 0.318 | 0.015 | 0.121 | 0.189 | 0.392 |
| Education | 13.577 | 5.239 | 11.105 | 5.921 | 15.445 | 3.687 |
| Female | 0.470 | 0.499 | 0.542 | 0.498 | 0.416 | 0.493 |
| English | 0.569 | 0.495 | 0.000 | 0.000 | 1.000 | 0.000 |
| Income | 75.739 | 68.926 | 47.594 | 50.835 | 98.046 | 73.104 |
| Age | 53.538 | 19.284 | 61.771 | 15.179 | 47.398 | 19.730 |
| Didn't vote '16 | 0.198 | 0.399 | 0.287 | 0.452 | 0.137 | 0.344 |
| California | 0.340 | 0.474 | 0.348 | 0.476 | 0.335 | 0.472 |
| Texas | 0.042 | 0.201 | 0.051 | 0.219 | 0.035 | 0.185 |
| New York | 0.104 | 0.305 | 0.091 | 0.287 | 0.114 | 0.317 |
| Florida | 0.036 | 0.185 | 0.032 | 0.176 | 0.038 | 0.192 |
| GOP Party ID* | 3.224 | 1.817 | 3.645 | 1.778 | 2.906 | 1.782 |

Table A3: This table provides descriptive statistics for Asian American respondents to the NAAS, with statistics for the full Asian American sample (n=4,362) in the left columns and only those interviewed in non-English languages (n=1,878) in the right columns. Party ID is measured via seven categories.

| | Fall 2016 | Fall 2018 | P-value |
|---|---|---|---|
| Pan-Ethnic | 4.916 | 4.680 | 0.017 |
| Nat'l Origin | 4.753 | 4.650 | 0.405 |
| American | 3.696 | 3.853 | 0.174 |
| Party | 7.466 | 6.426 | 0.000 |

Table A4: This reports the mean rank given to different identities as well as the corresponding p-values from two-sided t-tests. n=408-416.

# E Reweighting

Tables A1 and A2 illustrate that while the attrition across the three waves for the English-speaking Asian American sample is limited, the attrition for the Latino sample is evident across a wider range of variables. Given that, we generated weights for spring 2018 respondents based on the probability of individuals with similar demographic profiles failing to participate in the fall 2018 survey. Specifically, we fit logistic regression models separately to the full Asian American and Latino samples predicting spring 2018 survey completion as a function of various spring 2016 demographics, including citizenship, being born in the U.S., national origin, education, gender, income, age, and validated 2016 voting. From these models, we then calculated weights for each respondent indicating how likely they were to not participate in the fall 2018 wave. For Asian Americans, these weights vary from 0.64 to 1.30, while for Latinos, they vary from 0.36 to 1.34. Higher weights indicate respondents whose demographic profiles make them less likely to participate in the fall 2018 wave. In Figure A3, we reproduced the partisan transition results using these weights.

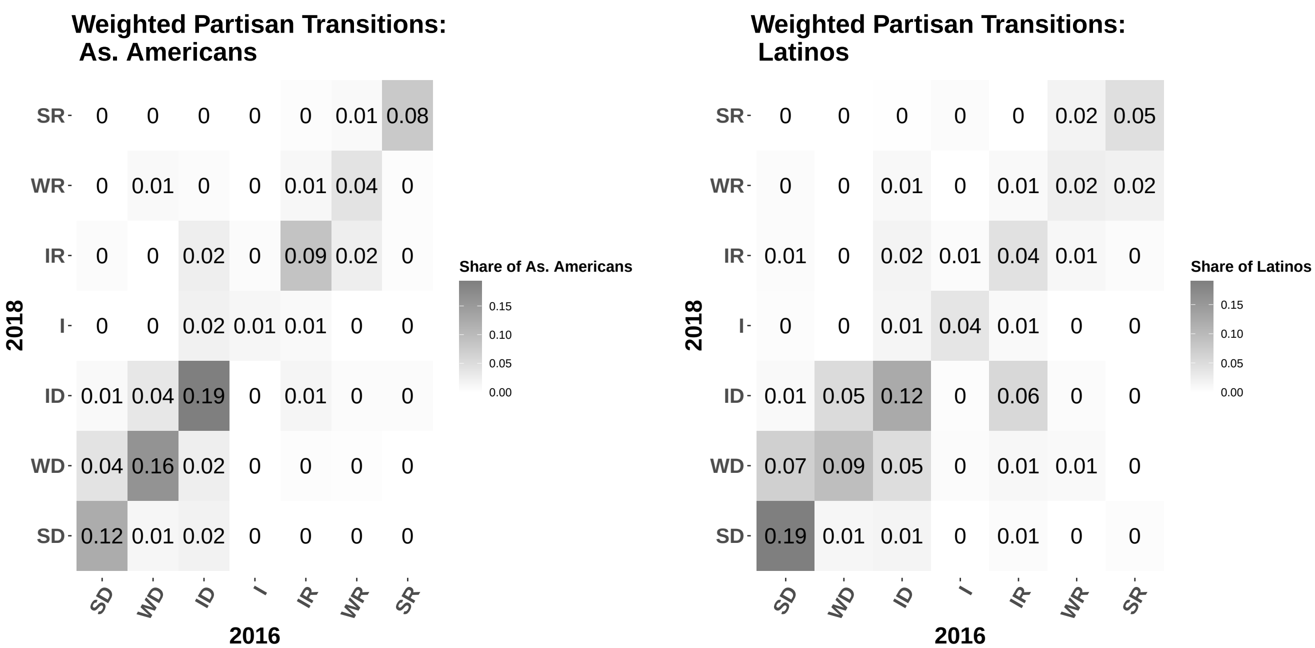


Figure A3: These figures present the partisan transitions after reweighting observations to address differential attrition.

| | Change in PID |
|---|---|
| Intercept | 0.55* |
| | (0.20) |
| Gain Employment | −0.02 |
| | (0.10) |
| Lose Employment | −0.04 |
| | (0.11) |
| Education, W1 | −0.01 |
| | (0.01) |
| Latino, W1 | 0.13* |
| | (0.06) |
| Spanish, W1 | −0.01 |
| | (0.09) |
| Female, W1 | 0.06 |
| | (0.05) |
| Age, W1 | −0.18 |
| | (0.19) |
| Income, W1 | 0.00 |
| | (0.00) |
| Employed, W1 | −0.12 |
| | (0.11) |
| Unemployed, W1 | −0.06 |
| | (0.12) |
| Retired, W1 | −0.11 |
| | (0.12) |
| Born in US, W1 | −0.05 |
| | (0.05) |
| $R^2$ | 0.04 |
| Num. obs. | 449 |

$^*p < 0.05$

Table A5: This table reports the wave-one predictors of changing partisanship over the period, with the outcome coded such that higher values indicate larger shifts toward either party.

# F Question Wording

## National Origin and Pan-ethnic Attachment

Respondents were asked two survey questions to assess their pan-ethnic and national-origin attachments. They were told: "Please tell us whether you agree or disagree with the following statements." The statements then included: "Being [pan-ethnic adjective/national origin adjective] is *not* important to my sense of what kind of person I am" as well as "Identifying with other [pan-ethnic noun/national origin noun] is central to who I am as an individual." Each item was coded 1-4, with 4 expressing higher attachment to the identity. The two items were added to generate each index.

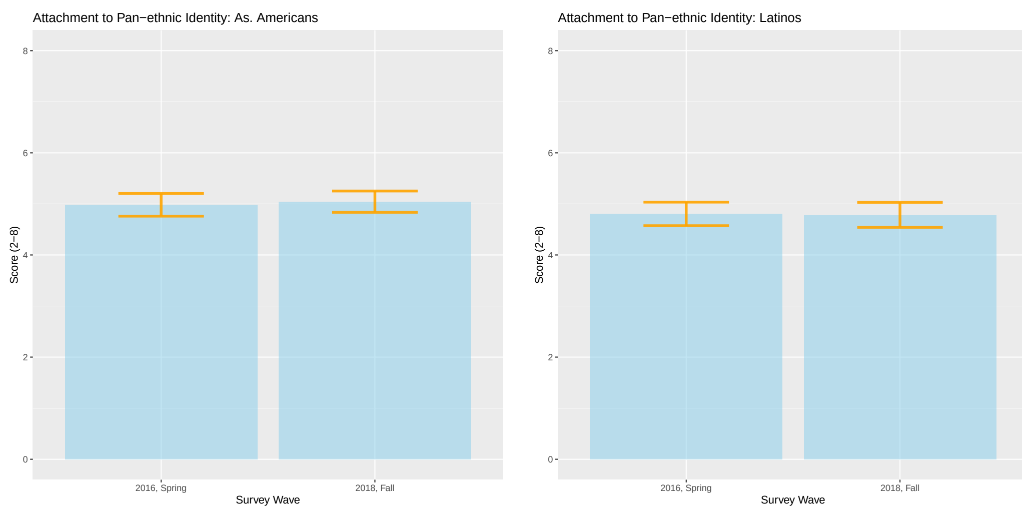


Figure A4: This figure illustrates scores on the index of pan-ethnic attachment separately for Asian Americans (left) and Latinos (right) in spring 2016 and fall 2018.

## Survey questions

- *Support for pathway to citizenship*: "On immigration, some people argue that U.S. policy should focus on [returning illegal immigrants to their native countries./creating a pathway to U.S. citizenship for illegal immigrants.] Other people argue that U.S. policy should focus on [creating a pathway to U.S. citizenship for illegal immigrants/returning illegal immigrants to their native countries]. Still others are somewhere in between. Where would you place yourself on this scale, or haven't you thought much about this?"

- *Identity attachment.* Please tell us whether you agree or disagree with the following statement: "Being [Latino/Asian American] is not important to my sense of what kind of person I am."
- *Identity attachment* Please tell us whether you agree or disagree with the following statement: "Identifying with other [Latinos/Asian Americans] is central to who I am as an individual."
- *Political Knowledge* "How much of a majority is required for the US Senate and House to override a presidential veto?"
- *Political Knowledge* "How long is the term of office for a United States Senator?"

# G Additional Tables and Figures

| | Party ID | Party ID | GOP FT | Dem FT |
|---|---|---|---|---|
| Intercept | 1.57*** | 1.59*** | 40.30*** | 19.48* |
| | (0.42) | (0.42) | (9.43) | (7.68) |
| Weak Democrat | 0.71*** | 0.72*** | | |
| | (0.15) | (0.15) | | |
| Lean Democrat | 1.57*** | 1.57*** | | |
| | (0.14) | (0.14) | | |
| Pure Independent | 2.86*** | 2.86*** | | |
| | (0.34) | (0.34) | | |
| Lean GOP | 2.55*** | 2.55*** | | |
| | (0.18) | (0.18) | | |
| Weak GOP | 3.54*** | 3.54*** | | |
| | (0.21) | (0.21) | | |
| Strong GOP | 4.76*** | 4.76*** | | |
| | (0.20) | (0.20) | | |
| Support Pathway | −0.13*** | −0.14*** | −1.93** | 0.76 |
| | (0.03) | (0.03) | (0.60) | (0.52) |
| Pan-Ethnic Attach. | −0.01 | −0.01 | −0.34 | 0.80 |
| | (0.03) | (0.03) | (0.67) | (0.57) |
| Education | 0.04* | 0.04* | −1.13* | −0.49 |
| | (0.02) | (0.02) | (0.45) | (0.38) |
| Latino | −0.01 | −0.09 | −0.80 | 2.87 |
| | (0.12) | (0.25) | (2.66) | (2.24) |
| Spanish Int. | 0.03 | 0.01 | 7.67 | 8.13* |
| | (0.17) | (0.18) | (3.99) | (3.39) |
| Female | −0.20* | −0.20* | 3.59 | 5.24** |
| | (0.10) | (0.10) | (2.26) | (1.92) |
| Age | 0.43 | 0.44 | 0.15 | −18.02** |
| | (0.32) | (0.33) | (7.45) | (6.28) |
| Income | −0.68 | −0.68 | −13.47 | −8.87 |
| | (0.84) | (0.84) | (19.28) | (16.36) |
| Support Pathway x | | 0.02 | | |
| Latino | | (0.05) | | |
| GOP FT | | | 0.57*** | |
| | | | (0.05) | |
| Dem FT | | | | 0.61*** |
| | | | | (0.04) |
| $R^2$ | 0.75 | 0.75 | 0.40 | 0.51 |
| N | 398 | 398 | 401 | 401 |

$^{***}p < 0.001$; $^{**}p < 0.01$; $^{*}p < 0.05$

Table A6: This table demonstrates regressions of fall 2018 partisanship or party-related affect on spring 2016 covariates.

| | Spring '16 | Fall '18 |
|---|---|---|
| (Intercept) | 3.261* | 2.614* |
| | (0.733) | (0.688) |
| Perceived Disc, W1 | −0.038 | |
| | (0.028) | |
| Latino, W1 | 0.017 | −0.039 |
| | (0.232) | (0.224) |
| Spanish, W1 | −0.748* | −0.546 |
| | (0.354) | (0.338) |
| U.S. Citizen, W1 | −0.019 | 0.114 |
| | (0.298) | (0.284) |
| Born in U.S., W1 | 0.216 | 0.152 |
| | (0.233) | (0.223) |
| Education, W1 | −0.033 | 0.017 |
| | (0.037) | (0.036) |
| Female, W1 | 0.037 | −0.126 |
| | (0.190) | (0.182) |
| Age, W1 | 1.150 | 1.029 |
| | (0.645) | (0.618) |
| Perceived Disc, W3 | | −0.047 |
| | | (0.028) |
| $R^2$ | 0.032 | 0.036 |
| N | 420 | 424 |
| RMSE | 1.936 | 1.860 |

$^*p < 0.05$

Table A7: This table shows regression models of 7-category GOP partisanship in spring 2016 (left) and fall 2018 (right) on a basic set of covariates including respondents' contemporaneous assessments of perceived discrimination in four aspects of their personal lives.

| | 2016 Asian Am | 2016 Asian Am | 2018 Latinos | 2018 Latinos |
|---|---|---|---|---|
| Intercept | 5.639* | 4.864* | 2.385* | 1.651* |
| | (1.197) | (1.172) | (0.850) | (0.808) |
| Nat'l Origin | −0.103 | −0.025 | −0.163* | |
| Idx, W1 | (0.072) | (0.070) | (0.071) | |
| Nat'l Origin | | | | −0.200* |
| Idx, W3 | | | | (0.063) |
| Education, W1 | −0.102 | −0.085 | −0.037 | 0.040 |
| | (0.062) | (0.061) | (0.046) | (0.044) |
| Female, W1 | −0.355 | −0.563* | 0.462 | 0.273 |
| | (0.257) | (0.252) | (0.260) | (0.249) |
| Age, W1 | 0.850 | 1.014 | 1.535 | 1.426 |
| | (0.841) | (0.823) | (0.874) | (0.829) |
| Income, W1 | −2.871 | −2.441 | 11.926* | 9.845* |
| | (1.957) | (1.916) | (2.781) | (2.707) |
| Born in US, W1 | −0.203 | −0.201 | 1.083* | 0.877* |
| | (0.274) | (0.269) | (0.377) | (0.365) |
| Citizen, W1 | 0.019 | 0.160 | −0.405 | −0.079 |
| | (0.400) | (0.392) | (0.421) | (0.396) |
| Interviewed in | | | 0.264 | 0.356 |
| Spanish | | | (0.411) | (0.393) |
| $R^2$ | 0.055 | 0.053 | 0.165 | 0.174 |
| N | 227 | 227 | 213 | 208 |

$^*p < 0.05$

Table A8: This table reports models predicting partisanship in spring 2016 (columns 1 and 3) or fall 2018 (columns 2 and 4).

| | Republican FT | | | Democrat FT | | |
|---|---|---|---|---|---|---|
| | 4/16 | 10/16 | 10/18 | 4/16 | 10/16 | 10/18 |
| Intercept | 56.412* | 40.850* | 68.425* | 14.593 | 18.523 | 23.890* |
| | (10.790) | (9.853) | (10.913) | (9.824) | (10.273) | (9.965) |
| Support Pathway, | −5.280* | −3.956* | −4.946* | 6.048* | 6.067* | 4.799* |
| Wave 1 | (0.651) | (0.595) | (0.659) | (0.593) | (0.620) | (0.601) |
| Natl Origin | −1.046 | −0.096 | −1.221 | 1.948* | 1.396 | 1.309 |
| Attach, W1 | (0.747) | (0.682) | (0.756) | (0.680) | (0.712) | (0.690) |
| Education, W1 | 0.106 | −0.083 | −0.753 | 0.036 | −0.122 | −0.157 |
| | (0.553) | (0.505) | (0.559) | (0.503) | (0.526) | (0.510) |
| Latino, W1 | 5.160 | 1.854 | 0.755 | 0.443 | −1.212 | 2.318 |
| | (3.208) | (2.929) | (3.244) | (2.920) | (3.054) | (2.963) |
| Spanish, W1 | 1.815 | 1.842 | 5.424 | 3.354 | 3.601 | 11.191* |
| | (5.160) | (4.712) | (5.219) | (4.698) | (4.913) | (4.766) |
| Female, W1 | −1.636 | 2.882 | 1.251 | 1.334 | 1.353 | 5.271* |
| | (2.764) | (2.524) | (2.795) | (2.516) | (2.631) | (2.552) |
| Age, W1 | 20.927* | 20.259* | 14.772 | 3.686 | 1.273 | −10.014 |
| | (8.889) | (8.117) | (8.990) | (8.093) | (8.463) | (8.209) |
| Income, W1 | 26.053 | 55.719* | 7.746 | 21.352 | 2.887 | −3.877 |
| | (23.285) | (21.265) | (23.552) | (21.200) | (22.171) | (21.506) |
| $R^2$ | 0.199 | 0.160 | 0.168 | 0.279 | 0.254 | 0.243 |
| N | 354 | 354 | 354 | 354 | 354 | 354 |

$^*p < 0.05$

Table A9: Here, we regress the Republican (left) and Democratic (right) feeling thermometers on wave one measures including a measure of attachment to one's national-origin identity (e.g. Chinese, Mexican, etc.).

# H Video Experiment Materials


Figure A5: Experiment: video of Donald Trump shown to Asian American respondents

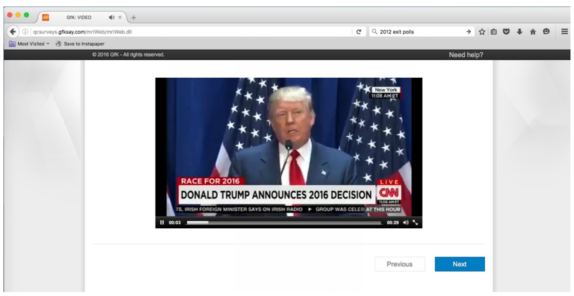


Figure A6: Experiment: video of Donald Trump shown to Latino respondents

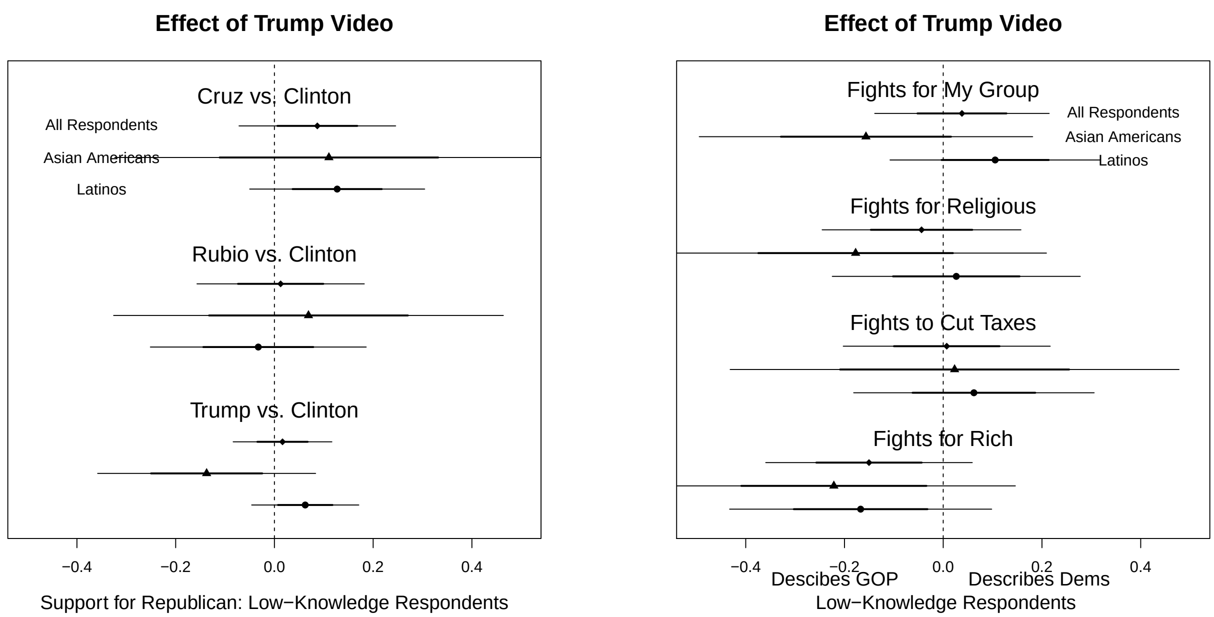


Figure A7: These figures illustrate the impacts of the two video experiments embedded in the spring 2016 survey for respondents with low political knowledge (n=536; 158 Asian American and 378 Latino). These respondents did not know the threshold for overriding a veto in Congress or the numbers of years in a Senator's term. The left panel shows impacts on expected general election vote choice while the right panel shows impacts on party-related perceptions.

# I Pre-analysis Plan

REDACTED, REDACTED

REDACTED, REDACTED

REDACTED, REDACTED

4/16/2016; 9:00 pm EDT

With support from REDACTED, in March and April of 2016, we implemented surveys to populations of approximately 800 Hispanic and 800 Asian American respondents in the United States via the survey firm GfK. Here, we detail our plans to analyze the resulting survey data. In all cases, we will check for balance on various baseline demographic characteristics across experimental treatments (education, income, age, gender, national origin, political partisanship measured prior to the survey, ideology measured prior to the survey, region of the country, immigrant generation, years in the U.S., and urban/rural residence).

**EXPERIMENT 2**

In experiment 2, respondents were separately randomized to view either 1) a brief video of GOP presidential front-runner Donald Trump making negative remarks related to the respondent's pan-ethnic group or 2) a control video about an asteroid which passed near earth in 2015. Following this experiment, respondents were asked to assess whether four statements pertained more to the Republican Party or the Democratic Party. Specifically, respondents were asked to identify which party (if either) better fit the statements: "Fights for religious people," "Fights for rich people," "Treats Hispanics/Latinos unfairly," "Fights to Reduce Taxes," and "Treats Asian Americans Unfairly." We will again use OLS regression to analyze whether respondents who saw Donald Trump make disparaging remarks are more likely to believe that the Republican Party treats Hispanics/Latinos or Asian Americans unfairly. We do not expect broader impacts on the other assessments, but will examine them to assess the breadth of the treatment effects.

**MODERATORS**

For each experiment, we will analyze as possible moderators:

- pre-survey political partisanship. Here, we will separate Republicans and Republican leaners from Democrats and Democratic leaners, with the expectation that Democratic identifiers and leaners will be more influenced by priming perceived discrimination.

- pre-survey strength of political participation, by analyzing the results separately for strong partisan identifiers of both types. Here, our expectation is that weak identifiers will be more influenced by priming perceived discrimination.

- immigrant generation, separating out individuals who arrived in the country as adults from those who arrived as children and those who were born within the U.S. Our expectation is that generations with more experience in the U.S. will show stronger treatment effects.

- strength of pan-ethnic identity (Q12), national origin identity (Q13), and American identity (Q14).

- Political knowledge, as measured by an index created by rescaling Q15 and Q16. Our expectation is that those with more political knowledge will demonstrate stronger effects of priming perceived discrimination.

- language of survey (for Latino respondents, who are able to respond in English or Spanish). Our expectation is that English-language survey takers will show stronger effects.

- perceived discrimination. Here, we will analyze responses for those who report levels of perceived discrimination that are above and below the median. We will pool the available, close-ended measures of perceived discrimination and use factor analysis to create a scale which will then provide a measure of perceived discrimination (see also REDACTED et al. 2015). If the fraction of people who report no discrimination in the open-ended question (Q5) is above 20%, we will create an alternative indicator of perceived discrimination using that. Our expectation is that those with higher levels of perceived discrimination will show stronger effects.

- national origin. For experiment 2 with Latino respondents, we chose a comment by Donald Trump which singled out immigrants coming across the border from Mexico. Accordingly, we expect a stronger response among people of Mexican descent.

For experiment 2 only, we will analyze responses separately for those who report that they could not see or hear the video, and for those who left the video screen in less than 5 seconds.

Finally, we will consider the possibility of interaction effects, such that exposure to the primes about discrimination in experiments 1 and 2 might have a stronger effect than exposure to both individually.