

RESEARCH ARTICLE

Constructing a New Measure of Macropartisanship Disaggregated by Race and Ethnicity

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(Received 24 February 2021; revised 10 July 2021; accepted 27 October 2021; first published online 26 August 2022)

Abstract

Macropartisanship is a measure of aggregate trends in party identification in the mass public that allows researchers to track partisanship dynamically. In previous research, macropartisanship was found to vary in concert with major political events and forces like presidential approval and the economy. However, studying macropartisanship as an aggregate trend assumes that group dynamics within the measure are equivalent. We present a series of new measures of macropartisanship using Stimson's (2018) dyad ratio approach disaggregated by race and ethnicity. We detail the creation of measures for White, Latino, and Black macropartisanship from 1983 to 2016 using more than 500 surveys from CBS News and CBS/New York Times. The resulting data collection is publicly available and can be downloaded in monthly, quarterly, or yearly format. Our initial analysis of these data show that thinking about macropartisanship as a single aggregate measure masks important and significant variation in our understanding of party identification. Change in the measures are uncorrelated. Latino macropartisanship is more volatile and responds more to economic conditions, Black macropartisanship is very stable and has become more Democratic in response to increased polarization, while White macropartisanship has become less responsive to economic conditions as has become more Republican as Republicans have moved to the right.

Keywords: macropartisanship; party identification; race and ethnicity; Latino politics; Black politics

1. Introduction

The importance of the development of party identification to the study of American political behavior cannot be overstated. Often acquired early in life, party identification acts as a great regulator and predictor of political action. It provides shortcuts to voters who are uncertain and organizing mechanisms for information to highly informed voters. And, it is generally quite stable. Changes to an individual's party identification are the exception rather than the rule. While static identification is the norm for individuals, changes observed in macropartisanship in many instances move in predictable and meaningful ways (MacKuen et al. 1989).

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Aggregate measures like macropartisanship, however, miss important realities about American history and democracy. According to the 2018 Current Population Survey, just over 60% of the U.S. population identifies as White, not Hispanic, which means 4 in 10 Americans come from a historically underrepresented racial or ethnic group. Notably, American society is often mischaracterized as fitting within the *liberal democracy* paradigm, and mono-theorizing in this way ignores rampant structures of inequality that have been critical to American political development. One cannot understand American democracy without a critical theory of inequality that accounts for race, ethnicity, and gender and the structures that created and perpetuated that inequality (Smith 1993). Indeed, in at least one account of party change, it is White Americans' attitudes about race that lead to durable change in party coalitions for decades (Carmines and Stimson 1986, 1989). While disputes occur in this literature about a variety of topics ranging from the stability of party identification (e.g. MacKuen et al. 1989; but see Green et al. 1998; 2002) to the nature of issue evolutions (Carmines and Stimson 1989; but see Abramowitz and Saunders 1998), little in these articles addresses the apparent stability in the partisanship of African Americans, or the apparent instability of the partisanship of Latinos, and how they differ fundamentally from those of non-Latino White Americans because of identity, shared history, and the politics of race, ethnicity and inequality. To be sure, a great deal of research exists on the political behavior of African Americans and Latinos. But our aggregate measures of macropartisanship need to be disaggregated if we are to be honest about the fact that (1) these communities differ in their partisanship, its stability and change, and (2) we want macropartisanship and its responsiveness to represent something other than *White* macropartisanship.

In this paper, we rectify this problem by introducing a new dataset of macropartisanship, disaggregated by race and ethnicity such that yearly, quarterly, and monthly calculations of White macropartisanship, Black macropartisanship, and Latino macropartisanship can be presented from 1983 to 2016 for the United States. These data were created through an aggregation of every CBS News and CBS/New York Times survey containing party ID from 1983 to 2016 and by applying Stimson's (2018) dyad ratio algorithm. After demonstrating the measure's validity, our initial analysis shows that traditional movers of macropartisanship, presidential approval and consumer confidence, as well as elite polarization, affect racial and ethnic groups quite differently. We provide for a public release of the data¹ and provide checks on external validity to deal with some potential pitfalls of using this data and time period.

2. Black and Latino macropartisanship—the case for disaggregation

We situate our inquiry at the intersection of the literature on party identification, notably its stability, and the literature on party identification for Blacks and Latinos. The conventional wisdom in American political behavior dates back to the development of the Michigan model (Campbell et al. 1960). Individuals develop party identification through a preadult socialization process. Lacking information about

¹See our dataverse at Johnson and Dyck (2021).

most basic matters of politics, party identification becomes a central organizing factor in personal identity for determining political behavior. Research has tended to show that party identification is stable over the life cycle (Green et al. 1998; 2002; Niemi and Jennings 1991; Sears and Funk 1999).

However, MacKuen et al. (1989) argue that the aggregate trend in macropartisanship is sensitive to short-term fluctuations in observable parts of the political environment. Most notably, they argue that changes in election results, presidential approval, economic conditions, and dramatic events can change party identification trends in predictable ways at the macro level—and that those changes can have long-lasting effects, particularly changes related to economic fluctuations. While Green et al. (1998) and others (Campbell 2010) dispute the volatility in party identification observed in macropartisanship, the measurement strategy opens up the notion that we can measure aggregate trends in party identification and that their fluctuations are meaningful.

While this literature created important new insights into political behavior, we argue there are several reasons that these expectations are on faulty footing for both Blacks and Latinos. For Blacks, we have for decades observed partisan identification and voting rates for the Democratic Party in the 85–95% range. Other factors also distinguish Black voting behavior. For instance, higher levels of religiosity among African Americans do not lead Blacks to be more likely to support Republican candidates as it does for Whites (McDaniel and Ellison 2008). Furthermore, economic differences are not apparent in Black party identification (Dawson 1994). Dawson (1994), whose theory underlies much of what is understood about Black partisanship and voting behavior today, argued that Blacks as individuals make voting decisions based on a community utility heuristic that considers the fate of the community as a whole. When we understand this, we come to understand how Black party ID may differ from White party ID—theories of White partisanship rooted in either the socialization (Campbell et al. 1960) or running tally (Fiorina 1981) framework do not use a group heuristic that focuses identification and voting behavior nearly as clearly and would allow for greater flexibility in responses to variations in the political environment (presidential approval, economic performance, etc.) A micro change in the economy will likely not be as persuasive as the fact that the Democratic Party is associated as the party of civil rights (Carmines and Stimson 1989) and that the Republican Party is viewed as the party of racial conservatism and often has used racialized appeals to court support from White voters (Hutchings and Valentino 2004).

And if the foundations of Black and White party ID are different, there is good reason to believe that Latino party ID may be less firmly rooted than the party ID of Whites and African Americans. There are two primary reasons why we might expect more volatile patterns of partisan change among Latinos (Dyck et al. 2012). The first explanation is a party conversion story. While it is true that party ID is generally stable for most Americans, *comparatively speaking*, immigrants have weaker party attachments (Cain et al. 1991; Uhlaner and Garcia 2005) and are less likely to have socialized bonds of party identification that span generations. The second explanation is a story of population change and immigration. According to the Pew Research Center, the median age among Latinos is 15 years younger than the median age for White Americans (Patten 2016). Therefore, any

observed changes in party identification may be driven by changes in the underlying population. To the extent that we observe any changes in Latino party ID, individuals may be changing party ID, but the process of generational replacement, or perhaps more accurately, generational growth, may be the underlying story. As Hajnal and Lee (2011, 89) write, “we maintain that the unique context of immigration implicates three distinct factors: information uncertainty, ideological ambivalence, and identity formation.” Under these conditions, we would expect that Latinos may not have the firmly rooted partisanship (Green et al. 2002) of non-Latino Whites or of African Americans. In short, we would expect greater stability in Black party identification and greater volatility in Latino party identification, but we can only examine this by disaggregating macropartisanship data by race and ethnicity.

3. Constructing the dataset

In order to disaggregate macropartisanship by race and ethnicity, we first need to determine a source for constructing such a measure. In the present day, we can turn to some large-scale data collection efforts like the Cooperative Congressional Election Survey (CCES), which surveys around 25,000–50,000 Americans every year. Other organizations like Latino Decisions do an admirable job of accurately surveying Latinos. The CCES, however, has only been around since 2006 and Latino Decisions was founded in 2007. Other projects like National Black Election Survey and the Latino National Survey have been fielded in 1 or 2 years. Pew, likewise, has some data trends, like the 2000–2006 series of surveys of Latinos.

Given the literature’s expectations on the stickiness of partisan identification, studying changes across race and ethnicity requires a longer time series than these datasets can provide. Consequently, we make use of the best available data from the 1980s and 1990s to construct an aggregate overtime measure of macropartisanship for Whites, Blacks, and Latinos. In fact, we turn to the same CBS/NYTimes polls used across a number of studies of macropartisanship (e.g. Green et al. 1998). The CBS and CBS/NYTimes data archive is best suited to this project for a variety of reasons. First, CBS/NYTimes has long asked a question on race and started asking respondents if they identified as Hispanic in 1983.² Second, unlike Gallup, CBS/NYTimes uses the classic wording of party ID consistent with the American National Election Study/Michigan Survey Research Center. The Gallup wording has been found in previous research to produce a noisier measure of partisanship subject to considerably more short-term variation, that is, an artifact of measurement error (Abramson and Ostrom 1991). Finally, while there are some other long-standing data collections available in either the ICPSR or Roper archives (e.g. ABC/Washington Post), only the CBS News and CBS/NYTimes polls go back to the early 1980s allowing for the construction of a longer time series.

We gathered all available CBS News and CBS/NYTimes surveys housed at the Inter-university Consortium for Political and Social Research and the Roper Center for Public Opinion Research for the period 1983–2016. We found a total of 392 CBS/NYTimes and 181 CBS News datasets that cover this period that asked

²Though in early years, the survey was inconsistent in regard to asking this question.

respondents both their party ID and race/ethnicity, for a total of 573. We encountered data errors in 12 of the datasets, giving us a total of 561 usable datasets.³ We also note that scores of early datasets failed to ask respondents if they identified as Latino or Hispanic, so these surveys could not be incorporated into our analysis. However, these 561 datasets allow us to construct a disaggregated time series of party identification by race/ethnicity that dates back more than three decades.⁴ For our study period, we have usable data for Whites, African Americans, and Latinos. We also collected data for those identifying as Asian and “Other” in their response to race, but their numbers in many of the surveys was very low. Furthermore, CBS/NYT surveys only regularly added Asian as a racial category in the late 1990s, so much of our sample lacks historical data for the fastest growing immigrant group today.

Our approach has some obvious limitations. First, each survey has a limited number of Black and Latino respondents. Across the 561 datasets, we have a total of 54,409 Black and 33,476 Latino respondents, that is an average of 97.0 Blacks and 58.4 Latinos per survey, meaning the CBS/NYT polls surveyed about 1650 Blacks and just over 1100 Latino respondents in an average year.

While the literature does expect higher volatility for Latino party identification due to the distinct socialization process for newer immigrant groups, we cannot completely rule out the possibility that a portion of this volatility could be a function of the small number of Latinos per survey. Second, the number of surveys per year is inconsistent and the time of year in which the surveys are conducted also varies considerably. In 2 years, 1983 and 1985, we have only two surveys with complete data, whereas after 2000 we have at least 12 surveys each year. Similarly, far more surveys are conducted in election years, while the months immediately preceding an election have the highest number of surveys. A final limitation of the data is that the CBS/NYTimes and CBS News survey is a three-point measure of party identification (Campbell 2010). Using a more comprehensive seven-point scale found in the ANES demonstrates that weak party identifiers and independent leaners are typically indistinguishable (Keith et al. 1992). Consequently, our measure may hide shifts in macropartisanship that take place between pure independents and independent leaners. Due to their distinct process of political socialization, Latinos may simply be more likely to identify with these independent leaners, despite behaving like weak partisans.

³The list of datasets with their accompanying ICPSR and Roper Center identification numbers and date of the survey is included in the dataverse. Some surveys listed the month of the survey, but not the date the survey started. When no date was listed, we listed the start of the month as the date the survey started. If more than one survey was conducted, we listed the start of the month as the date of the first survey, the 15th as the date of the second survey, and the 25th as the date of any month in the event of a third survey.

⁴Barreto and Pedraza (2009) note that the ANES pooled cluster sampling is problematic when trying to select a subsample of Latinos. However, the CBS/NYTimes poll uses random digit dialing using primary sampling units. Similarly, we know that recent CBS/NYTimes polls were conducted in both English and Spanish. Techniques for surveying in Spanish are obviously better today than they were in the 80s and 90s, but these data at least allow us to attempt to create a reasonable overtime measure that spans four decades.

Nevertheless, these data provide the most comprehensive way to construct a measure of macropartisanship that can be disaggregated by race and ethnicity.⁵ As noted above, the number of Black and Latino respondents was often small in a given survey, but as we will show in the next section, when we aggregate the data, this concern is greatly reduced. For example, if we look at quarterly data then we typically have over 400 African American respondents and about 280 Latino respondents. These samples are larger than those of racial and ethnic subgroups in seminal micro-level analyses of group interest and shared fate (c.f. Chong and Kim's 2006).⁶ Additionally, we find that while macropartisanship changes for all racial and ethnic groups, the standard deviations do not vary substantially over time, meaning changes are not likely related simply to changing sample sizes. In sum, we are confident that changes in macropartisanship across racial and ethnic groups is not simply a function of sample size.

4. Measuring macropartisanship

4.1 *The standard measure*

The standard measure of macropartisanship reflects the relative distribution of the two-party partisan identifiers at any given moment. This is simply calculated as the number of Democrats in a sample divided by the number of Democrats and Republicans in the sample multiplied by 100 (Erikson et al. 1998; Green et al. 1998; MacKuen et al. 1989). We find that the Democratic Party held a small, consistent advantage with a median of 53.7% of the two-party identifiers. We also see that the Republican Party gained ground in the 1980s, lost ground in the middle of Clinton's term, gained ground in the early part of the presidency of George W. Bush, but again lost ground in the Obama years (see Figure 1). That macropartisanship slightly favors the Democratic Party and appears to track with presidential approval and economic fortunes is consistent with expectations. We also see that overall macropartisanship is fairly stable, typically hovering around 54% and very rarely going below 50% or over 58%. Essentially, overall macropartisanship indicates a narrowly divided American public that makes only slight course corrections over time.

While aggregate macropartisanship demonstrates both the overall stability and subtle shifts in the population, it fails to examine whether stability and shifts hold across racial and ethnic divisions. Using the same methodology mentioned above, we calculated macropartisanship measures for Blacks, Latinos, and Whites and plotted them over the course of our study (see Figure 2). For African Americans, we see an enormous gap in macropartisanship, with more than 90% of partisan

⁵While it would also be desirable to be able to measure macropartisanship for Asian-Americans and other racial and ethnic subgroups, our sample sizes are only large enough to compute time trend estimates for non-Latino Whites, Latinos, and African Americans.

⁶A random sample of 400 carries a margin of error of +/-4.9% and a random sample of 280 carries a margin of error of +/-5.9%. Additionally, we generated scatterplots and local polynomial smooth plots that show Latino, White, and Black macropartisanship over time with confidence intervals. These figures do show Latino macropartisanship is more volatile than White or (especially) African American macropartisanship (see Online Appendix Figures 3-5).

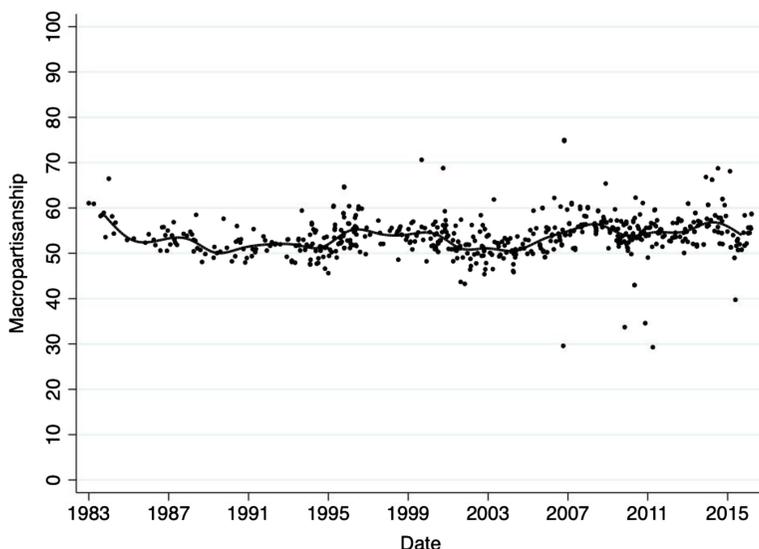


Figure 1. Overall Macropartisanship using CBS/NYTimes Data. June 1983 to August 2016⁷

identifiers falling in the Democratic camp. Furthermore, even though there are more than eight times as many non-Latino Whites than African Americans across all of our datasets, they have similar standard deviations. This provides *prima facie* evidence that African American partisanship is not only distinct from overall macropartisanship but is also more stable. This is consistent with the Dawson's (1994) contention that Black party identification is strongly predicted by linked fate, which downplays short-term forces. One of the remarkable things to observe here is that, running counter to conventional wisdom, the nomination and election of Barack Obama in 2008 produced little movement in Black macropartisanship, as group identity had already placed 9 out of every 10 African American partisans squarely within the Democratic Party. If anything, African Americans began trending even more toward the Democratic Party in the early 1990s.

For Latinos, we also find that the Democratic Party enjoys a significant advantage, though the advantage is not nearly so great, with almost 7 in 10 partisans identifying with the Democrats. We also note that Latino macropartisanship appears less stable. The passage of the IRCA in 1986 seems to have garnered the party of Reagan some additional support, though Latinos drifted back to the Democratic Party for most of the Clinton administration; this also coincides with major shocks to the party system that occurred in California, where the Republican Party targeted Latinos and paid the price in terms of a generation of lost party identifiers (Bowler et al. 2006; Dyck et al. 2012). The Republican Party again picked up support in the wake of 9/11, before Latinos moved away from the Republican Party during

⁷Here, we present the combined scatterplot and the median spline plots; median spline plots are used to illustrate trends. Median splines smooth out functional shapes and their use is exceptionally common in biostatistics and medical research (Perperoglou et al. 2019).

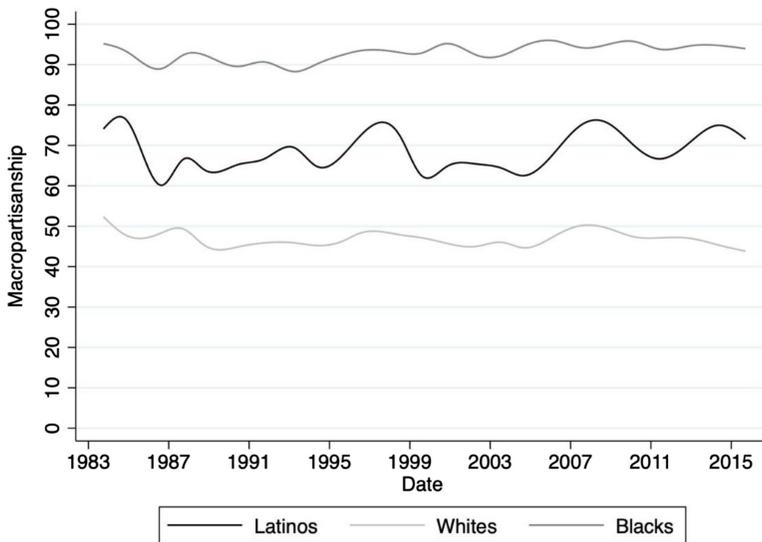


Figure 2. Macropartisanship by Race and Ethnicity. June 1983 to August 2016⁸

the Bush administration in the later years.⁹ Democrats gained their biggest advantage in the wake of the Great Recession and the election of Obama but lost ground as the Spanish-language media and the activist community became increasingly concerned over deportations during Obama's first term in office. However, support for the Democratic Party among Latinos appears to spike in presidential election years. The greater variation observed among Latinos is consistent with the notion that many from the Latino community lack firm roots and ties that span generations with one of the American parties, yet return to the Democratic Party in important election years, all of which is consistent with Hajnal and Lee (2011) core arguments.¹⁰ The Democratic advantage, the volatility in Latino partisanship, and the recent strengthening of this advantage are consistent with other recent work in

⁸Here, we again use median splines to plot macropartisanship, this time by race and ethnicity. We have suppressed the scatterplots to make the figures easier to read.

⁹This likely had far more to do with hard line immigration stances taken by Congressional Republicans and by policies enacted in the states than by Bush himself, who advocated for a pathway to citizenship for undocumented immigrants and was by all accounts, an immigration moderate within his party.

¹⁰Despite the fact that our samples for Latinos have the smallest number of respondents, the variation in sample sizes over the study period gives us some leverage over understanding the trends in the data. We observe fewer respondents in the first half of our time period than in the second half. However, mean Latino macropartisanship for the first half is 68.4% with a standard deviation of 6.9 and for the second half is 69.7% with a standard deviation of 7.7. If changes for Latino macropartisanship were due to small sample sizes, we should expect a much higher standard deviation in the first half compared with the second half, yet we actually find a slightly higher standard deviation in the second half of our sample. This gives us a great deal of confidence that the volatility we are observing is not a function of sampling error, but instead consistent with the theoretical arguments we have outlined. See also footnote 12 regarding this point in relation to Stimson's dyad ratio algorithm.

the field (see Hajnal 2020). Regardless, Figure 2 supports our contention that Latino macropartisanship is distinct from overall macropartisanship.

For the non-Latino White population, we find a small advantage for the Republican Party. Non-Latino Whites drift away from the Democratic Party in the Reagan years consistent with the movement of White men away from the Democratic Party and the creation of the durable gender gap (Kaufmann and Petrocik 1999). White macropartisanship ebbs toward the Democratic Party a bit in the George W. Bush administration's later years, and then slightly away from the Democrats during the Obama years, reflecting the thermostatic nature of public opinion trends more generally (Erikson et al. 2002). However, all of these shifts are small. In short, non-Latino Whites and African Americans seem to shift gradually, supporting Green et al. (1998) contention that macropartisanship changes slowly, as if eroded by the wind, not quickly, as if eroded by the tide (see Figure 2). White partisans are simply far more Republican than African Americans or Latinos, at times appear to drift in the opposite direction of African Americans, and are more stable than Latino partisans.¹¹ In short, the overall measure of macropartisanship found in Figure 1 masks important variations by race and ethnicity found in Figure 2.

4.2 Dyad ratio algorithm

While we have over 500 surveys to construct this 34 year time series of data, the inconsistent time between surveys has to be considered, and the data transformed in some way to correct for the fact that at different times the precision of our estimates differ and that there are some time windows where there are holes in the data. Stimson's dyad ratio algorithm (1991, 2015, 2018) presents a solution for this problem. This factor-analytic algorithm compiles time series data by weighting each individual survey by how well it correlates with the latent concept—macropartisanship (Carlin et al. 2018; Claassen 2019). Point estimates are then exponentially smoothed (see Stimson 2018 for a detailed explanation). Previous macropartisanship studies and policy mood studies used this approach in the United States (see Enns 2016; Stimson, MacKuen, and Erikson 1995; Erikson et al. 2002), while comparativists use this approach to create cross-national measures of presidential approval (see Arnold et al. 2017; Carlin and Hellwig 2019).

Described by Stimson as an algorithm “for building a continuous regular time series from the scraps of dated survey results that are typically available for public opinion analysis,” the algorithm stores data in a sparse matrix, with variable-indexed rows and time-indexed columns (Stimson 2015, 1). Therefore, within each column, there is a time step with a corresponding variable value. In this case, that variable is macropartisanship as the Democratic percentage of two-party PID. The matrix is sparse because surveys are only performed every so often or clustered around certain events (like elections), so most of the time steps will have no data. A dyad is the values for the same variable at different time steps. The dyad ratio, then, is a ratio of one of

¹¹The correlations for macropartisanship across race and ethnicity are all positive and statistically significant at $p < 0.10$. However, these correlations are quite low, ranging from 0.11 for Blacks and Whites to 0.21 for Latinos and Whites.

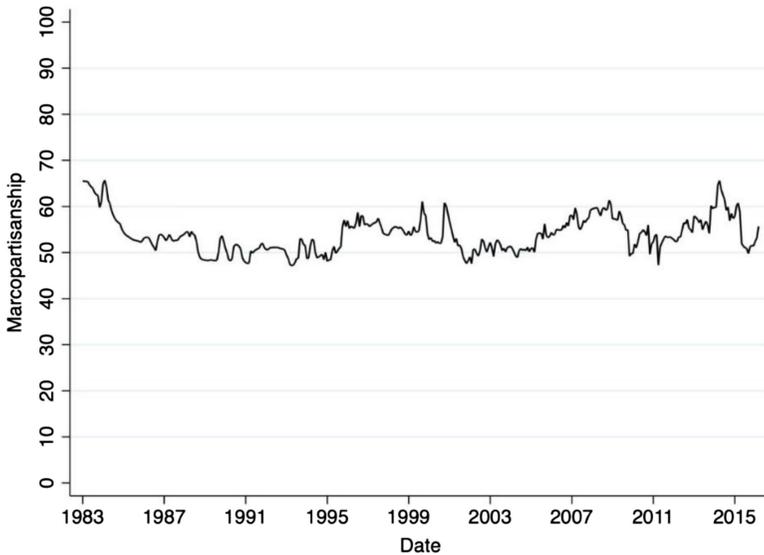


Figure 3. Overall Macropartisanship using Stimson's Dyad Ratio Algorithm. June 1983 to August 2016

these dyads, or the ratio of the value of the variable at one time step to the value at a different time step. The algorithm uses both forward and backward recursion to calculate two dyad ratios at each time step. These dyad ratio estimates in both directions are then averaged to produce an estimate that uses all the available information and that weights all the items equally. Stimson's implementation of the algorithm also smooths the data and accounts for validity estimation.

This is important because it allows us to generate monthly, quarterly, and annual estimates of macropartisanship from 1983 to 2016, even when the frequency of polling varies considerably over time. Here, we see immense value in having 561 surveys over 34 years and we go well beyond other available estimates of macropartisanship that might be gleaned from a single number taken from the American National Election Study or the CCES.

Drawing on Stimson (2018), we start by recalculating macropartisanship for the entire population; here we present the monthly version of the data. A simple line graph of our measure of macropartisanship created by Stimson's dyad ratio algorithm sees that these data closely match what we found in the previous section where we plotted cubic splines of the raw data (see Figure 3). In the raw data, we found the mean of macropartisanship was 53.84, while in our dyad ratio algorithm data have a mean of 53.91, meaning the Democratic Party tended to hold a small advantage. The Republican Party reached approximate parity in the late 1980s until the mid-1990s, again in the immediate wake of 9/11, and during the height of the Tea Party in 2010 and 2011. The Democratic Party held advantages early in the Reagan Administration, in the second term of the Clinton Administration, the second Bush II term, and in Obama's second term. Again, these ebbs and flows in this measure of macropartisanship largely track with the larger literature, with Republicans gaining ground when a president of their party is popular or when

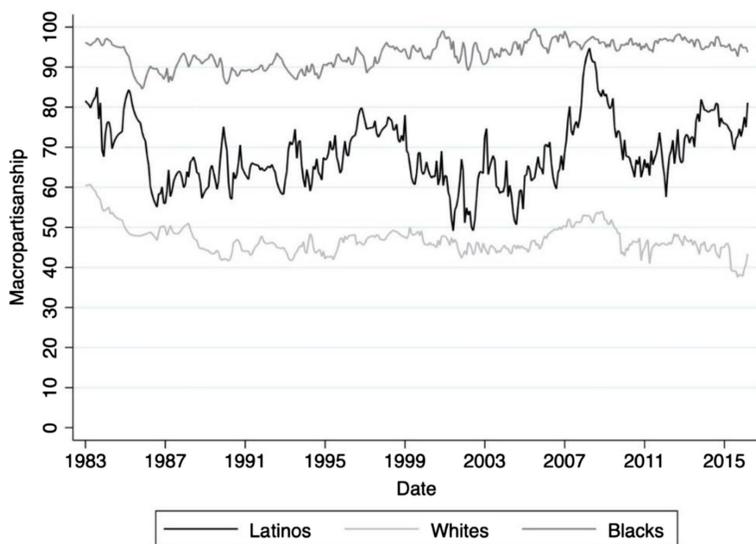


Figure 4. Macropartisanship by Race and Ethnicity using Stimson’s Dyad Ratio Algorithm June 1983 to August 2016

a Democratic President loses popular support, while the Democratic Party gains ground under the opposite conditions.

Now that we have an overall measure of macropartisanship, we can also use Stimson’s dyad ratio algorithm to recalculate macropartisanship by race and ethnicity (see Figure 4). Again, we see the Stimson method produces a figure that closely tracks the cubic spline plots of the raw data found in the previous section (see Figure 2). Using the dyad ratio method, we find that over 9 out of 10 African American partisans identify with the Democratic Party, with a mean macropartisanship score of 93.31, compared to 93.22 in the raw data. The small shift toward the Republican Party that took place in the 1980s to the mid-1990s was overwhelmed by a shift back to the Democratic Party starting in the mid-1990s. That stabilized even before the nomination and inauguration of Barack Obama.

Latinos also follow the same pattern when we compare the data using the dyad ratio algorithm (see Figure 4) and the raw data (see Figure 2). We find a mean macropartisanship score of 68.58 for Latinos using the dyad ratio algorithm, while the raw data produced a mean of 68.49. We also see the same movements of Latinos over time, as well as the higher rate of volatility among Latinos compared with African Americans or non-Latino Whites.¹²

¹²The estimation report from the dyad ratio algorithm shows higher standard deviations for the Latino series than for the White or Black series. However, the dyad ratio algorithm’s estimation report also shows a higher eigen estimate for the Latino series (0.63) than the White series (0.60), while the Latino series also shows a higher percent of the variance explained (45.23) compared with the White series (42.47). Only the African American series has a higher eigen estimate (0.77) and percent of variance explained (54.88). These all speak to the validity of the measures despite the smaller sample sizes for Latinos and African Americans, as Stimson (2018) notes that mood indicators of policy preferences are typified by percent of variance explained in the high 30s.

Finally, we turn to the non-Latino White segment of the population. Again, the use of dyad ratio algorithm (see Figure 4) results in similar trends as when we plotted the raw data earlier (see Figure 2). The Democratic Party's advantage among the White population evaporated during the Reagan Administration and, with the exception of the latter years of the second Bush II Administration, the Republican Party held a consistent advantage. The macropartisanship score of 46.80 in the raw data is matched by a score of 46.90 in the monthly data using Stimson's algorithm. We can also see that non-Latino Whites do not necessarily follow the same patterns as the other groupings. While African Americans stuck with the Democratic Party during the Obama Administration and Latinos trended toward the Democrats, the White population moved toward the Republican Party. In sum, Stimson's dyad ratio algorithm produces substantially similar results to what was seen in the plots found in the previous section and allows us to develop monthly, quarterly, and annual measures of macropartisanship by race and ethnicity.¹³

4.3 Independents

One limitation of the preceding analysis is that it calculates macropartisanship as a relative measure of the two-party PID. This is interesting if the variation in those identifying as independent is stagnant. However, macropartisanship largely ignores the dealignment/realignment debate found in the literature (Dalton and Flanagan 2017; Hawley and Sagarzazu 2012; Lawrence 2018). Additionally, the volatility of Latino macropartisanship and the theoretical roots of that instability, we believe that this new dataset also unveils other interesting patterns worthy of further exploration. As has been argued by Hajnal and Lee (2011), independent party identification among groups with less rooted partisanship, particularly Latinos, reflects non-partisan or apolitical origins that come from a group feeling disenfranchised and disconnected from the political process. Therefore, beyond the relative strength of voting behavior for the Democratic party as captured by the share of the two-party PID, variation in independent identification may also give us an insight into macropartisanship among Latinos.

We begin with a presentation of dyad ratio estimates of independent identification for the overall sample (see Figure 5). Unlike changes in Democratic or Republican Party identification, or even macropartisanship, independent identification is remarkably stable in the overall sample. After an initial decline, independents consistently form a bit over one-third of the sample from 1987 for about 20 years, with a mean of 36.73%. However, starting around 2007, we start to see a sustained upward trend in independent identification. For the last 5 years of our data, we find independent identification averaged 41.87%, a five-point rise from the overall average.

Next, we present the dyad ratio estimates of independent identification for African Americans, Latinos, and non-Latino Whites in Figure 6. Here, we start with the non-Latino White independent identification, as it most closely follows the

¹³The measures are indeed distinct. We calculated the pairwise correlations in change in macropartisanship for Whites, Latinos and Blacks and found that Latino-White was .071, Latino-Black was .005, and Black-White was -0.050 . None of them reached standard levels of statistical significance.

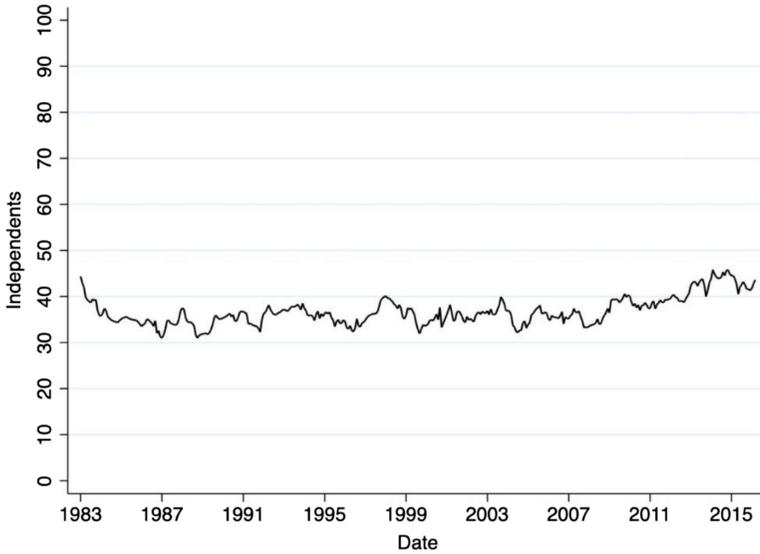


Figure 5. Overall Independent Identification using Stimson’s Dyad Ratio Algorithm. June 1983 to August 2016

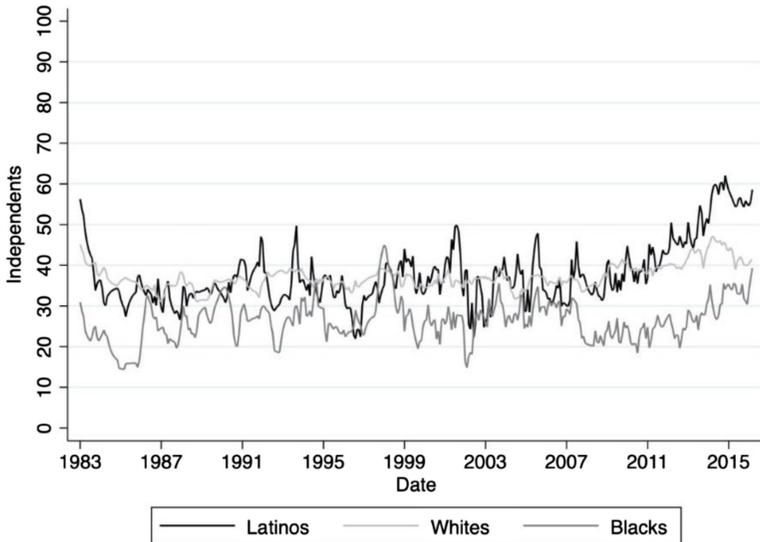


Figure 6. Independents by Race and Ethnicity using Stimson’s Dyad Ratio Algorithm. June 1983 to August 2016

overall trend, with an average of 37.21%. It also follows the initial decline, 20 years of stability, and then uptick overall the last decade of our data. The Black population consistently demonstrates lower independent identification, with an average of only 26.43%. They also demonstrate a sharp upward trend starting in around 2012, after

the uptick seen in the overall population and among non-Latino Whites. Turning to the Latino population, we see a fair amount of overlap with the White population, averaging 37.72%. More importantly, we see a sharp upward trend in independent identification starting around 2008, with more than 50% of Latinos identifying as independents in the final 3 years of the surveys. This provides *prima facie* evidence that dealignment is a substantial trend among Latino respondents in the last decade.

This decline in party identification in both major parties is hidden in macropartisanship data, and especially hidden without considering variation by race and ethnicity. To be sure, one limitation of our data is that we do not have information on the partisan leanings of independents over the study period. As has been well established (Keith et al. 1992), independents who lean toward the Democratic or Republican parties tend to have voting behavior patterns that look much more similar to those of weak party identifiers than they do to pure independents. Therefore, we do not wish to overstate these trends, as picking “leaning Independent” instead of “weak Partisan” can be viewed as a social desirability effect (Klar and Krupnikov 2016). Still, as Hajnal and Lee (2011) have argued, independent identification means something different for Latinos and we are especially interested by the fact that as Latino macropartisanship has trended toward the Democratic Party, we also observe a higher share of independent identification. While it is beyond the scope of this paper, we believe that this merits further investigation.

5. Validity

In this section, we seek to test the validity of our measures. We do this by (1) comparing the aggregated measure of macropartisanship to the measure constructed by Erikson et al. (2002), (2) comparing the disaggregated measure to the bi-annual data points from the American National Election Study archive,¹⁴ and (3) by validating that macropartisanship varies with model covariates from previous work (notably presidential approval and consumer confidence). In this last section, we can also observe how the disaggregated models for Whites, Latinos, and Blacks differ from the overall model.

5.1 The original measure

The most direct way to (1) assess the validity and (2) assess the utility of our new measure of macropartisanship is to compare it to the extended data presented by Erikson et al. (2002) available from Stimson’s (2017) website. While in the previous section we used monthly dyad ratio calculations, in this section we use quarterly dyad ratio calculations, as the MacKuen, Erikson, and Stimson’s dataset contains quarterly measures. In Figure 7, we see that our measure of macropartisanship tracks with MacKuen, Erikson, and Stimson’s original dataset.¹⁵ We do find that

¹⁴Data from the American National Election Studies.

¹⁵A simple correlation shows a robust relationship ($r=0.69$, $p<0.0001$), though perhaps not as strong as we might have expected. This is likely due to distinct methodologies in survey techniques and questions, as well as how MacKuen, Erikson, and Stimson built their measure.

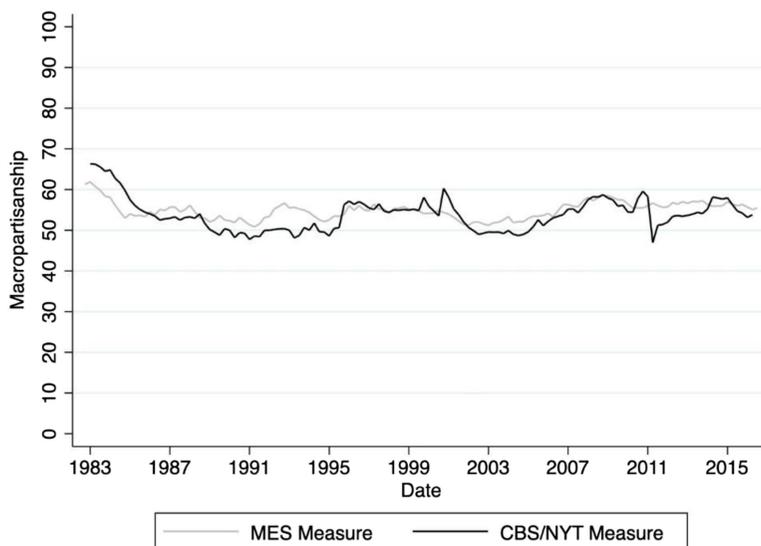


Figure 7. MacKuen, Erikson, and Stimson (MES) versus CBS/CBSNYTimes Quarterly Measures of Macropartisanship June 1983 to August 2016

MacKuen, Erikson, and Stimson’s measure is slightly more Democratic than our measure, with means of 54.82% and 53.84, respectively.¹⁶

In addition to comparing our overall measure of macropartisanship with the original measure created by MacKuen, Erikson, and Stimson, we also compare our measure of macropartisanship for non-Latino Whites with MacKuen, Erikson, and Stimson’s original measure. We compare the two because the literature on macropartisanship fails to account for variations by race and ethnicity and the traditional measure was developed when the White population accounted for a far larger proportion of the population. In Figure 8, we see that our measure of macropartisanship for Whites also tracks closely with MacKuen, Erikson, and Stimson’s data.¹⁷ In sum, our measure of macropartisanship based on CBS and CBS/NYTimes polls, as well as our subset of the non-Latino White population, closely tracks MacKuen, Erikson, and Stimson’s original measure.

¹⁶This measurement, difference, however, is to be expected as Erikson, MacKuen, and Stimson noted that in-person Gallup surveys were more Democratic than telephone surveys that Gallup moved to exclusively in the post-1991 era. To correct for this in the data, they adjusted the telephone surveys to conform to the in-person surveys, creating a pro-Democratic adjustment in the data. All of the CBS and CBS/New York Times surveys used here are from telephone surveys and no adjustments were made to the final data.

¹⁷As we saw with the overall measure of macropartisanship, we find MacKuen, Erikson, and Stimson’s measure and our measure of White macropartisanship are significantly correlated ($r = 0.56, p < 0.001$), while our measure of White macropartisanship also correlates with our overall measure at a very high rate ($r = 0.79, p < 0.001$).

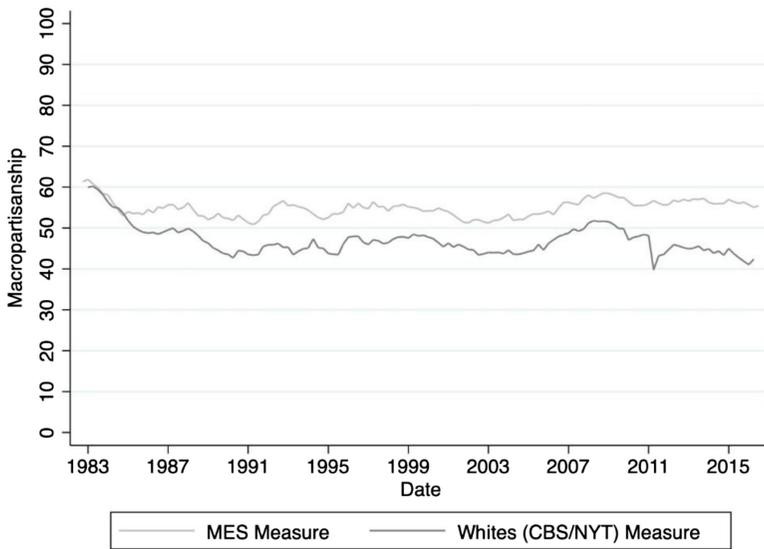


Figure 8. MacKuen, Erikson, and Stimson (MES) versus CBS/CBSNYTimes Quarterly Measures of White Macropartisanship June 1983 to August 2016

5.2 Validity by comparison with the ANES

A careful reader might wonder how we are certain that our disaggregated measures of macropartisanship are actually measuring what we assert they are measuring. Here, we might be concerned that data collection techniques, particularly for interviewing Latinos, were not as good in the 1980s and 1990s as they are today. While the CBS/NTY surveys did complete Spanish interviews, we know little about their early sampling procedures and can only assume that they are a far cry from what are considered best practices for reaching the Latino population today, particularly the Spanish-speaking Latino population.¹⁸ As a check on external validity, we check our yearly estimates of macropartisanship for White, Latinos and Blacks against existing data points available from the American National Election Studies.

The ANES is suitable for this task, as they have consistently taken biennial surveys of approximately 2000 Americans since the 1990s, with many interviews taking place in person for much of the data collection period. Given the ANES is only conducted every 2 or 4 years, here we used Stimson's dyad ratio algorithm to calculate annual measures of macropartisanship.¹⁹ Remarkably, the data track very well with our annual estimates from CBS News and CBS/New York Times surveys (see Figure 9). For African American respondents, we observe a slight decline in the 1980s and then a small but sustained upward trend identification in the Democratic Party. Even though we are looking at a very small sample size ($N = 14$), we see a positive and significant correlation ($r = 0.48$, $p = 0.08$).

¹⁸See for instance the procedures detailed from Latino Decisions (2020) on best practices for getting an accurate sample of the American Latino population.

¹⁹ANES data cover 1984–2016, with no data for 2006, 2010, or 2014.

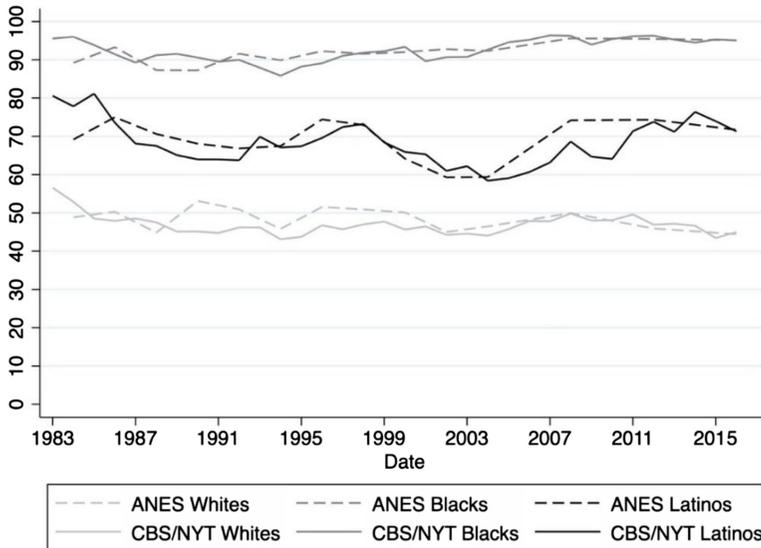


Figure 9. ANES versus CBS/CBSNYTimes Annual Measures of Macropartisanship 1983 to 2016

Similarly, Latinos show a dip in Democratic Party identification during the 1980s, a return in the Clinton years, another dip in the early part of the Bush II administration, and a sustained move toward the Democrats in both datasets. Again, despite the small sample size, we find a strong statistical correlation ($r = 0.78, p < 0.01$). Interestingly, we find similar levels of White identification in both datasets but find the weakest statistical relationship ($r = 0.28, p = 0.34$). Nevertheless, starting in the later years of the Clinton era, we see White macropartisanship track much more closely in both datasets, with a small upward trend in the Bush II administration and a slight downward trend in the Obama years. Overall, a comparison of our new dataset to the ANES data provides strong evidence of external validity, particularly for African American and Latino samples.

5.3 Disaggregated macropartisanship and the known predictors of macropartisanship

Erikson et al. (2002) establish a series of correlates of macropartisanship—notably presidential approval (direct) and economic indicators such as consumer confidence (indirect). As a means of establishing the external and construct validity of disaggregating macropartisanship by race and ethnicity, we examine the relationships between White, Latino, and Black macropartisanship and (1) presidential approval, (2) consumer confidence, (3) political polarization, and (4) presidential administration. Our examination here reveals five central observations. First, there is strong evidence that we are presenting a new measure of macropartisanship that is valid and performs like the previous measure presented by Erikson et al. (2002). Second, we find evidence of a more muted relationship between consumer confidence and macropartisanship, consistent with literature finding that partisan motivated

reasoning has subsumed economic voting (Duch et al. 2000). Third, it appears that there is a notable exception to this trend among Latinos, whose partisanship we have shown to be the least stable/most variable; Latino macropartisanship appears most responsive to economic conditions, while macropartisanship for Blacks, and possibly for Whites, is not.²⁰ Fourth, increased political polarization, particularly the Republican Party's sharp shift to the right, has moved Whites to the GOP and away from the Democratic Party. Finally, we also see mixed evidence of administration effects, with the greatest effect seeming to be a substantial shift among Latinos toward the Democratic Party in the Obama years. This shift appears to have occurred even when controlling for traditional factors and polarization.

First, we follow the methodology laid out in MacKuen et al. (1989) to calculate the mean deviates for macropartisanship. This is simply an observed value for macropartisanship minus the mean for macropartisanship. We calculated this for overall macropartisanship, for Latinos, for Whites, and for African Americans. A score above zero signifies that Democratic Party identification was above average for the observed time period, while a score below zero indicates Democratic identification was below average.

Second, we collected monthly presidential approval data from Gallup (UCSB 2018).²¹ Again, we calculated the mean deviate for presidential approval, so a higher score means a president is more popular than the overall average, while a negative score means a president is less popular. Next, we lag these scores by 1 month so that last month's presidential approval would predict this month's macropartisanship. We also multiply this measure of approval by -1 in the case of Republican presidents so that higher approval rates for Republican presidents should lead to lower Democratic Party identification and lower macropartisanship numbers.

Third, we gathered another of the original factors linked with changes in macropartisanship, consumer confidence. Consumer confidence is designed to capture the effect of the economy on changes in partisanship. A good economy will presumably lead to not only more popular presidents, but a shift in partisanship toward the incumbent party (Erikson et al. 1998; Green et al. 1998; MacKuen et al. 1989). The Consumer Confidence Index (CCI) is collected monthly via survey and surveys were lagged 1 month. Again, we calculated the mean deviate, so a score above zero indicates above average confidence, while a score below zero shows below average consumer confidence. Again, we multiply this economic measure by -1 for Republican administrations.

Fourth, we assembled data on political polarization. Since the advent of Nixon's Southern strategy, the parties have increasingly diverged and Congress has become increasingly polarized (Lewis 2020). This led to a long process of rebooting the American political process so that liberals identified as Democrats and conservatives

²⁰Given these findings related to consumer confidence, we also ran models with presidential approval excluded (see Online Appendix Table 1). Here, we find that consumer confidence is significant in most models, meaning that a good economy does appear to cause shifts in macropartisanship. However, we do find the effects appear much stronger for Latinos than for Whites, while the effects for African Americans are small and even statistically insignificant in the Newey–West models.

²¹Gallup data were gathered at The American Presidency Project (2021). These data usually provide multiple surveys per month, so we took the monthly average approval score. A few months did not produce estimates, so we used linear interpolation to fill in data for these months.

as Republicans (Levendusky 2009). At the elite level, this polarization was asymmetrical, with Republicans in Congress moving much further to the right than Democrats moved to the left. Similarly, Poole and Rosenthal's (1997) well-known claim regarding a two-dimensional spatial model, where the second dimension captured social issues, including race, evaporated during the period under study (Hare and Poole 2014; Voteviewblog 2017). Consequently, we calculate elite polarization using the median of first dimension NOMINATE estimates, where we take the median Republican score in a given congress and subtract the median Democratic score to create our measure of polarization (Lewis et al. 2021).²² Based on the polarization literature and the race/ethnicity literature, we expect Latinos, African Americans, and Whites will respond to this polarization accordingly, with Latinos and African Americans increasingly identifying with the Democratic Party and Whites moving to the Republican camp.

Finally, we added a series of presidential dummy variables for each administration included in our study: Reagan, Bush I, Clinton, Bush II, and Obama. Again, this harkens back to MacKuen et al. (1989) original model that showed substantial variation across administration. We decided not to include events dummies, given this analysis is simply designed to provide additional validation of our measures of macropartisanship.

We start with simple Pearson's correlation coefficients, presented in Table 1. We look to see if presidential approval or CCI are correlated with overall macropartisanship, or with macropartisanship broken down by race or ethnicity. First, we see that presidential approval is a reliable predictor of macropartisanship across all groups. However, we see that it is substantially weaker for Latinos and African Americans than it is for Whites.

We conducted both a Hotelling's T-Test and a Steiger's Z-Test to determine if any of these correlations were different from one another (Weiss 2011). We found that the relationship between White macropartisanship and presidential approval is significantly stronger than the relationship is for Latino macropartisanship or African American macropartisanship.²³ The White population simply seems to respond to presidential approval cues more than other groups. Second, we see that

²²If we use the mean, rather than median scores on the first dimension, we find substantively and statistically similar results. While Hare and Poole (2014) demonstrate, NOMINATE's second dimension evaporates during the period of our study and we double checked. We find a strong, positive correlation between first and second dimensions ($r=0.97$, $p<0.0001$) on our measure of polarization, so we did not include the second dimension in our analysis. Finally, we also ran analyses with Republican NOMINATE scores in one set of models and Democratic Nominate scores in another (see Online Appendix Tables 2 and 3). A positive coefficient indicates that as a party becomes more conservative the group became more Democratic, while a negative coefficient indicates the group became more Republican as the party moved to the right. In the table with only Republican NOMINATE scores, we see little change in the effect of presidential approval and consumer confidence, while the effect for Republican conservatism is only significant for Whites. As the Republican Party moved to the right, Whites moved away from the Democratic Party to the Republican Party (see Online Appendix Table 2). In the table with only Democratic NOMINATE scores, we also see little change in the effect of presidential approval and consumer confidence. However, we see a significant effect for Whites and African Americans in the Newey-West models. A more conservative Democratic Party made Whites more Democratic and Blacks less Democratic, which fits expectations. Further study of the effects of elite and mass divergence on issues of policy, particularly issues of race and ethnicity, is warranted in future studies.

²³See Appendix Tables 1 and 2 for the results of Hotelling's T-Test and Steiger's Z-Test.

Table 1. Presidential Approval, Consumer Confidence, and Macropartisanship

Macropartisanship	Presidential approval	Consumer confidence	Polarization
Latino	0.29***	0.09*	0.18***
White	0.42***	-0.04	-0.32***
African American	0.31***	-0.18***	0.59***
Overall	0.36***	0.01	0.16***

Pearson's correlation coefficients. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

CCI is not significant for Whites but is significant for Latinos and African Americans. For Latinos, the effect of a good economy is as predicted. Good economic times lead to shifts in macropartisanship toward the president's party, while bad times lead to shifts away. However, for African Americans, the reverse is true. Good economic times actually move people away from the president's party, while bad times move them toward the president. This unusual finding might be a function of who oversaw economic good times, or it might be a function of the distribution of gains during recent economic good times. Either way, this is worth additional investigation in future works. Third, we see that polarization has differential effects across racial and ethnic groups, with the correlation being positive for Latinos and African Americans and negative for Whites. As the parties diverged on policies over the last several decades, we see that African Americans moved even more sharply into the Democratic camp, Latinos moved somewhat to the Democratic camp, while Whites moved substantially toward the Republican camp.²⁴ Taken together, presidential approval provides additional evidence for the validity of our data, while consumer confidence introduces some questions.

However, Mackuen et al. (1989) argue that the effect of economic performance on macropartisanship is likely to be indirect, working through its connection to presidential approval. Additionally, we cannot discount the fact that (1) the importance of economic evaluations in macropartisanship may vary by race and (2) may be muted by considerations in individuals and the political environment (see Anderson 2007 for a review of this literature). To consider the effect of presidential approval, consumer confidence, and presidential administration, we present two types of analysis: (1) regression models with a lagged dependent variable (Achen 2000; Keele and Kelly 2006) and (2) Newey–West (Newey and West 1987) models with six lags (see Table 2). The use of a lagged dependent variable is quite common in time series models given the likelihood of serial autocorrelation. Newey–West estimators not only account for serial autocorrelation but also for heteroskedasticity that is common in time series data.²⁵ In these models, we include the variables for presidential approval, consumer confidence, and

²⁴These correlations are significantly different from one another (see Appendix Table 2).

²⁵The precise structure of the serial autocorrelation can be deduced by plotting autocorrelations and partial autocorrelations (see Online Appendix Figures 1 and 2). The first figure shows strong time serial components, with significant autocorrelations lasting over a year. However, the second figure shows that an AR1 structure largely captures concerns related to autocorrelation. When we conducted Durbin–Watson tests on the OLS models with the lagged dependent variable, we determined that only the model for Blacks produced residuals that violated an AR1 process ($p=0.06$). Still, we followed the advice of leading scholars

Table 2. Presidential approval, consumer confidence, political polarization, administration, and macropartisanship.

	Latinos	Whites	Blacks	Overall	Latinos	Whites	Blacks	Overall
	OLS	OLS	OLS	OLS	Newey–West	Newey–West	Newey–West	Newey–West
Approval _{t-1}	0.029*	0.021***	0.014**	0.017**	0.207***	0.145***	0.093***	0.119***
	(0.017)	(0.006)	(0.006)	(0.008)	(0.068)	(0.027)	(0.019)	(0.025)
Consumer index _{t-1}	0.042*	0.015*	0.007	0.018*	0.285***	0.050	-0.023	0.094***
	(0.023)	(0.008)	(0.008)	(0.010)	(0.109)	(0.031)	(0.017)	(0.032)
Polarization	4.289	-5.279*	0.670	0.151	-44.830	-41.686***	12.205*	-4.915
	(7.778)	(2.821)	(2.691)	(3.526)	(36.517)	(14.448)	(7.118)	(15.579)
Reagan	1.109	-0.209	-0.334	0.365	1.792	-0.229	-1.582	3.376
	(1.080)	(0.376)	(0.377)	(0.502)	(5.972)	(2.017)	(1.427)	(2.241)
Bush I	0.246	-0.916**	-0.546	-0.446	-6.149	-6.395***	-2.740**	-3.202
	(1.203)	(0.437)	(0.419)	(0.552)	(5.823)	(1.953)	(1.273)	(2.117)
Clinton	-0.033	-0.510**	-0.495**	-0.244	-0.727	-2.832***	-2.854***	-1.257
	(0.614)	(0.221)	(0.224)	(0.283)	(3.397)	(0.903)	(0.619)	(1.036)
Obama	0.934	0.316	0.150	0.483	14.954***	2.897**	0.176	4.263***
	(0.849)	(0.276)	(0.268)	(0.367)	(3.392)	(1.415)	(0.667)	(1.497)
Macropartisanship _{t-1}	0.889***	0.872***	0.865***	0.868***	-	-	-	-
	(0.023)	(0.022)	(0.025)	(0.023)				
Constant	-0.422	0.206	0.213	-0.068	-2.398	1.211*	1.298***	-0.707
	(0.428)	(0.150)	(0.151)	(0.196)	(2.325)	(0.718)	(0.479)	(0.756)
Observations	398	398	398	398	399	399	399	399
R-squared	0.863	0.912	0.891	0.871	-	-	-	-
F	-	-	-	-	7.79	10.97	31.60	13.86

Standard errors in parentheses.

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

polarization but also add in presidential administration dummies. We use the Bush II administration as the excluded category as its macropartisanship scores average closest to the mean.²⁶

These models produce interesting results. We see that presidential approval is positive and significant in all models. A popular president simply brings all racial and ethnic groups toward the president's party, while an unpopular president hurts the president's party. It is with economic evaluations and political polarization that differential effects appear. A strong economy matters for Latinos, may matter for Whites, but has no effect on African Americans. This is consistent with our expectations, as weaker Latino partisanship allows for a greater role for the economy. Similarly, whether it is the result of group identity or the skewed distribution of economic gains over the last several decades, African Americans appear unresponsive to economic ups and downs.

Our findings related to political polarization may be the most interesting. Polarization makes no difference for Latinos, clearly matters for Whites, and may also matter for Blacks. Once we controlled for other factors, the positive correlation between polarization and Latinos shifting to the Democratic Party evaporates. However, we see a strong, negative effect for Whites. As the associations between the Democratic Party and civil rights and the Republican Party and racial conservatism accelerated, Whites moved substantially away from the Democratic Party (see Carmines and Stimson 1989; Hutchings and Valentino 2004). The effect for African Americans is positive and significant, but only in the Newey–West model. This initial analysis cannot definitively resolve the effects of polarization on macropartisanship but does provide important insights that merit study in future works.

The administration dummies produce some interesting insights. When comparing administrations to the Bush II administration, we see that Whites and Blacks were a bit less Democratic during the Bush I and Clinton administrations, though the effect for Blacks is only significant in the Newey–West model. Interestingly, the Obama administration may have been a boon to Democrats among Latinos and Whites, but not among African Americans, with positive and significant coefficients in the Newey–West models. For Whites, it may just mean that Obama was able to make small gains among Whites despite the Republican Party's gains related to polarization. Again, this merits less exploratory analysis in future works.

6. Discussion

Macropartisanship is an aggregate measure of American partisanship that varies over time. But political science theories about macropartisanship—whether it is stable or variable, by how much, and what moves it—are caught up in a vision of American politics that thinks centrally about the political behavior of *White*

and also included Newey–West estimators to examine the robustness of our findings. While we included six lags in these models, even going out to 12 lags makes not statistical or substantive difference.

²⁶Based on the advice of a reviewer, we also ran ARMA (1, 1) models (see Online Appendix Table 4) to address possible errors-in-variables bias due to survey sampling error given the time series nature combined with the subdivision of the data by race and ethnicity (Beck 1991). Our analysis produced somewhat distinct findings regarding the effects of presidential approval and consumer confidence.

Americans. There are scholarly traditions which have defined different expectations for Black political behavior—whereby we might expect greater overtime partisan stability, and different expectations for Latino political behavior—whereby we might expect greater responsiveness to apolitical trends.

In this paper, using Stimson's (2018) dyad ratio algorithm to aggregate more than 500 CBS News and CBS/New York Times surveys from 1983 to 2016, we have presented a new measure of macropartisanship that can be disaggregated by race and ethnicity to create valid and reliable estimates of Black, Latino and non-Latino White macropartisanship. These data, which are available for download monthly, quarterly, or yearly, are presented as a free public download.

We view this paper as the first step to investigating the divergent trends in disaggregated macropartisanship, but indeed, the differential variation needs to be explored further. Can we connect theories of linked fate to the relative stability in Black macropartisanship? Is Latino macropartisanship, which exhibits considerably more overtime variation, reacting to protest events and changes in the partisan disposition toward issues like citizenship and immigration? Finally, should models of White macropartisanship account for the growing presence of White identity in recent work? These questions, all beyond the scope of this project, are all worthy of future investigation and research we hope will be spurred by this project and data release.

The exploratory analyses that are presented are driven by the theoretic traditions of the Black and Latino political behavior literatures. We find, consistent with expectations, that Black macropartisanship is remarkably stable. Interestingly, the overtime variation in Black, White, and Latino macropartisanship is predicted by variation in presidential approval, but evaluations of the economy, measured by consumer sentiment, matter mostly for Latinos. We also find that political polarization led Whites to abandon the Democratic Party for the Republican Party, while having a mixed effect on Blacks and no apparent effect on Latinos. These findings help put into context the importance of increasingly ideological extremism and White identity within the Republican Party (Jardina 2019).

However, as we observe Latino macropartisanship to be considerably more variable, we also find that the economy affects the swings in Latino macropartisanship to a greater degree. We also note that Latino macropartisanship, despite its variation, has trended toward a Democratic advantage since 2008. In the same era, however, we have also seen a dramatic rise in Latino independent identification, at a rate that far outpaces both Blacks and Whites. These findings are of particular note given the research in California that has found that Latinos moved away from the Republican Party in response to a nativist policy response from the Republican Party (Barreto et al. 2005; Bowler et al. 2006; Dyck et al. 2012; Korey and Lascher 2006, but see Hui and Sears 2018). Given the nativism surrounding the national Republican Party under Trump, it remains to be seen if we will see the emerging trend in Latino macropartisanship that favors the Democrats continue. While Democrats have made gains in the last decade in the data, the sharp rise in independent identification is equally worthy of investigation in future research.

Supplementary material. To view supplementary material for this article, please visit <https://doi.org/10.1017/rep.2021.35>

Acknowledgements. We are grateful for research assistance from Dan Herschel, Gabriel Carvajal, and Jack Cordell. Additionally, Shanna Pearson-Merkowitz, Ryan Carlin and Greg Love provided helpful advice related to using Stimson's dyad ratio algorithm in R. Finally, we appreciate the helpful advice our anonymous reviewers provided to strengthen the manuscript.

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Appendix A

Table A1. Testing Whether the Correlations between Presidential Approval and Macropartisanship Vary by Race and Ethnicity

	Hotelling's <i>T</i> -Test	Steiger's <i>Z</i> -Test
White vs. Black	2.110** (0.035)	2.100** (0.036)
White vs. Latino	2.541** (0.011)	2.523** (0.012)
Black vs. Latino	0.190 (0.849)	0.190 (0.849)

Hotelling's *T*-Test reports a *t*-statistic. Steiger's *Z*-Test reports a *Z*-statistic; *p*-values in parentheses. **p* < 0.10, ***p* < 0.05, ****p* < 0.01.

Table A2. Testing Whether the Correlations between Polarization and Macropartisanship Vary by Race and Ethnicity

	Hotelling's <i>T</i> -Test	Steiger's <i>Z</i> -Test
White vs. Black	-22.822*** (<0.001)	15.296*** (<0.001)
White vs. Latino	-10.057*** (<0.001)	8.974*** (<0.001)
Black vs. Latino	8.091*** (<0.001)	7.616*** (<0.001)

Hotelling's *T*-Test reports a *t*-statistic. Steiger's *Z*-Test reports a *Z*-statistic; *p*-values in parentheses. **p* < 0.10, ***p* < 0.05, ****p* < 0.01.

Cite this article: Dyck, JJ and GB Johnson (2022). Constructing a New Measure of Macropartisanship Disaggregated by Race and Ethnicity. *The Journal of Race, Ethnicity, and Politics* 7, 433–459. <https://doi.org/10.1017/rep.2021.35>