Bounding Partisan Approval Rates Under Endogenous Partisanship: Why High Presidential Partisan Approval May Not Be What It Seems

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Abstract: The presidential approval rate among a president's co-partisans has received a great deal of attention and is an important quantity for understanding accountability of the executive branch. Observed partisan approval rates may be biased when the composition of the president's party changes. We show that the composition of the president's party is endogenous to presidential popularity in Gallup polls, with the party growing and becoming more ideologically moderate as presidential popularity increases. We derive bounds on the compositionally-corrected partisan approval rate under a theoretically-motivated monotonicity condition. We examine how the bounds have evolved during the Obama and Trump presidencies. The proportion of survey respondents who identify with the Republican party has decreased rapidly from the pre-election benchmark during the Trump presidency and, as a result, the lower bound on Trump's partisan approval rate is much lower than at a comparable point in the Obama presidency.

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Introduction

In this short paper, we examine how compositional changes in public opinion polls affect estimation of the presidential approval rate among a president's co-partisans and employ Manski-style (2007) bounds to correct for these compositional changes. We show that the proportion of the president's co-partisans in a survey sample relative to a benchmark level is an essential (and typically ignored) component in bounding the partisan approval rate. By accounting for changes in the proportion of president co-partisans in a survey sample relative to an electorally-important benchmark level, we demonstrate that partisan approval may be lower than observed when fewer respondents report presidential co-partisanship or higher than observed when more respondents report presidential co-partisanship.

The partisan presidential approval rate has received a great deal of popular attention recently with many journalists and observers of American politics commenting on the high approval rate that President Trump enjoys among self-identified Republicans even while his aggregate approval rate is low by historic standards (Shepard 2017). Commentators have also noted that the partisan approval rate has implications for governance (Dropp and Nyhan 2017, Bump 2017, Wilkinson 2017). The partisan approval rate is substantively important as high presidential partisan approval rates may insulate the president from electoral and legislative accountability. Because committee chairs set the ground rules for investigations into the executive branch (Kriner and Schickler 2014), the presidential partisan approval rate is especially important under unified government when the president's party holds majorities in the House and Senate. Committee chairs may condition their willingness to hold the executive branch accountable for its actions on the president's approval rate among the legislator's partisan electorate. More broadly, co-partisans of the president seek to represent the views of partisan constituents either due to electoral or normative concerns regarding representation and these efforts have implications for the success or failure of a president's legislative agenda (Canes-Wrone and De Marchi 2002).

In cross-sectional surveys, estimating outcomes in the sample of respondents with a particular self-reported characteristic can result in a misleading estimate when self-reported characteristics are endogenous to external developments.[1] These considerations are particularly important in the context of the presidential partisan approval rate. As we show below, when presidents are losing support, the naive partisan approval rate can overestimate the approval rate among presidential co-partisans because of compositional change.

We first provide evidence that the probability that an individual respondent identifies with the president's party increases with presidential approval, but that the probability an individual identifies with the president's ideology is unresponsive to presidential approval. We then show how the observed partisan approval rate can be a misleading estimate when the sample composition changes over time and we construct Manski-style bounds for the compositionally-corrected partisan approval rate. We show that under a theoretically-motivated monotonicity condition, the true

[1] In recent years, polls such as the RAND American Life Panel (Gutsche et al. 2014) have employed high-frequency panels that allow researchers to examine how survey responses change among the same class of individuals over time. While these panels allow researchers to correct for compositional change by examining the outcome of interest for the same individuals over time, we focus on bounding outcomes under compositional change with cross-sectional polls, such as Gallup's presidential approval poll, because of their widespread use and the vast amount of popular attention that they receive.

compositionally-corrected partisan approval rate must lie within easily-calculable bounds. The bounds crucially depend on the proportion of survey respondents who identify with the president's party relative to a benchmark proportion. The upper bound is calculated by assuming that all missing (*resp.* excess) partisans approve (*resp.* disapprove) of the president while the lower bound assumes that the missing (*resp.* excess) partisans disapprove (*resp.* approve) of the president. These results emphasize the importance of accounting for the proportion of respondents in the president's party when interpreting observed partisan approval rates. We conclude by examining how the bounds have evolved over Obama's presidency and the first five months of the Trump presidency. As we show, the proportion of survey respondents who identify with the Republican party has decreased rapidly from the pre-election benchmark during the Trump presidency and, as a result, the lower bound on Trump's partisan approval rate is much lower than at a comparable point in the Obama presidency. Moreover, the magnitude of the difference between the lower bound and the observed level of support is much higher under Trump.

Our paper makes a methodological contribution to the study of compositional change in public opinion polls and has important implications for empirical studies that employ observed partisan approval rates as the dependent variable. Gelman et al. (2016) correct for endogenous survey response rates by using multi-level regression with poststratification where the poststratification is performed with information on respondents' partisanship and political characteristics, in addition to standard demographic variables, that are collected when respondents first enter into the panel. Our approach is a complementary method to account for changes in the composition of a survey, but instead of point identifying the compositionally-corrected outcome using an ignorability assumption on selection in the survey we derive bounds on this quantity under theoretically-motivated monotonicity conditions. On the applied side, there are numerous papers that examine how political and economic events affect partisan approval rates of the president (Hibbs, Rivers and Vasilatos 1982, Fox 2009, Newman and Siegle 2010). For example, Lebo and Cassino (2007) document an asymmetry in how same and opposing partisans' presidential approval responds to macroeconomic developments. Only opposing partisans change their approval evaluations of the president in reaction to macroeconomic changes. While Lebo and Cassino ascribe these dynamics to the psychological mechanism of motivated reasoning, a complementary explanation for their findings is that the sample composition of individuals who are willing to self-report as members of the president's party changes in response to macroeconomic events.

Other work in political science has employed Manski bounds for the purpose of set identifying a quantity of interest. Ashworth et al. (2008) bound the attributable risk of suicide terrorism due to military occupation in the presence of selection on the dependent variable. Wilkins (2012) bounds the probability that an incumbent U.S. House member losses reelection when retiring members' reelection outcomes are not observed. As a benchmark before considering more restrictive assumptions, Jackman (1999) uses data from the 1996 Australian Election Study to calculate the Manski bounds on the probability that an individual would vote if compulsory voting were eliminated by accounting for survey non-response. We contribute to this literature on political science bounding applications by considering the context when we are missing data on an individual's partisanship that would be observed under a different counterfactual state of the world. Our method is widely applicable in other settings where researchers are interested in the behavior of partisans or other groups whose composition may be endogenous.

Evidence of Endogenous Partisanship

Before proceeding to the analysis of how compositional changes affect partisan approval rates and the description of the bounding procedure, we first present evidence that the composition of the president's party is endogenous to presidential popularity using monthly Gallup polls from the Roper Center archives.[2] We employ the first poll of the month from January 2000 - June 2015, when the last Gallup poll in the Roper archive is reported, for our analysis. In the Online Appendix, we report summary statistics for these polls. Gallup polls may be especially sensitive to the problem of endogenous partisanship due to question ordering effects. Typically, the presidential approval question is the very first substantive[3] question in the survey (Sigelman 1981) and in all of the codebooks that we have reviewed this question is asked before the partisanship question.[4] While these dynamics may represent a real

[2]The effects that we document may be the result of respondents declining to complete the survey when they are unsatisfied with the president's performance or changing their self-reported partisanship in response to their assessment of the president's performance. We suspect that both mechanisms are driving the observed patterns in the data and are agnostic on the relative magnitudes of these two mechanisms. Regardless of whether survey non-response or changing item response is driving the results, the composition of individuals in the survey who identify as partisans of the president is endogenous and the observed partisan approval rate must be corrected to account for compositional change.

[3]Gallup respondents are asked a series of screening and, in the case of cell phone calls, safety questions, before the substantive portion of the survey.

[4]Randomizing the ordering of the partisanship and approval questions is a promising avenue for quantifying the extent to which the endogeneity we document is an artifact of question ordering.

effect of citizens becoming more inclined to identify as members of the president's party when the president is performing admirably in office, we can partially rule out this explanation by examining whether the probability that an individual identifies with the president's ideology changes in response to presidential approval.

To investigate these possibilities, we estimate linear probability models of the individual probability of identifying with the president's party and ideology[5] regressed on the previous month's average presidential approval rate[6] with each respondent weighted by the survey weight. The upper panel of Figure 1 reports the predicted probability of identifying as a member of the president's party and ideology as a function of the previous month's average presidential approval. The probability that a respondent identifies with the president's party is increasing with previous month's presidential approval while the ideology identification probability exhibits a much flatter relationship.[7]

The coefficient estimate on the regression of presidential party response on previous month presidential approval is approximately 0.17. Over the entire range of observed presidential approval, the probability that a respondent identifies with the president's party increases from a low of 0.411 to a high of 0.515. In the sample, the standard deviation of monthly presidential approval is 0.1219 so a one-standard deviation increase in presidential approval increases the probability of identifying with the president's party by approximately 0.021. In contrast, a one-standard deviation

[5] We include individuals who lean toward the president's party as identifying with the president's party and in the case of a Democratic (Republican) president, we classify both strong and weak liberals (conservatives) as identifying with the president's ideology.

[6] We downloaded these data from Gerhard Peters' American Presidency Project website.

[7] As we show in the Online Appendix, the coefficient estimates are statistically distinguishable from 0 in the partisanship regressions, but are statistically insignificant in the ideology regressions.

increase in presidential approval decreases the probability of identifying with the president's ideology by merely 0.006. These findings illustrate that self-reported partisanship is endogenous to the president's performance in office and that the pattern is not the result of citizens changing their political ideology in response to presidential performance.[8]

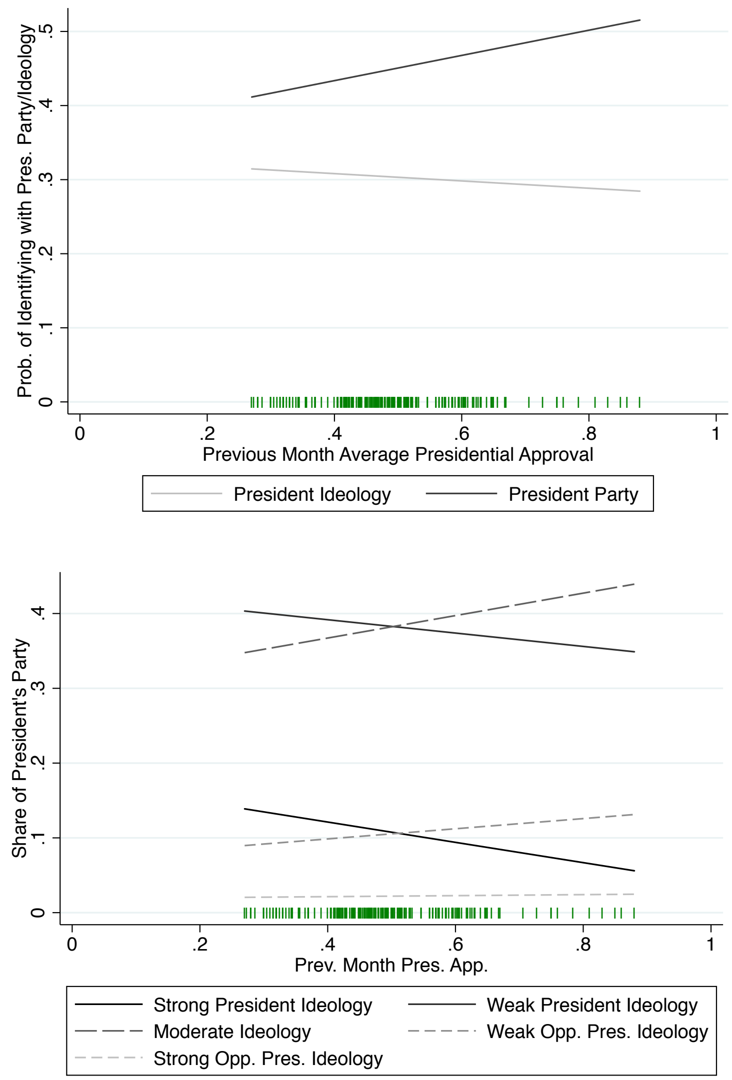
We next examine how the self-reported ideology of members of the president's party changes with presidential approval.[9] Under endogenous partisanship, we would expect that unpopular president's parties are mostly made up of the ideologically extreme. As the president becomes more popular, moderates and even ideological opponents begin to self-report as members of the president's party.

The lower panel of Figure 1 reveals how the ideological composition of the president's party evolves in response to presidential approval. Consistent with our expectations, individuals who strongly and weakly agree with the president's ideology become a smaller proportion of the president's party as the president becomes more popular. The largest estimated effect is the increase in the share of moderates in the president's party. The share of respondents with weakly opposing ideology in the president's party also increases, but the magnitude of this relationship is smaller than the moderate share relationship. The flat line revealing the change in the share of individuals with strongly opposed ideology shows that these ideologically oppo-

[8]The vast literature in political behavior showing that partisan attachments are durable, for example Green and Palmquist (1990), also suggests that the observed compositional changes are likely to be a survey artifact as opposed to a real change in national partisanship in response to the previous month's presidential performance.

[9]Individual respondents self-reported ideology may be endogenous to real-world developments and survey question effects, but our finding that the probability of identifying with the president's ideology does not respond to presidential approval reduces concern about the most serious potential endogeneity problem.

Figure 1: Probability of Identifying With President's Party and Ideology and Ideology Shares in President's Party as Function of Presidential Approval



site individuals are unlikely to be converted to presidential partisans even when the president is extremely popular.

In the Online Appendix, we show that these results hold when including individual-level demographic controls,[10] interacting these demographics with an indicator for a Democratic president, and only using within-president changes in presidential approval through the inclusion of president-specific fixed effects in the regressions. We account for the temporal dynamics of the data by aggregating the individual-level responses to one poll-level observation using a first-differenced dependent variable with one autoregressive and one moving average component in an ARIMA model. We also examine the sensitivity of the results to the use of unweighted regressions. The results are quite similar across these specifications.

While our analysis has focused on the endogeneity of partisan self-reports in cross-sectional surveys, conditioning on other attitudinal and political variables also has the potential to change the sample composition. For example, calculating approval rates for subsets of respondents based on past reported vote choice is subject to similar problems due to biased recall of voting history, which may depend on presidential popularity (Himmelweit, Biberian and Stockdale 1978). Given this evidence that the sample of respondents changes in response to presidential approval rates, we now turn to developing a bounding approach to account for the changing composition of survey respondents who identify with the president's party.

[10]We include indicators for state of residence and whether the respondent is female, African-American, Asian, Latino, multi-racial, Native American, Native Hawaiian, or white, has a high school degree or less, some college, college degree, or post graduate degree, is in the 18-34, 35-54, or 55 plus age categories. We also include income category indicators

Bounding Compositionally-Corrected Partisan Approval Rates

In order to derive our empirical bounds on the the partisan approval rate, we model poll respondents as being drawn from a finite set of types[11] and make one behavioral assumption about the likelihood of identifying as a party member, which we call the stable alignment assumption. We assume that as one type of respondent becomes more likely to identify as a presidential co-partisan then all types are weakly more likely to identify as presidential co-partisans. We make no other assumptions about the likelihood of approving of the president or the relationship between identity as a co-partisan and approval.[12] We believe that this monotonicity condition is relatively mild, but recognize that it could be violated in the event of a persistent partisan realignment.[13] This assumption also highlights that our method is most appropriate for making comparisons over relatively short periods of time, say between elections. It would not be appropriate for using the survey responses of Democrats in 2016 to reach conclusions about the counterfactual of how the Democratic party of the 1950s would approve of President Obama. In the Online Appendix, we show that under the stable alignment assumption the compositionally-corrected partisan approval rate that we define must lie within the bounds in expectation.

[11] These types characterize the probability that a member identifies with the president's party and approves of the president. The types could represent demographic or attitudinal groups or even individuals. As will become clear below, the exact definition of types is not essential for deriving bounds on the president's co-partisan approval rate.

[12] An alternative approach would impose stronger behavioral and functional form restrictions on the probability that an individual identifies with the president's party and approves of the president's performance in office to point identify the compositionally-corrected partisan approval rate. While such an exercise has some merit it requires stronger assumption that would be untestable and could lead to significant bias.

[13] In the presence of a persistent partisan realignment, we could calculate the bounds relative to the newly realigned electorate.

We define our bounds under two different empirical conditions: when there is a surplus of presidential co-partisans and when there is a deficit relative to a benchmark proportion. Let γ_P be the proportion of respondents that report presidential co-partisanship in the benchmark.[14] Let T be the total number of respondents in the survey and let P be the number reporting presidential co-partisanship. Each respondent reports either approval or disapproval.[15] Let the number of presidential co-partisans who report approval be P_A. The observed partisan approval that does not account for compositional changes is simply $\frac{P_A}{P}$.

If $\frac{P}{T} < \gamma_P$, then the proportion reporting presidential co-partisanship is lower in the survey than in the benchmark and we say that we have a deficit of presidential co-partisans equal to $T * \gamma_P - P$ respondents. To calculate the upper bound of compositionally stable partisan presidential approval, we account for the deficit of missing presidential co-partisans by assuming that each "missing" co-partisan approves so we add to both the numerator (number of approvers) and the denominator (number of presidential co-partisans) arriving at $\frac{P_A+(T*\gamma_P-P)\times 1}{P+(T*\gamma_P-P)} = \frac{P_A+(T*\gamma_P-P)}{T*\gamma_P}$. To derive the lower bound, we assume that each missing presidential co-partisan does not approve and only add to the denominator so that the co-partisan approval rate is $\frac{P_A+(T*\gamma_P-P)\times 0}{P+(T*\gamma_P-P)} = \frac{P_A}{T*\gamma_P}$.[16]

[14] In our empirical application, we use the proportion reporting co-partisanship with the presidential winner in the last available pre-election poll, but the analyst can employ any proportion.

[15] In our base case, we remove respondents who do not answer the approval question from the sample.

[16] To account for the fact that there may not be a sufficient number of non-presidential co-partisan approvers or non-approvers to relabel as co-partisans when there is a deficit of presidential co-partisans, we further adapt our bounds. Let A^{np} and D^{np} be the number of non-partisan approvers and non-approvers respectively. The upper bound accounting for the number of approvers is $\frac{P_A+Min\{(T*\gamma_P-P),A^{np}\}}{T*\gamma_P}$ and the lower bound is $\frac{P_A+Max\{0,(T*\gamma_P-P)-D^{np}\}}{T*\gamma_P}$. In practice these modifications rarely hold.

If $\frac{P}{T} > \gamma_P$, then the proportion reporting presidential co-partisanship is higher in the survey than in the benchmark and we say that we have a surplus of presidential co-partisans equal to $P - T * \gamma_P$ respondents. To calculate the upper bound of compositionally stable partisan presidential approval, we account for the surplus of presidential co-partisans by assuming that each "extra" co-partisan disapproves and so while we adjust the denominator down, we do not adjust the numerator down arriving at $\frac{P_A}{P-(P-T*\gamma_P)} = \frac{P_A}{T*\gamma_P}$. To derive the lower bound, we assume that each extra co-partisan does approve, and we adjust both the numerator and the denominator down, so that the co-partisan approval rate is $\frac{P_A-(P-T*\gamma_P)\times 1}{P-(P-T*\gamma_P)} = \frac{P_A-(P-T*\gamma_P)}{T*\gamma_P}$.[17]

We have defined our bounds in the simplest possible context, but the procedure can be extended to accommodate more complex situations. Survey weights can be incorporated by weighting the benchmark proportion of presidential co-partisans and approval responses. We can further extend the bounds to account for non-response to the approval question by assuming that all non-respondents approve of the president for the upper bound and that all non-respondents disapprove of the president for the lower bound. We could impose additional theoretical restrictions to tighten the bounds. For example, the analyst might be willing to assume that the probability a missing presidential co-partisan approves of the president is no greater than the probability that the average observed presidential co-partisan approves of the president and that the probability an extra presidential co-partisan approves of the president is no less than the probability that that the average observed presidential

[17]Similarly, we account for the fact that there may not be a sufficient number of presidential co-partisan approvers or non-approvers to relabel as non-partisans when there is a surplus of presidential co-partisans, we further adapt our bounds. The upper bound accounting for the number of approvers is $\frac{P_A - Max\{0,(P_A - T*\gamma_P)\}}{T*\gamma_P}$ and the lower bound is $\frac{P_A - Min\{(P-T*\gamma_P),P_A\}}{T*\gamma_P}$.

co-partisan approves of the president. These assumptions would tighten the upper bound to the observed approval rate when there is a deficit and the lower bound to the observed approval rate when there is a surplus.

Evolution of the Bounds During Obama and Trump Presidencies

We apply our bounding procedure to examine how the compositionally-corrected partisan approval rate has evolved over the course of the Obama and Trump presidencies using presidential approval polls from the Gallup Analytics web-based tool. Gallup Analytics reports cross-tabs for presidential approval by partisanship and the number of respondents identifying as Democrats, Republicans, and Independents every week.[18] We use all available presidential approval data from 2009-2017 for the analysis.[19]

We calculate the partisanship deficit or surplus relative to the partisanship proportions in the last poll available before the previous election.[20] We believe that

[18]The cross-tabs exclude partisan leaners from the two major parties and because we lack access to the underlying individual-level survey responses we cannot recreate partisanship proportions that include leaners.

[19]Gallup Analytics reports the presidential approval rate using their U.S. daily tracking poll. The lowest level of temporal aggregation that includes the party affiliation of respondents is the week. The first presidential approval results reported in Gallup Analytics are from the week of January 19-26, 2009 and the final week available when this analysis was conducted was June 26-July 2, 2017.

[20]For Obama's first term we use the final pre-election Gallup Poll fielded from October 31 - November 2, 2008 to calculate the proportion of respondents who identify as Democrats (excluding leaners) because Gallup Analytics presidential approval results are not available before January 19-26, 2009. For Obama's second term, we use the proportion of Democrats in the week of September 24, 2012 to September 30, 2012 polls. For Trump's presidency, we use the proportion of Republicans in the week of October 31 to November 6, 2016 polls. In all cases, we remove respondents who do not answer the presidential approval question from the sample before calculating these proportions.

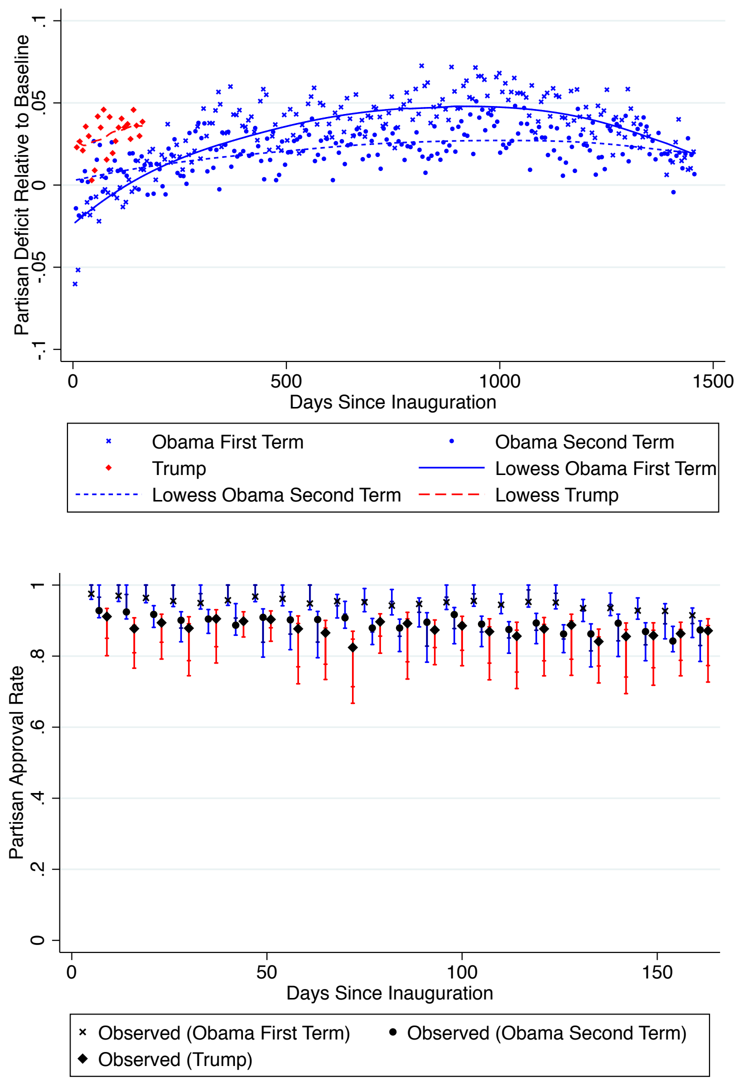
the last pre-election poll is the best measure of the size of the electorally-relevant coalition, but our approach is flexible enough to employ alternative benchmarks in different contexts. We use unweighted partisanship proportions because Gallup Analytics reports the raw number of Democrats, Republicans, and Independents in the sample as opposed to a weighted quantity. To calculate standard errors, we bootstrap the upper and lower bound using 200 bootstrap replications to compute the observed standard deviation of the upper and lower bounds across the bootstrap replications.[21]

In the upper panel of Figure 2, we plot the partisan deficit relative to the pre-election baseline partisanship proportions against the number of days since the inauguration for the two terms of the Obama presidency and the first five months of the Trump presidency. We also include separate lowess plots[22] for each presidential term. The most striking finding in the plot is that the deficit of president partisans during the Trump presidency is very high compared to the analogous period in Obama's first term and is even higher than during the comparable period in Obama's second term. The Trump presidency lowess curve has already reached a higher level than at any point during Obama's second term and is approaching the partisan deficit highwater mark in 2011. The smoothed partisan deficit 149 days into the Trump presidency is at the level reached 509 days into Obama's first term and it is never reached during

[21]The lack of individual-level data requires us to construct pseudo-individual level data when conducting the bootstrap resampling. For example, if there are 1,000 independents in a poll and 450 approve of the president and 550 disapprove, the pseudo-individual level data for the independents consists of 1,000 observations with 450 observations set equal to 1 for approval and 550 observations with 0s for the approval variable.

[22]We use local linear regression, a bandwidth of 0.8, and weight the observations using the tri-cube function to create the lowess curves.

Figure 2: Gallup Analytics Partisan Proportions and Partisan Approval Rates



Obama's second term. While we cannot predict how the partisan deficit will evolve in the future, the high partisan deficit does have important implications, which have been ignored by popular commentators, for interpreting President Trump's partisan approval rate.

In the lower panel of Figure 2, we plot observed partisan approval rates, the bounds on the compositionally-corrected partisan approval rate, and 95 percent confidence intervals for the upper and lower bound during the first 163 days of each presidential term.[23] The marker is the observed partisan approval rate, the first capped line extending out from the marker are the bounds, and the second set of capped lines are the 95 percent confidence intervals on the lower and upper bounds.[24]

Trump's observed partisan approval rates are very low compared to the same period during Obama's first term, but are roughly comparable to Obama's second term. More relevant for our analysis is how the bounds evolve over time. The lower bound on the compositionally-corrected partisan approval rate is quite low during Trump's presidency. In 14 of the 23 weeks, the lower bound is below 0.8. With only one exception, the lower bound on Trump's compositionally-corrected partisan approval rate is lower than the lower bound from the analogous poll during Obama's second term. The observed partisan approval rate is partially an artifact of missing respondents who would have previously reported Republican partisanship. While President Trump's observed partisan approval rate has received much attention, the data are also consistent with the possibility that his partisan approval rate is quite low relative to recent presidential history.

[23] The final poll from our sample is 163 days after Trump's inauguration so we restrict attention to the in order to make the visual comparison as uncluttered as possible. In the Online Appendix, we report a graph that shows the entire history of the observed partisan approval rates, bounds, and confidence intervals over the complete 2009-2017 period.

[24] In some instances, the standard errors are so small that the confidence intervals are visually indistinguishable from the point estimate for the bounds.

Conclusion

We have shown that self-reported partisanship is endogenous to presidential approval in Gallup polls and derived bounds on the compositionally-corrected partisan approval rate under a relatively mild and theoretically-justified monotonicity condition. We also documented how the bounds on the compositionally-corrected partisan approval rate have evolved over Obama's presidency and Trump's first five months in office. The lower bounds on the compositionally-corrected partisan approval rates are much lower than at equivalent points in Obama's first and second terms.

While we have framed our discussion in terms of endogenous partisanship and compositionally-corrected presidential partisan approval rates, there are many other potential applications for our approach. Whenever analysts calculate a quantity of interest for a particular attitudinal or political subpopulation there is the potential that individuals with that characteristic will decline to answer the survey or change their responses on the self-reported characteristic of interest. Using information on the proportion of respondents in the survey relative to their baseline proportion can be used to bound the unknown compositionally-corrected quantity of interest.

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Online Appendix A

In this appendix, we define our quantify of interest and establish that it is contained in our empirical identification region. Let each respondent be represented by a type $\tau \in \{1, 2, 3...\overline{\tau}\}$. We assume that each respondent type has a propensity for identifying with the party of the president that depends on a underlying state of the world $\omega \in \Omega \subseteq \mathcal{R}$ and denote this propensity by $\rho(\tau, \omega)$ and note that it lies in $[0, 1]$. We label states of the world, such that a higher ω is associated with higher probabilities of reporting presidential co-partisanship (PCP) for all respondent types. Factors that could affect ω include incumbent performance, media coverage, the performance of the economy, elections, or world events. Our key assumption is that, conditional on our labeling of states of the world, that higher states of the world are associated with a weakly higher probability of reporting PCP for all respondent types. Substantively, this assumption rules out states associated with a wholesale realignment of the political environment where, for example, those who formerly self-reported as Republican become less likely to report being Republican while those that previously identified as Democrats become more likely to identify as Republican. While such a realignment may occur over time, we limit our analysis to a single presidential term following an election where we believe the assumption is likely to hold. In any case, any method that seeks to account for compositional change would need to make a similar assumption in addition to stronger, less theoretically-motivated assumptions.

We call this assumption the *Stable Assignment Assumption* and state it formally:

Stable Alignment Assumption. *If $\omega' > \omega$ then $\rho(\tau, \omega') \geq \rho(\tau, \omega)$ for all $\tau \in 1, 2, 3...\overline{\tau}$ and $\rho(\tau, \omega') > \rho(\tau, \omega)$ for at least one $\tau \in 1, 2, 3...\overline{\tau}$.*

In addition to a probability of aligning with the president's party each respondent has a propensity for reporting approval of the president that can vary with their type or the state of the world: $\alpha(\tau, \omega) \in [0, 1]$. We allow this probability to vary arbitrarily with respondent type and state and allow for any cross-dependence between state and type.

In what follows, we abstract away from sampling variation that might change the distribution of types in the survey. For simplicity of presentation, we assume that the distribution of types is fixed and equal to $p(\tau)$ where $p(\tau) \geq 0$ for all $\tau \in 1, 2, 3...\overline{\tau}$ and $\sum_{\tau=1}^{\overline{\tau}} p(\tau) = 1$. Sample variability is a well studied issue and we follow Manski (2007) by bootstrapping standard errors for all our empirical analogues by recalculating the bounds for each bootstrap replication and computing the standard deviation across the bootstrap replications.

We can now calculate the proportion of respondents who report PCP and the approval rate among those reporting PCP. Let $\mathbb{P}(\omega)$ be the proportion of respondents reporting PCP when the state is ω. Formally:

$$\mathbb{P}(\omega) = \sum_{\tau=1}^{\overline{\tau}} p(\tau)\rho(\tau, \omega)$$

Of particular interest is the set of respondents that report PCP at a benchmark point in time. Let the state during the election be ω_0 and call those that report PCP in state ω_0 *core supporters*. The proportion of core supporters is a fixed quantity of interest and is equal to $\mathbb{P}(\omega_0)$ and the probability that a respondent of type τ is a core supporter is $\rho(\tau, \omega_0)$.

In addition to the proportion of respondents that report PCP, we can calculate the average level of support among those that report PCP. Let $\tilde{\rho}(\tau,\omega) = \frac{\rho(\tau,\omega)}{\sum_{t=1}^{\overline{\tau}} \rho(\tau,\omega)}$ be the probability mass function for those that report PCP given ω and let $IS(\omega)$ be the average approval rate among them. Formally:

$$IS(\omega) = \sum_{\tau=1}^{\overline{\tau}} \alpha(\tau,\omega)p(\tau)\tilde{\rho}(\tau,\omega)$$

While $IS(\omega)$ is observable in surveys, it is subject to change for two distinct reasons. First, it may change as the rate of approval for each type τ changes. Second, it may change due to compositional changes as the rate of PCP changes. We are interested in changes in approval that only reflected changes in the state and not changes in partisan composition. To account for the compositional changes, we define our quantity of interest, $CS(\omega)$, to be the average level of support among *core supporters*, the group where the probability of reporting PCP is fixed at the level for state ω_0. Formally:

$$CS(\omega) = \sum_{\tau=1}^{\overline{\tau}} \alpha(\tau,\omega)p(\tau)\tilde{\rho}(\tau,\omega_0)$$

To see the relationship between our quantity of interest $CS(\omega)$ and the observable $IS(\omega)$, we introduce some more notation. Let $\delta(\tau,\omega) = |\rho(\tau,\omega) - \rho(\tau,\omega_0)|$ be the absolute difference between the probability that a person reports PCP in state ω and the probability that a person reports PCP in in state ω_0. Given our Stable Alignment Assumption, $\delta(\tau,\omega) = \rho(\tau,\omega) - \rho(\tau,\omega_0)$ if and only if $\omega > \omega_0$, otherwise $\delta(\tau,\omega) = \rho(\tau,\omega_0) - \rho(\tau,\omega)$. Let $\Delta(\omega) = |\mathbb{P}(\omega) - \mathbb{P}(\omega_0)|$ be the difference in proportion between respondents that report PCP and those that are core supporters.

Case I: $\omega > \omega_0$

Let us first consider the case where $\omega > \omega_0$, so that $\Delta(\omega) = \mathbb{P}(\omega) - \mathbb{P}(\omega_0)$ and $\delta(\tau, \omega) = \rho(\tau, \omega) - \rho(\tau, \omega_0)$. The set of PCP responders includes all core supporters, but additionally contains respondents who are not core supporters. Let $\tilde{\delta}(\tau, \omega) = \frac{\delta(\tau,\omega)}{\sum_{t=1}^{\overline{\tau}} \delta(\tau,\omega)}$ be the probability mass function for the additional PCP respondents given ω and let $DS(\omega)$ be average approval among these additional supporters. Formally:

$$DS(\omega) = \sum_{\tau=1}^{\overline{\tau}} \alpha(\tau, \omega) p(\tau) \tilde{\delta}(\tau, \omega)$$

We can now decompose $IS(\omega)$ into a mixture of core supporter approval and additional supporter approval:

$$IS(\omega) = \frac{\mathbb{P}(\omega_0)}{\mathbb{P}(\omega)} CS(\omega) + \frac{\Delta(\omega)}{\mathbb{P}(\omega)} DS(\omega)$$

Re-arranging, we can express our quantity of interest, $CS(\omega)$ in terms of observable quantities, $IS(\omega)$, $\mathbb{P}(\omega)$, $\mathbb{P}(\omega_0)$, and $\Delta(\omega)$ and an unknown, but bounded quantity, $DS(\omega)$:

$$CS(\omega) = \frac{\mathbb{P}(\omega)}{\mathbb{P}(\omega_0)} IS(\omega) - \frac{\Delta(\omega)}{\mathbb{P}(\omega_0)} DS(\omega)$$

Simple algebra gives us:

$$CS(\omega) = \frac{\mathbb{P}(\omega) IS(\omega) - \Delta(\omega) DS(\omega)}{\mathbb{P}(\omega) - \Delta(\omega)}$$

While we cannot observe the approval rate among the excess respondents that report PCP, we know that their approval rate is bounded between 0 and 1, thus we can construct a theoretical upper bound $\overline{CS}^{surplus}(\omega)$ by assuming that each additional PCP identifier does not approve and construct a theoretical lower bound $\underline{CS}^{surplus}(\omega)$ by assuming each additional PCP identifier does approve:

$$\overline{CS}^{surplus}(\omega) = \frac{\mathbb{P}(\omega)IS(\omega)}{\mathbb{P}(\omega) - \Delta(\omega)}$$

and

$$\underline{CS}^{surplus}(\omega) = \frac{\mathbb{P}(\omega)IS(\omega) - \Delta(\omega) \times 1}{\mathbb{P}(\omega) - \Delta(\omega)}$$

Clearly, $CS(\omega) \in [\underline{CS}(\omega), \overline{CS}(\omega)]$ and so we arrive at our main proposition for the case where there is a surplus (relative to the election) of respondents that report PCP.

Proposition 1. *Under the stable alignment assumption, if $\omega > \omega_O$ then $CS(\omega) \in [\underline{CS}^{surplus}(\omega), \overline{CS}^{surplus}(\omega)]$*

Case II: $\omega < \omega_0$

We now consider the case where $\omega < \omega_0$, so that $\Delta(\omega) = \mathbb{P}(\omega_0) - \mathbb{P}(\omega)$ and $\delta(\tau, \omega) = \rho(\tau, \omega_0) - \rho(\tau, \omega)$. The set of core supporters includes all those that report PCP, but additionally contains respondents that do not report PCP. Now, $DS(\omega)$ defined above, is the average approval among these missing core supporters.

We can decompose our quantity of interest, $CS(\omega)$, into a mixture of the observable approval rate among those that report PCP, $IS(\omega)$, the observable weights

$\mathbb{P}(\omega)$, $\mathbb{P}(\omega_0)$, and $\Delta(\omega)$ and the unknown, but bounded quantity, $DS(\omega)$, which is the approval rate among missing core supporters:

$$CS(\omega) = \frac{\mathbb{P}(\omega)}{\mathbb{P}(\omega_0)} IS(\omega) + \frac{\Delta(\omega)}{\mathbb{P}(\omega_0)} DS(\omega)$$

Simple algebra gives us:

$$CS(\omega) = \frac{\mathbb{P}(\omega)IS(\omega) + \Delta(\omega)DS(\omega)}{\mathbb{P}(\omega) + \Delta(\omega)}$$

While we cannot observe the approval rate among the missing core supporters, we know that their approval rate is bounded between 0 and 1, thus we can construct a theoretical upper bound $\overline{CS}^{deficit}(\omega)$ by assuming that each additional missing core supporter does approve and construct a theoretical lower bound $\underline{CS}^{deficit}(\omega)$ by assuming each additional missing core supporter does not approve:

$$\overline{CS}(\omega)^{deficit} = \frac{\mathbb{P}(\omega)IS(\omega) + \Delta(\omega) \times 1}{\mathbb{P}(\omega) + \Delta(\omega)}$$

and

$$\underline{CS}(\omega)^{deficit} = \frac{\mathbb{P}(\omega)IS(\omega)}{\mathbb{P}(\omega) + \Delta(\omega)}$$

Clearly, $CS(\omega)^{deficit} \in [\underline{CS}^{deficit}(\omega), \overline{CS}^{deficit}(\omega)]$ and so we arrive at our second main proposition.

Proposition 2. *Under the stable alignment assumption, if $\omega < \omega_O$ then $CS(\omega) \in [\underline{CS}^{deficit}(\omega), \overline{CS}^{deficit}(\omega)]$*

To transition between our theoretical bounds and our empirical bounds, we describe our empirical analogues. Our empirical measure of the proportion of core supporters, $\mathbb{P}(\omega_0)$, is the percentage of respondents reporting being partisans with the party of the eventual winner of the presidential election in a benchmark period. Our empirical measure of the proportion of respondents reporting PCP in any given poll, $\mathbb{P}(\omega)$, is the percentage of respondents reporting presidential co-partisanship. Finally, our empirical measure of PCP respondent support, $IS(\omega)$, is simply the approval rate among those reporting partisanship with the president. Which set of bounds we use depends on whether ω is greater or less than ω_0. Under the stable alignment assumption, $\omega > \omega_0$ if and only $\mathbb{P}(\omega) > \mathbb{P}(\omega_0)$ or empirically when the proportion of strong presidential partisans is greater than in the pre-election poll. Thus, when the proportion of strong presidential partisans is greater than in the pre-election poll, we apply the bounds from proposition 1 otherwise we apply the bounds from proposition 2.

Online Appendix B

Table A.1: Summary statistics

Variable	**Mean**	**Std. Dev.**	**N**
Previous Month Average Presidential Approval	0.486	0.12	191942
President Party	0.45	0.497	188740
President Ideology	0.305	0.46	186113
President Strong Ideology	0.062	0.242	186113
President Weak Ideology	0.243	0.429	186113
Moderate Ideology	0.388	0.487	186113
Opp. President Weak Ideology	0.237	0.426	186113
Opp. President Strong Ideology	0.07	0.255	186113
Female	0.507	0.5	188839
African-American	0.076	0.265	189044
Asian	0.01	0.097	189044
Latino	0.048	0.214	189044
Multi-Racial	0.033	0.178	189044
Native American	0.006	0.078	189044
Native Hawaiian	0.014	0.119	189044
White	0.813	0.39	189044
HS or Less	0.144	0.351	190393
Some College	0.264	0.441	190393
College Graduate	0.266	0.442	190393
Post Graduate	0.326	0.469	190393
Age 18-34	0.168	0.374	172786
Age 35-54	0.377	0.485	172786
Age 55+	0.455	0.498	172786
Income < 10K	0.04	0.196	149499
Income 6-12K	0.012	0.107	149499
Income 10-14K	0.062	0.241	149499
Income 12-24K	0.031	0.172	149499
Income 15-20K	0.087	0.282	149499
Income 20-30K	0.109	0.311	149499
Income 24-36K	0.036	0.187	149499
Income 30-50K	0.163	0.369	149499
Income 48-60K	0.03	0.169	149499
Income 50-75K	0.166	0.372	149499
Income 60-90K	0.043	0.203	149499
Income > 75K	0.222	0.415	149499

Table A.2: Prob. of Identifying With President's Party and Ideology as a Function of Presidential Approval

	(1) Pres. Party	(2) Pres. Party	(3) Pres. Party	(4) Pres. Ideology	(5) Pres. Ideology	(6) Pres. Ideology
Prev. Month Pres. App.	0.170***	0.219***	0.204***	-0.0493	-0.0401	-0.0190
	(0.0215)	(0.0265)	(0.0189)	(0.0576)	(0.0448)	(0.0164)
Observations	188,740	130,475	130,475	186,113	128,854	128,854
Controls	No	Yes	Yes	No	Yes	Yes
Controls X Dem. Pres.	No	No	Yes	No	No	Yes
President FEs	No	No	Yes	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.3: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party

	(1) Strong Pres. Ideology	(2) Strong Pres. Ideology	(3) Strong Pres. Ideology
Prev. Month Pres. App.	-0.142***	-0.170***	-0.161***
	(0.0186)	(0.0163)	(0.0161)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.4: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party

	(1) Weak Pres. Ideology	(2) Weak Pres. Ideology	(3) Weak Pres. Ideology
Prev. Month Pres. App.	-0.101	-0.0764	-0.0575*
	(0.0723)	(0.0485)	(0.0240)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.5: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party

	(1) Moderate Ideology	(2) Moderate Ideology	(3) Moderate Ideology
Prev. Month Pres. App.	0.154***	0.172***	0.156***
	(0.0398)	(0.0345)	(0.0246)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.6: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party

	(1) Weak Opp. Pres. Ideology	(2) Weak Opp. Pres. Ideology	(3) Weak Opp. Pres. Ideology
Prev. Month Pres. App.	0.0814^{+}	0.0669^{*}	0.0534^{***}
	(0.0435)	(0.0271)	(0.0114)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.7: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party

	(1) Strong Opp. Pres. Ideology	(2) Strong Opp. Pres. Ideology	(3) Strong Opp. Pres. Ideology
Prev. Month Pres. App.	0.00779	0.00737	0.00923^{+}
	(0.00769)	(0.00512)	(0.00506)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.8: ARIMA(1,1,1) Time-Series Regression of Share of President's Party/Ideology on Previous Month's Presidential Approval

	(1) Pres. Party	(2) Pres. Ideology
main		
D.Prev. Month Pres. App.	0.198^{***}	-0.0737
	(0.0517)	(0.123)
Observations	177	177
President FEs	Yes	Yes

Heteroskedasticity robust standard errors in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.9: ARIMA(1,1,1) Time-Series Regression of Ideological Composition of President's Party on Previous Month's Presidential Approval

	(1) St. Pres	(2) Weak Pres	(3) Mod.	(4) Weak Opp. Pres	(5) St. Opp. Pres
main					
D.Prev. Month. Avg. Pres. App	-0.104***	-0.156	0.150**	0.0958	0.00618
	(0.0242)	(0.184)	(0.0555)	(0.110)	(0.00980)
Observations	177	177	177	177	177
President FEs	Yes	Yes	Yes	Yes	Yes

Heteroskedasticity robust standard errors in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.10: Prob. of Identifying With President's Party and Ideology as a Function of Presidential Approval (Unweighted)

	(1) Pres. Party	(2) Pres. Party	(3) Pres. Party	(4) Pres. Ideology	(5) Pres. Ideology	(6) Pres. Ideology
Prev. Month Pres. App.	0.157***	0.208***	0.199***	-0.0771	-0.0674	-0.0291*
	(0.0129)	(0.0191)	(0.0168)	(0.0616)	(0.0489)	(0.0145)
Observations	188,740	130,475	130,475	186,113	128,854	128,854
Controls	No	Yes	Yes	No	Yes	Yes
Controls X Dem. Pres.	No	No	Yes	No	No	Yes
President FEs	No	No	Yes	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.11: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party (Unweighted)

	(1) Strong Pres. Ideology	(2) Strong Pres. Ideology	(3) Strong Pres. Ideology
Prev. Month Pres. App.	-0.153***	-0.175***	-0.158***
	(0.0171)	(0.0143)	(0.0136)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.12: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party (Unweighted)

	(1) Weak Pres. Ideology	(2) Weak Pres. Ideology	(3) Weak Pres. Ideology
Prev. Month Pres. App.	-0.129^{+}	-0.0976^{+}	-0.0707***
	(0.0730)	(0.0519)	(0.0201)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.13: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party (Unweighted)

	(1) Moderate Ideology	(2) Moderate Ideology	(3) Moderate Ideology
Prev. Month Pres. App.	0.191***	0.196***	0.171***
	(0.0440)	(0.0377)	(0.0220)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.14: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party (Unweighted)

	(1) Weak Opp. Pres. Ideology	(2) Weak Opp. Pres. Ideology	(3) Weak Opp. Pres. Ideology
Prev. Month Pres. App.	0.0839*	0.0704**	0.0528***
	(0.0397)	(0.0250)	(0.00900)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.15: Prob. of Ideology Identification as a Function of Presidential Approval Conditional on Identifying with President's Party (Unweighted)

	(1) Strong Opp. Pres. Ideology	(2) Strong Opp. Pres. Ideology	(3) Strong Opp. Pres. Ideology
Prev. Month Pres. App.	0.00750	0.00605	0.00514
	(0.00588)	(0.00383)	(0.00383)
Observations	83,451	58,262	58,262
Controls	No	No	Yes
Controls X Dem. Pres.	No	Yes	Yes
President FEs	No	No	Yes

Heteroskedasticity robust standard errors clustered by poll in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.16: ARIMA(1,1,1) Time-Series Regression of Share of President's Party/Ideology on Previous Month's Presidential Approval(Unweighted)

	(1) Pres. Party	(2) Pres. Ideology
main		
D.Prev. Month Pres. App.	0.188***	-0.114
	(0.0363)	(0.144)
Observations	177	177
President FEs	Yes	Yes

Heteroskedasticity robust standard errors in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.17: ARIMA(1,1,1) Time-Series Regression of Ideological Composition of President's Party on Previous Month's Presidential Approval (Unweighted)

	(1) St. Pres	(2) Weak Pres	(3) Mod.	(4) Weak Opp. Pres	(5) St. Opp. Pres
main					
D.Prev. Month. Avg. Pres. App	-0.0961***	-0.183	0.159**	0.119	0.0124^{+}
	(0.0250)	(0.198)	(0.0606)	(0.122)	(0.00700)
Observations	177	177	177	177	177
President FEs	Yes	Yes	Yes	Yes	Yes

Heteroskedasticity robust standard errors in parentheses.
$^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Figure A.1: Observed Partisan Approval Rate, Bounds, and Confidence Intervals for Full Period