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

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The presidential approval rate among a president's copartisans has received a great deal of attention and is an important quantity for understanding accountability of the executive branch. We show that the reported composition of the president's party is endogenous to presidential popularity, with the party growing and becoming more ideologically moderate as presidential popularity increases. As a result, observed partisan approval rates may be biased because of compositional change in respondents who self-identify with the president's party. 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 preelection 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.

In this short article, we examine how compositional changes in public opinion polls affect estimation of the presidential approval rate among a president's copartisans and employ Manski (2007) style bounds to correct for these compositional changes. We document that the composition of self-reported partisans changes in response to the president's popularity. As a result, observed partisan approval rates may be biased because of changes in the composition of survey respondents who self-identify with the president's party. We show that the proportion of the president's copartisans 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 copartisans in a survey sample relative to an electorally important benchmark level, we demonstrate that partisan approval may be meaningfully lower than observed when fewer respondents report presidential copartisanship or higher than observed when more respondents report presidential copartisanship.

The partisan presidential approval rate has received a great deal of popular attention recently with many 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). 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

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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 legislators' partisan electorate. More broadly, copartisans of the president seek to represent the views of partisan constituents due to either 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] This endogeneity may be the result of respondents either declining to complete the survey when they are unsatisfied with the president's performance (Gelman et al. 2016; Hartman 2018) or changing their self-reported partisanship in response to their assessment of the president's performance. These possibilities are particularly important in the context of the presidential partisan approval rate. In appendix B (apps. A–C are available online), we 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. As we demonstrate, when presidents are losing support, the naive partisan approval rate can overestimate the approval rate among presidential copartisans because some prior weak partisans no longer identify with the president's party. Similarly, when presidential approval is high nonpartisans' probability of identifying with the president's party increases, so the rate of nonpartisans' approval may be underestimated.

We show how the observed partisan approval rate can mislead 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 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 (excess) partisans approve (disapprove) of the president, while the lower bound assumes that the missing (excess) partisans disapprove (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 the Obama and Trump presidencies. The proportion of survey respondents who identify with the Republican party has decreased rapidly from the preelection 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.

We make a methodological contribution to the study of compositional change in public opinion polls that 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 multilevel regression with poststratification, where the poststratification is performed with information on respondents' partisanship and political characteristics, in addition to standard demographic variables. 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.

Other work in political science has employed Manski (2007) 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 US House member loses reelection when retiring members' reelection outcomes are not observed. 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 nonresponse. We contribute to this literature on 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 applicable in other settings where researchers are interested in the behavior of partisans or other groups whose composition may be endogenous.

1. Polls such as the RAND American Life Panel (Gutsche et al. 2014) employ 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. Moreover, even panel surveys are subject to attrition and nonresponse bias that may necessitate compositional correction, and in some settings, such as historical studies, panels may not exist.

BOUNDING COMPOSITIONALLY CORRECTED PARTISAN APPROVAL RATES

To derive our empirical bounds on the partisan approval rate, we model poll respondents as being drawn from a finite set of types and make one behavioral assumption about the likelihood of identifying as a party member, which we call the stable alignment assumption.[2] We assume that as one type of respondent becomes more likely to identify as a presidential copartisan then all types are weakly more likely to identify as presidential copartisans. We make no other assumptions about the likelihood of approving of the president or the relationship between identity as a copartisan and approval. In appendix A, we show that under the stable alignment assumption the compositionally corrected partisan approval rate that we define must lie within the bounds in expectation.[3]

We define our bounds under two different empirical conditions: when there is a surplus of presidential copartisans and when there is a deficit relative to a benchmark proportion. Let γ_P be the proportion of respondents that report presidential copartisanship in the benchmark.[4] Let T be the total number of respondents in the survey, and let P be the number reporting presidential copartisanship. Each respondent reports either approval or disapproval.[5] Let the number of presidential copartisans who report approval be P_A. The observed partisan approval that does not account for compositional changes is simply P_A/P.

If $P/T < \gamma_P$, then the proportion reporting presidential copartisanship is lower in the survey than in the benchmark, and we say that a deficit of presidential copartisans is equal to $T \times \gamma_P - P$ respondents. To calculate the upper bound of compositionally stable partisan presidential approval, we account for the deficit of missing presidential copartisans by assuming that each "missing" copartisan approves, so we add to both the numerator (number of approvers) and the denominator (number of presidential copartisans) arriving at

$$\frac{P_A + (T \times \gamma_P - P) \times 1}{P + (T \times \gamma_P - P)} = \frac{P_A + (T \times \gamma_P - P)}{T \times \gamma_P}.$$

To derive the lower bound, we assume that each missing presidential copartisan does not approve and only add to the denominator so that the copartisan approval rate is

$$\frac{P_A + (T \times \gamma_P - P) \times 0}{P + (T \times \gamma_P - P)} = \frac{P_A}{T \times \gamma_P}.$$

If $P/T > \gamma_P$, then the proportion reporting presidential copartisanship is higher in the survey than in the benchmark, and we say that we have a surplus of presidential copartisans equal to $P - T \times \gamma_P$ respondents. To calculate the upper bound of compositionally stable partisan presidential approval, we account for the surplus of presidential copartisans by assuming that each "extra" copartisan disapproves, and so while we adjust the denominator down, we do not adjust the numerator down, arriving at

$$\frac{P_A}{P - (P - T \times \gamma_P)} = \frac{P_A}{T \times \gamma_P}.$$

To derive the lower bound, we assume that each extra copartisan does approve, and we adjust both the numerator and the denominator down, so that the copartisan approval rate is[6]

$$\frac{P_A - (P - T \times \gamma_P) \times 1}{P - (P - T \times \gamma_P)} = \frac{P_A - (P - T \times \gamma_P)}{T \times \gamma_P}.$$

We have defined our bounds in the simplest possible context, but the procedure can be extended to accommodate more complex situations such as survey weights and question nonresponse. Additional theoretical restrictions can also be applied to tighten the bounds.

EVOLUTION OF THE BOUNDS DURING THE 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 Gallup Analytics. Gallup Analytics reports cross-tabs for presidential approval by partisanship and the number of respondents identifying as Democrats, Republicans, and Independents every week.[7] We use all available presidential approval data from 2009 to 2017 for the analysis.[8]

2. The finite set of types characterizes the probability that a member identifies with the president's party and approves of the president. The types could represent demographic or attitudinal groups. The exact definition of types is not essential for deriving the bounds.

3. In the appendix section "Relaxing the Monotonicity Condition," we discuss how potential violations of the stable alignment assumption would make the bounds less informative.

4. In our application, we use the proportion reporting copartisanship with the presidential winner in the last available preelection poll, but the analyst could employ any proportion.

5. In our base case, we remove respondents who do not answer the approval question.

6. We further account for the fact that there may not be a sufficient number of missing or excess copartisan of a specific type in a poll to add or remove. In practice these modifications rarely hold.

7. 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 re-create partisanship proportions that include leaners.

8. Gallup Analytics reports the presidential approval rate using their US daily tracking poll. The lowest level of temporal aggregation that includes

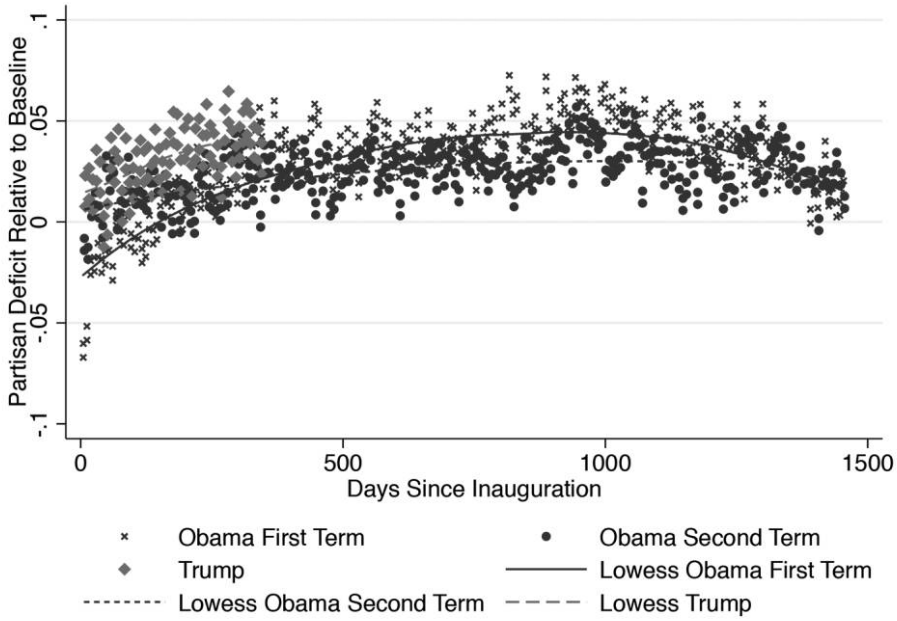

Figure 1. Gallup Analytics partisan proportions

We calculate the partisanship deficit or surplus relative to the partisanship proportions in the last poll available before the previous election. We believe that the last preelection poll is the best measure of the electorally relevant coalition size, but our approach is flexible enough to employ alternative benchmarks in different contexts.[9] We use unweighted partisanship proportions because Gallup Analytics reports the count 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 by computing the standard deviation of the upper and lower bounds across 200 bootstrap replications.[10]

In figure 1, we plot the partisan deficit relative to the preelection baseline partisanship proportions against the number of days since the inauguration for the two terms of the Obama presidency and the first seven months of the Trump presidency.[11] We also include separate LOWESS (locally weighted smoothing scatterplot) plots for each presidential term. The most striking finding in the plot is that the deficit of presidential 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 smoothed partisan deficit 346 days into the Trump presidency (approximately 0.0484) is never reached in Obama's first or second terms (which achieved a maximum of 0.048–902 days into Obama's first 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 Trump's partisan approval rate.

In figure 2, we plot observed partisan approval rates, the bounds on the compositionally corrected partisan approval rate, and 95% confidence intervals for the upper and lower bound during the first 346 days of each presidential term.[12] The marker is the observed partisan approval rate, the dark

the party affiliation of respondents is the week. The data start with January 19–26, 2009, and the final week available when this analysis was conducted was December 25–31, 2017.

9. As a robustness check, we replicate our analysis using average partisanship from a broader set of polls in figs. C.2 and C.3 in app. C.

10. The lack of individual-level data requires us to construct pseudo-individual 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 data for the independents consists of 1,000 observations with 450 observations set equal to 1 for approval and 550 observations with zeroes for the approval variable.

11. Consistent with our findings in app. B, there is a negative relationship between presidential approval and the magnitude of the partisan deficit in the Gallup Analytics sample. The correlation is −0.5665, and in a regression with robust standard errors of the president's partisan deficit on presidential approval the coefficient estimate is −0.176 and the *t*-statistic is a highly significant −15.28. The results are similar if we use presidential approval from the previous week as the independent variable.

12. The final poll from our sample is 346 days after Trump's inauguration, so we restrict attention to this time span in order to ease visual comparison. In fig. C.1 of app. C, we show the entire history of the observed partisan approval rates, bounds, and confidence intervals over the complete 2009–17 period.

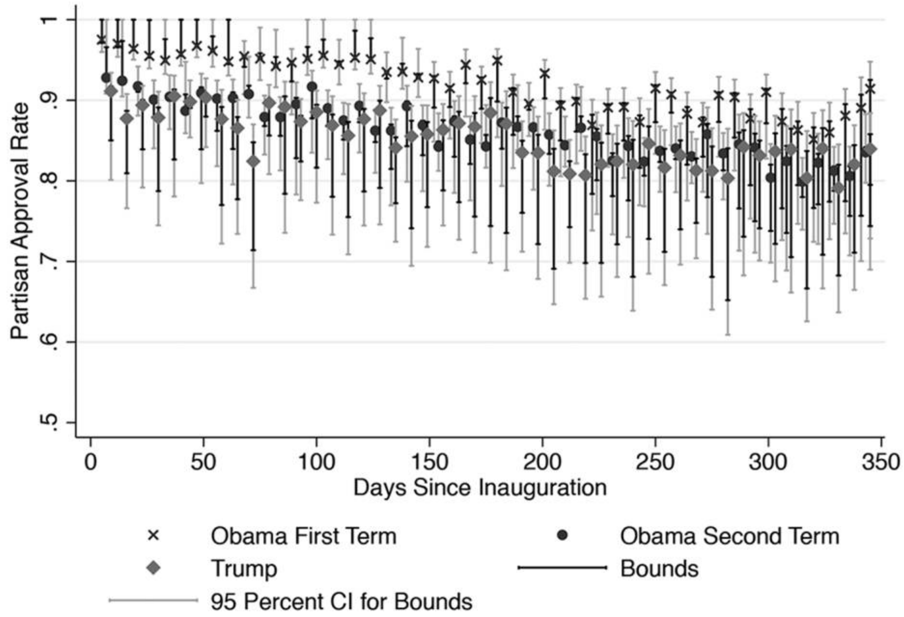

Figure 2. Gallup Analytics partisan approval rates

capped line extending out from the marker are the bounds, and the lighter capped lines are the 95% confidence intervals on the lower and upper bounds.[13]

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 40 of the 49 weeks, the lower bound is below 0.8. In all instances, the lower bound on Trump's compositionally corrected partisan approval rate is lower than the lower bound from the analogous poll during Obama's first term.[14] The observed partisan approval rate is partially an artifact of missing respondents who would have previously reported Republican partisanship. While 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.

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 seven 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.

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

14. Moreover, the mean lower bound on the partisan approval rate during the first 49 weeks of Obama's first term is approximately 0.886 compared to 0.755 during the Trump presidency. Even in Obama's second term, the mean lower bound during the first 49 weeks is 0.832.

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