Alphabet Soup: Randomized Ballot Order and the Representation of Marginalized Candidates

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Abstract

Ballot order has the power to shape electoral outcomes: even when randomly assigned, candidates listed first receive more votes. Meanwhile, candidates from historically marginalized groups—such as women and ethnic minorities—tend to fare worse in low-salience or low-information contexts. We examine whether being listed first—an increase in salience—benefits these candidates more or less than their white and male counterparts. Leveraging a natural experiment from over 29,000 California local elections between 1995 and 2021, in which ballot order was randomly assigned, we estimate the causal effect of being listed first on vote share by candidate gender and ethnicity. We find that while all candidates receive a premium from appearing first, non-white candidates benefit significantly more than white candidates. This advantage varies with the partisan and racial composition of the electorate and with election timing. We close by discussing the implications of election design for equality of representation.

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Introduction

Prior work finds that candidates benefit from being listed first on the ballot. This positional advantage—often termed "donkey voting"—appears across a wide range of electoral contexts, especially in local, nonpartisan, and complex elections (Australia: King and Leigh 2009; Orr 2002; Britain: Webber et al. 2014; Rallings and Thrasher 1997; Colombia: Gulzar, Robinson, and Ruiz 2022; Korea: Jun and Min 2017; Spain: Lijphart and Pintor 1988; U.S.: Meredith and Salant 2013; Koppell and Steen 2004; Miller and Krosnick 1998). This advantage is usually attributed to cognitive bias: the more candidates to be examined, the higher the informational costs to the voter (Ho and Imai 2008; see also Meredith and Salant 2013; Miller and Krosnick 1998; Bain and Hecock 1957).[1]

A large and separate literature documents how cognitive biases and stereotypes shape the evaluations and electoral outcomes of women and ethnic minorities (e.g., Bauer, 2020*a*,*b*; Berinsky et al., 2020; Crowder-Meyer et al., 2020; Fulton and Dhima, 2020; Doherty, Dowling, and Miller, 2019). Candidates from historically marginalized groups—here, women and ethnic minorities—often fare worse in low-salience and low-information contexts such as on-cycle local elections (Anzia and Bernhard, 2022; de Benedictis-Kessner, 2018). These patterns are commonly attributed to increased reliance on stereotypes and heuristics when voters know little about local candidates (Bernhard and Freeder, 2020; McDermott, 1997, 1998; Kreiss, Lawrence, and McGregor, 2020; Holman, Merolla, and Zechmeister, 2017). In such settings, social categories like race and gender that are easily inferred from names can become shortcuts for inferring less visible attributes like ideology or partisanship (Badas and Stauffer, 2023; Crowder-Meyer et al., 2015, 2020).

Taken together, we believe these literatures produce a natural question: does increased salience from being listed first benefit marginalized candidates more or less than majority-group candidates? On the one hand, being listed first may help overlooked candidates by

[1] This mechanism differs from biases in ballots that order candidates by seniority, in which first place signals rank; or systems that order candidates by past performance (e.g., in a primary, or by past party performance), in which first place signals strength rather than visibility.

drawing voters' attention to them. On the other, the same increase in visibility could also amplify stereotyping, particularly when additional information is absent. Randomized ballot order therefore offers a rare opportunity to disentangle the effects of low salience from those of low information in shaping voter behavior toward marginalized candidates.

We explore these hypotheses about salience, information environments, and the representation of marginalized groups by drawing on a natural experiment contained within California elections. Using data on more than 29,000 local elections from 1995-2021 in which ballot order is randomized anew for each cycle, we assess how ballot order affects candidates' electoral success. We code whether the candidates in these races are (1) women and (2) ethnic minorities to test how such candidates' share of the vote is influenced by randomly being listed first on the ballot relative to male and white candidates, respectively.

In keeping with previous scholarship, we find that all candidates benefit significantly from being listed first, but we also find evidence of substantial and previously unrealized heterogeneity of these effects. These "salience boosts" are largest for non-white candidates and, most of all, for non-white women. The probability that a non-white woman will win a seat increases by almost nine percentage points if she is listed first, while the same effect is just under 2.5 percentage points for white men. Consistent with our theory, we also find these differences are widest in information environments where voters are more likely to satisfice (races with many candidates) or know little about local candidates (races held concurrent with national elections).

To deepen our understanding of the mechanisms and scope conditions that produce these effects, we then explore three sets of potential moderators. First, we look at variation by incumbency (de Benedictis-Kessner, 2018; Trounstine, 2013). Does the familiarity of being an incumbent candidate "drown out" any salience boost from ballot position? Second, we explore whether these effects might be driven by negative stereotypes about or inferred partisanship of non-white and women candidates (Anzia and Bernhard, 2022; Bernhard, 2020). Are these effects smallest in heavily Democratic constituencies, where the effect of salience might pale in comparison to ideological choices due to stereotypes that women and ethnic minorities are more liberal (Cargile, 2023; Jones, 2014)? And can being listed first provide a cue that compensates for the negative effects of stereotypes in more conservative districts? Third, we investigate whether ethnic minority population size conditions these salience effects (Marschall, Ruhil, and Shah, 2010; Trounstine and Valdini, 2008). Do non-white candidates benefit most from salience effects in majority-minority districts where they have a natural advantage, or among white-majority districts where they are more marginalized?

Our exploratory analyses of these moderators demonstrate just how heterogeneous the effects of a simple salience cue can be. We find some evidence that incumbency drowns out the differences in the size of the salience cue between candidates of different racial backgrounds. We further find that the salience boosts to non-white candidates are largest in Republican districts, which may mean that being listed first helps these candidates overcome voter stereotypes about their liberalism, or simply that there is a ceiling effect to these candidates' advantages in Democratic districts. Then, we show that the salience boosts are greatest for non-white candidates running in more heavily Hispanic jurisdictions; whether this is evidence of linked fate or affinity effects (McConnaughy, 2010) is something that we hope future research, especially survey and experimental work, will explore.

In short, we find considerable and previously unobserved variation in just how salient candidates listed first can be. Given that these boosts make a large difference in the probability that candidates win their races, this has a sizable impact on the representation of women and ethnic minorities at the local level. In turn, who is elected at the local level shapes many other outcomes, from the local policy outcomes it produces (Holman, 2014, 2017; Shi and Singleton, 2023) to the likelihood that these candidates will stay in politics (Bernhard and de Benedictis-Kessner, 2021) or go on to run for higher office (Einstein et al., 2020; Sweet-Cushman, 2018). As we lay out in the discussion, these differential returns to randomly being listed first may have major implications for unequal representation.

Theory

Our argument rests on a simple premise: ballot position affects how much attention voters pay to a candidate. When information is limited and choices are numerous, many voters satisfice—selecting the first option that seems acceptable rather than evaluating all alternatives (Bernhard and Freeder, 2020; Lijphart and Pintor, 1988; Burnett and Kogan, 2015). Being listed first therefore supplies a small but meaningful boost in salience: the candidate becomes more visible and more cognitively accessible at the moment of choice (Ho and Imai, 2008; Miller and Krosnick, 1998; Bain and Hecock, 1957).

Where we depart from existing literature is in theorizing that there are uneven returns to being listed first. We argue this occurs in two main ways. First, returns to salience should be larger for candidates who are otherwise less likely to attract voter attention: those who are less familiar (McDermott, 1997), have fewer resources (Johnson and Selin, 2012; Sorensen and Chen, 2022; Grumbach and Sahn, 2020; Grumbach and Staszak, 2022), or are perceived as atypical for the office (Bernhard and Holman, 2025; Sweet-Cushman, 2021; Huddy and Terkildsen, 1993). In local elections, these are often women and ethnic minority candidates. When voters are inattentive or overwhelmed, such candidates may be overlooked; random salience from first position can help level that playing field. This logic yields our baseline expectation that being listed first will increase vote share for all candidates, but that the gains will be larger for women and non-white candidates than for white men.

Second, we expect the magnitude of these effects depends on the broader information environment. When voters face heavy cognitive loads due to long ballots, crowded races, or low campaign visibility, they should rely more on these position and social cues (McDermott, 1998).[2] One feature of the information environment around elections that has been examined in previous research is the timing of those elections. In California and elsewhere

[2]We also thank a reader for suggesting that voting by mail (VBM) might give voters a less time-pressured and more information-rich environment and thus reduce these effects. While data on VBM rates was not available for most of our offcycle election races, we hope future research will explore this possibility in more detail.

in the U.S., many local elections happen either in the spring or in November of odd years (off-cycle elections) rather than in November of even years (on-cycle with presidential and midterm elections). Past research has documented both how the origins of election timing decisions were shaped by interest groups (Anzia, 2012*b*) and how this crucial institution has shaped policy outcomes (Anzia, 2012*a*, 2014, 2011). Yet the timing of elections can also influence representation and accountability (Anzia and Bernhard, 2022; de Benedictis-Kessner, 2018; Dynes, Hartney, and Hayes, 2021; Payson, 2017, cf. de Benedictis-Kessner and Warshaw, 2023), potentially via altering the composition of the electorate (Hajnal, Kogan, and Markarian, 2022; Kogan, Lavertu, and Peskowitz, 2018).

In particular, off-cycle elections may make it more likely that those people who do show up to vote are more knowledgeable (Berry and Gersen, 2011; Hajnal and Lewis, 2003; Oliver and Ha, 2007). Off-cycle elections may also alter the choice task of actually casting a ballot: when local elections are held at a different time from national elections, and thus the number of races on the ballot is usually fewer, voters may be able to devote more time and effort to decision-making rather than relying on heuristics such as incumbency, demographics, or ballot position, which they may do if they experience choice fatigue (Augenblick and Nicholson, 2016; de Benedictis-Kessner, 2018; Iyengar and Lepper, 2000; Selb, 2008). Election timing may be a particularly influential institutional feature of elections given that it shifts both the overall turnout of voters *and* the types of voters who show up to vote. This might lead to potentially larger effects on outcomes than other administrative procedures governing elections (Grimmer and Hersh, 2023).

Following this literature, we expect that the effects of being listed first on the ballot should be largest in local elections where many voters do not know much about the candidates and are most likely to satisfice—as may be the case in on-cycle presidential and midterm elections (Anzia and Bernhard, 2021; de Benedictis-Kessner, 2018)—and smallest when voters know more about local candidates—as may be the case in off-cycle elections.

These are our main theoretical expectations, and straightforward to test with the right kind of data (which we argue in the next section a natural experiment contained in California local elections offers us). But we also want to better understand the mechanisms and scope conditions underlying these effects. As an exploratory set of analyses, we propose three sets of potential moderators.

We first examine whether incumbency status moderates the first-position effect. We ask whether incumbency crowds out the salience boost or whether ballot primacy reinforces incumbents' existing advantages. On the one hand, incumbent candidates are already more salient due to name recognition, larger networks, and greater resources (de Benedictis-Kessner, 2018; Trounstine, 2013). As such, first-place salience may have limited additional effects. Conversely, if salience operates as a general attention multiplier, it could strengthen incumbency advantages. We therefore model the effects among non-incumbents and among incumbents: the first-position boost may either be attenuated (crowding out) or amplified (reinforcement) for incumbent candidates.

Next, we test whether partisan context shapes the differences in our effects among marginalized candidates who we expect to benefit especially from an early ballot position. There are two primary reasons this might happen. Candidates' increased salience from being first on the ballot may offset negative stereotypes that voters hold about non-white candidates and women. On the one hand, there is a close relationship between racial attitudes and partisanship in the US (e.g. Engelhardt, 2021; Westwood and Peterson, 2022) as well as between gender attitudes and partisanship (Huddy and Lizotte, 2008; Deckman and McTague, 2015; Barnes and Cassese, 2017). This correlation with voters' partisan leanings may therefore condition the way that ballot order heuristics interact with the race and gender of candidates. Specifically, Republican voters might be less likely to vote for non-white candidates or women than white candidates or men, but more likely to vote for these candidates when they are listed first because of the strength of the ballot order heuristic. This would lead to a larger effect of ballot position for these types of candidates among Republican voters due to a lower baseline level of support.

Alternatively, voters of all partisan stripes may be more likely to engage in ideological stereotyping or ideological projection. Voters tend to stereotype women (Anzia and Bernhard, 2022; Bauer, 2018*b*,*a*; Sanbonmatsu, 2002) and ethnic minorities as more ideologically liberal (Berinsky et al., 2020; Cargile, 2023; Jones, 2014; Juenke and Shah, 2016; Lerman and Sadin, 2016; McDermott, 1997, 1998; Meier et al., 2005; Sigelman et al., 1995)—often argued to be a form of statistical discrimination, as most women and minority candidates run as Democrats (Juenke and Shah, 2016). Under these theories, voters might be more likely to assume that non-white and women candidates are more liberal. Democratic voters may take this as a signal of ideological similarity, and choose non-white women regardless of ballot position. Voters in more Democratic areas may already have a high baseline support for such candidates, reducing the size of the effects that candidates gain from being at the top of the ballot.

Finally, we consider whether local ethnic composition conditions the impact of salience, assessing whether non-white candidates gain more from being listed first in majority-minority districts—where co-ethnic affinity may reinforce support (Krupnikov and Piston, 2015; Marschall, Ruhil, and Shah, 2010; McConnaughy, 2010; Trounstine and Valdini, 2008; Ansolabehere and Fraga, 2016)—or in predominantly white districts, where they are more often overlooked or even resisted (McConnaughy, 2010). Indeed, Moy and Vitale (2025) show that affinity effects exist even within race: white candidates of various ethnicities (e.g., Polish Americans) receive more votes in neighborhoods where their co-ethnics are clustered than from other white neighborhoods. Importantly, however, (Matsubayashi and Ueda, 2011) find that only informed white voters change their votes away from black candidates; uninformed white voters do not, perhaps because black names are less racially distinctive than Hispanic and Asian American names (Butler and Homola, 2017). In heavily Hispanic (the largest ethnic group in California) or otherwise diverse districts (Moy and Vitale, 2025; Shah, Marschall, and Ruhil, 2013), increased visibility may strengthen co-ethnic or cross-minority affinity voting, as the salience boost from being listed first makes shared identity more cognitively accessible. However, as with Democratic partisanship, these environments may already provide high baseline support for minority candidates, again producing ceiling effects. Observing how ballot position effects shift with demographic composition allows us to probe whether salience magnifies or diminishes such affinity dynamics.

In short, we expect ballot order matters, and might matter especially to 1) historically marginalized candidates (who might be discounted otherwise) and 2) when voters are inattentive or overwhelmed, as with long ballots or during on-cycle elections. We are more agnostic about our moderators—incumbency, partisanship, and district composition—but given a perfect setting to explore the limits of salience cues, we find them worth exploring descriptively. Accordingly, in the next section, we explain how we test these expectations using a natural experiment run over several decades of California local elections, where ballot order is randomized in every contest.

Data and Methods

To better understand how voters' cognitive biases shape election outcomes, we examine a natural experiment occurring in California local elections to assess whether there are real-world heterogeneous treatment effects of ballot position in a context that randomly varies ballot order. Our election data come from the California Elections Data Archive (CEDA), which encompasses county, city, community college, school district, and special district elections in California from 1995 to 2021 (CEDA, 2022; de Benedictis-Kessner and Bernhard, 2022). In total, the CEDA data contain information on candidates in 29,076 elections for multiple levels of local government office ranging from directors of small local community services districts that provide water treatment and fire protection, to mayors and city councilors in large cities that supervise multi-billion dollar budgets.

Importantly, local elections in California are nonpartisan. This means that California is a case in which voter reliance on heuristics and stereotypes related to race and gender may be especially visible. On the other hand, it is also an immensely diverse state, with approximately 39 million residents and no dominant ethnic group. According to the 2022 American Community Survey, 40% of Californians are Latino or Hispanic-identifying, 35% are white, 15% are Asian American or Pacific Islander (AAPI), 5% are Black, 4% are multiracial, and roughly 1% are Native American or Alaska Natives (Johnson, McGhee, and Mejia, 2024). Such diversity may mean that there is no singular heuristic or stereotype about candidates being applied, and thus may be difficult to observe.

To assess the effect of ballot order using these elections, we leverage the fact that the ballot order of candidates in California is decided by a random drawing for each unique election date. In many contexts, the jurisdiction lists candidates in alphabetical name order, producing an alphabetic bias. In 1975, California implemented random alphabets for each election to circumvent alphabetic bias and ensure that candidates whose surnames are closer to 'A' do not benefit election after election (Miller, 2010; Scott, 1972). Specifically, prior to every election, the California Secretary of State randomly draws all letters in the alphabet and records the order in which they appear. This randomized alphabet then serves as the "alphabetical order" in which candidates are listed on every ballot.[3] This means that it is entirely random in each election whether or not a candidate will appear first—or in any given position—on the ballot.

California is an ideal case for studying ballot order randomization as it is one of only thirteen American states that fully randomizes its ballot order[4], and the remaining twelve states are far less ideal, for two reasons. First, many of these states (Arkansas, Iowa, Kentucky, New Hampshire, Oregon, Rhode Island, South Dakota, Utah, West Virginia) have much lower levels of ethnic diversity, producing fewer elections in which people of color run, and thus reducing the number of useful observations. Second, California has randomized ballot order for fifty years, far longer than any other state, with the majority of these states having adopted ballot order randomization only in 2019 or later (Iowa, Kentucky, New Hampshire,

[3]For more details, see: `https://www.sos.ca.gov/elections/randomized-alphabet`.
[4]For a list of states, see: `https://ballotpedia.org/Rules_on_ballot_order_and_party_labeling`.

New Mexico, New Jersey, Oregon, Virginia, West Virginia). California is therefore the only state with sufficiently high ethnic variation and sufficient elections over time to produce a sufficient number of observations for analysis.

We use this exogenous variation in ballot position to identify candidates that appeared first within a race on the ballot and assess the effect on voters' choices by regressing each candidate's vote share (i.e. their share of the total votes cast in a race) on this indicator of ballot position. To account for additional variation induced by state-wide shifts in voting patterns or competitiveness, we include fixed effects for year, county, and office. This helps to adjust for additional confounding that may affect candidates' electoral success that happens to coincide with the ballot-order alphabet picked for that election.

We code nearly every candidate algorithmically for gender (dichotomized as "woman" and "man") using the `gender` package in `R`. Similarly, we code race and ethnicity of first names using mortgage data and of surnames using the `wru` package in `R` to code first names and surnames, creating a joint probability of being in a given ethnic group in a given district in California (details in Appendix A). Only a few (e.g., those missing a first name, for gender, or those missing a surname, for race) are uncoded.

Coding our moderator variables—incumbency, district partisanship, and district racial composition—is simpler. The CEDA data already contains information on incumbency, which we check against de Benedictis-Kessner and Bernhard (2022)'s more updated database. Because some of our races are municipal (e.g., mayor), some districted in some places and some at-large in others (e.g., city council, school board), and some county (e.g., sheriff), we code both partisanship and racial composition at the county level. This choice obscures some greater variation (there are much more Republican districts than counties, for instance), which in turn should bias us against finding effects of these variables. For partisanship, we use data on presidential vote returns (Voting and Election Science Team, 2020) aggregated at the county level (Tausanovitch and Warshaw, 2013), which we then match to the county of each election. For racial and ethnic composition, we use data from the Census Bureau's decennial census for each county in which an election in our data occurred, and calculate the percentage of the county's population coming from various racial/ethnic groups delineated by the Census. Specifically, we use data on the proportion of a county that identifies as white when listing a single race alone. We also use data on the proportion of a county's population with Hispanic/Latino origin (regardless of race).

Results

Election data enable us to assess the real-world effects of ballot position on candidates' electoral success. Here, we analyze only races where two or more candidates competed. Overall, we find that candidates who appeared first on the ballot earned 33% of the vote on average, while those who appeared later on the ballot earned an average of 22.9% of the vote.

This difference in vote shares between first and later candidates masks substantial heterogeneity in candidate success based on other factors, such as the overall competitiveness of the race. For instance, the average candidate in competitive races with a greater total number of candidates mechanically earns less of the vote share than she would if she had earned the same number of votes against fewer candidates. Furthermore, ballot ordering effects may be stronger in certain years, certain jurisdictions, for certain types of races, and between certain types of candidates, given various imbalances in candidate salience, such as relative differences in campaign spending by candidates, absolute differences in election interest in different years and races, familiarity gaps between incumbents and challengers, and so on.

To address these sorts of factors, we examine the effect of being listed first on the ballot more formally in a regression format and display the results in Figure 1. We do this for our main outcome (candidates' vote share in the election) and for a secondary outcome (candidates' probability of winning), and display results for each outcome in the two panels

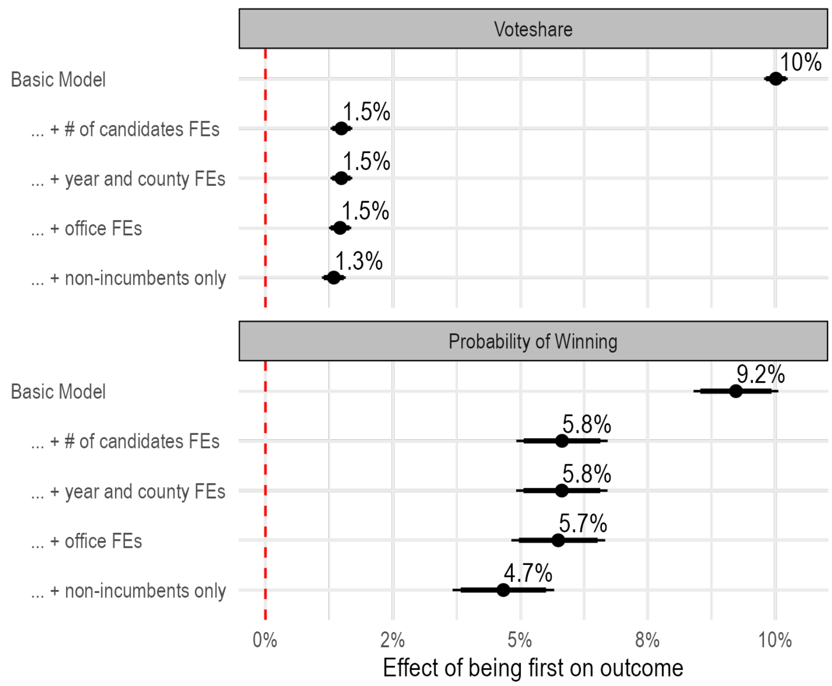

Figure 1: Effect of First Ballot Position. Each model includes all control variables from the preceding model, and standard errors in all models are clustered at the unique contest-level. Thicker bars depict 90% confidence intervals and thinner bars, 95% confidence intervals.

of Figure 1. We control for a number of features of the election and candidates in order to improve our statistical efficiency and the precisions of our estimates, and also to account for a number of features of the elections that might drive differences in candidates' electoral success that are unrelated to ballot position. We regress vote share on "first position" (Model 1) and then progressively adjust our model specification by adding fixed effects for the number of candidates in the race (2), adding election year fixed effects and county fixed effects (3), adding fixed effects for the type of office sought (4), and restricting to non-incumbent candidates only (5).

As displayed in Figure 1, our most basic specification yields an estimate of a 10 percentage point increase in vote share from being listed first on the ballot, but after accounting for the number of candidates in a given race, this effect is smaller in size.[5] Once controlling for the number of candidates with fixed effects, the effect of being listed first on vote share is

[5]We show how the size of our estimates of the effect of ballot position varies as a function of the number of candidates in the race in Appendix B.

relatively consistent in size across specifications in the top panel of Figure 1. Being listed first on the ballot yields a reward of 1.3–1.5 percentage points in vote share. Our results using the alternative outcome of win probability similarly are statistically and substantively significant, as shown in the bottom panel. Candidates who appear first on the ballot are between 4.7 and 5.8 percentage points more likely to win than candidates who do not appear first.[6] For ease of reference, we stick to analyzing vote share in subsequent figures, but the results for wins by race and gender are given in Appendix E.

Race and Gender

We next examine the effects of ballot position on candidates of different races and genders. We continue to formally assess the effect of random ballot position using a regression format, but subset our analyses by candidates' gender and race. In each regression, we use fixed effects to control for the total number of candidates in the race, year, county, and office, and restrict to non-incumbents (equivalent to the fifth model in Figure 1 above). We display these results in Figure 2, which plots the effect of being first on vote share among these subsets of non-incumbent candidates.

We see clear evidence of a larger advantage for non-white candidates who are listed first. Non-white candidates receive a larger boost to their vote share (1.9 percentage points) from being randomly listed first on the ballot than white candidates (1.2 percentage points). The difference in the size of these effects is significant ($p = 0.008$). This boost is slightly larger for non-white women candidates (2.2 percentage points) than for non-white men candidates (1.8 percentage points), but this difference is not statistically significant. Women generally receive a slightly larger boost (1.5 percentage points) relative to men (1.3 percentage points), but this difference is also not statistically significant. Overall, a candidate of color can expect

[6] We conduct a number of robustness checks of our main analyses. These include analyzing heterogeneity in the effects of ballot position among different types of office in Appendix C. We find that these effects are largest in races for lower-salience offices such as treasurer, clerk, school board, and county legislator, while smaller (and insignificant) in higher-salience races for offices such as mayor. We also analyze heterogeneity in these effects in the earlier period of our data and the later period and present these results in Appendix D, which show little difference in the size of the effects before vs. after 2012.

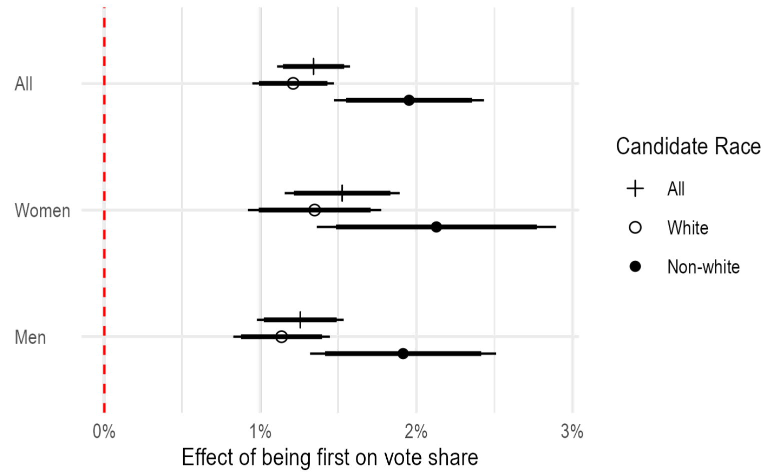

Figure 2: Effect of Ballot Order by Race and Gender, among Non-incumbents. Models include fixed effects for the number of candidates, year, county, and office. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

to receive a larger premium for being listed first relative to a white candidate.[7]

Election Timing

Our results indicate that ballot position has a substantial effect on voters' choices and therefore candidates' success. This influence benefits non-white candidates who appear first relative to white candidates who appear first, but does not appear to produce differential effects by gender. However, other features of the campaign and election context that differ across races may also influence candidates' success and the influence of ballot order. Given that these ballot position effects may be driven by voter inattention or being overwhelmed, it is important to explore variation in contextual factors that might influence voters' levels of knowledge.

In this section, we examine whether the timing of elections—either concurrent with presidential elections, concurrent with midterm federal elections, or off-cycle—moderates the effect of ballot position on candidates' success. To do so, we subset our analyses by the

[7]Using our alternative outcome of candidates' win probability shows similar heterogeneity: ballot position influences candidates' electoral success to a greater degree for non-white candidates, as we show in Appendix E.

timing of elections and display the results in Figure 3. Our theory predicts that ballot order effects will be larger in on-cycle elections that mobilize greater numbers of low-information voters and which feature more distractions from local races in favor of national politics.

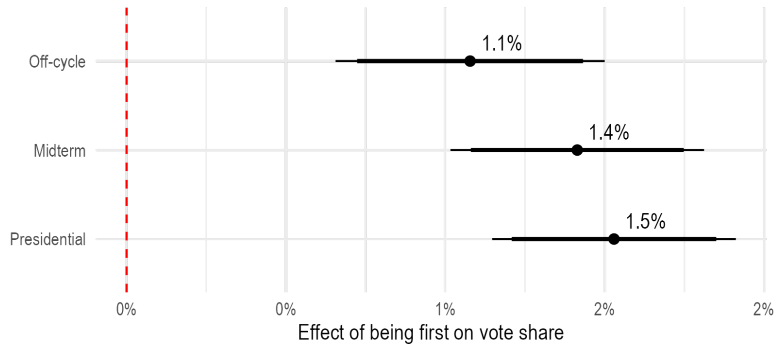

Figure 3: Effect of Ballot Order by Election Timing. Models use standard errors clustered by contest, include fixed effects for year, county, and office, and are limited to non-incumbent candidates only. Thicker bars depict 90% confidence intervals and thinner bars, 95% confidence intervals.

In line with these predictions, we find that the ballot order effect is somewhat moderated by election timing. On the top line, we plot the effect of being first on the ballot in races held off-cycle, where we see an effect of approximately 1.1 percentage points. In races concurrent with midterm (middle line) and presidential (bottom line) elections, however, we observe larger effects of ballot position of 1.4 and 1.5 percentage points, respectively. The difference in effect size between off-cycle and presidential races is not statistically significant ($p = 0.12$). Yet these differences are directionally consistent with our theoretical predictions for the mechanism behind the ballot order effects we observe.

We next assess how this moderation of ballot order effects extends across different types of candidates. Figure 4 plots the effects of appearing first on the ballot across different election timing contexts, divided up between candidates of different racial/ethnic backgrounds and genders. Each panel shows these differences within groups of elections held at different times, with presidential on-cycle elections in the left panel, midterm on-cycle elections in the middle panel, and off-cycle elections in the right panel.

In on-cycle elections (the left and middle panels), we find large differences in the size of the effect of ballot position between white and non-white candidates, as demonstrated by the difference between the size of the effects plotted with open circles (white candidates) and filled circles (non-white candidates) in the top set of points in both panels. The substantive size of this racial heterogeneity in the effects of ballot position in on-cycle elections bears some attention. In on-cycle (both presidential and midterm) elections, non-white candidates receive a boost to their vote share from being listed first on the ballot that is 70%–85% larger than the boost that white candidates receive.[8]

In presidential elections, this difference appears to be concentrated among women candidates (plotted in the middle set of points in the left panel), while in midterm elections, it appears to be concentrated among men candidates (plotted in the bottom set of points in the middle panel). In presidential elections, non-white women (filled circle in the middle set of points) receive a 3.4 percentage point boost from being listed first on the ballot, compared to just 0.6 percentage points for white women (open circle in the middle set of points).[9] Meanwhile, in midterm elections, non-white men receive a 2 percentage point boost from being listed first, while white men only receive a 0.8 percentage point boost.[10]

In off-cycle elections, however, the effects of being listed first on the ballot are smaller overall (as shown in Figure 3), but also display little heterogeneity in their size between men and women candidates or between white and non-white candidates. None of the differences displayed in the right panel of Figure 4 are meaningfully different in size nor statistically distinguishable from one another.

Our evidence indicates that election timing moderates the effect of being listed first, with reliance on this heuristic higher in contexts like on-cycle elections which have higher numbers of low-information national voters, as well as more distracting information available to voters due to the greater number of races on the ballot. We also find that the use of this

[8]In both cases, the differences between non-white and white candidates are also statistically significant. The 0.9 percentage point difference in presidential elections is statistically significant ($p = 0.009$), as is the 1 percentage point difference in midterm elections ($p = 0.01$).

[9]This 2.8 percentage point difference is statistically significant ($p = 0.028$).

[10]This 1.2 percentage point difference is also statistically significant ($p = 0.012$).

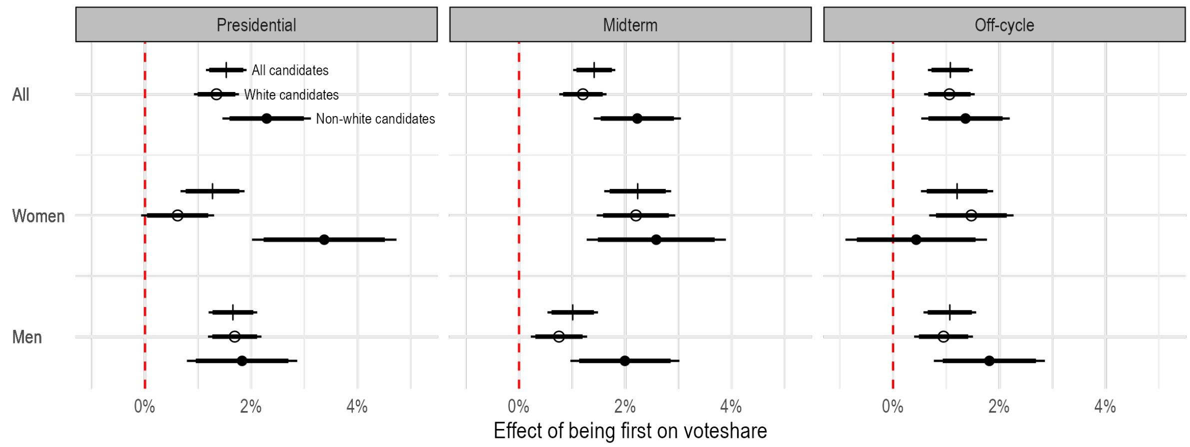

Figure 4: Effect of Ballot Order by Race, Gender, and Election Timing. Models include fixed effects for number of candidates, year, county, and office, and are limited to non-incumbent candidates only

heuristic impacts non-white and white candidates differently. Non-white candidates benefit most in on-cycle presidential and midterm elections relative to white candidates. Together, this evidence is consistent with a theoretical mechanism of voter inattention and heuristic use driving ballot position effects more generally, which is exacerbated in contexts with longer ballots, greater numbers of low-information voters, and more distracting information about national politics.

Moderators

Observational data at the election level alone does not allow us to draw firm conclusions about the ultimate causes of variation in the size of effects that we observe. Yet there are a number of theoretically relevant mechanisms that we can investigate using variation in several measureable features of the context in which these elections occur. Using data on several of these contextual factors, we next explore potential moderators of the size of our effects.

Incumbency

Heterogeneity in the effects of ballot order by race and gender is clear among non-incumbent candidates, but incumbency may swamp the effects of ballot order due to the increased salience from which incumbent candidates benefit among voters. We replicate our analyses for incumbent candidates only, and present these results below in Figure 5. Consistent with our theory, these analyses display little heterogeneity in the effects of ballot order by race or gender among incumbents.

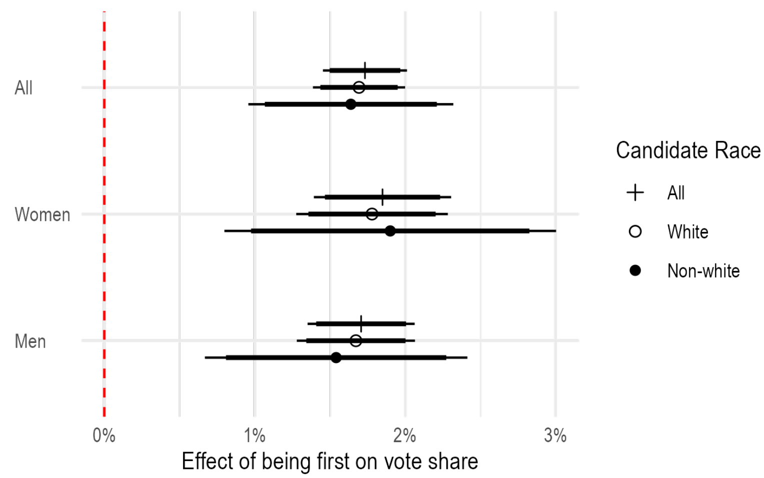

Figure 5: Effect of Ballot Order by Race and Gender, among Incumbents. Models include fixed effects for number of candidates, year, county, and office. Only incumbent candidates included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

Taken together, simply increasing the salience of candidates by (randomly) listing them first on the ballot does not appear to generate a significantly bigger reward—or penalty—for women relative to men. However, increasing the salience of non-white candidates by listing them first delivers a significantly larger reward than white candidates receive. One strong potential explanation is that non-white candidates consistently receive less media attention, or are able to run fewer campaign ads, and therefore benefit more from a salience bump relative to white candidates who are already more well-known. Indeed, candidates of color receive fewer campaign funds relative to white candidates (Johnson and Selin, 2012; Sorensen and Chen, 2022), which in turn is explained by a combination of co-ethnic donation patterns and the under-representation of people of color in the donor class (Grumbach and Sahn, 2020; Grumbach and Staszak, 2022). A third possibility is that any advantages accrued from being white result in a larger proportion of incumbents who are white, and thus challengers are more likely to be non-white and therefore are more likely to benefit from the salience boost; this would be consistent with our finding that the non-white premium vanishes when comparing non-incumbents to incumbents (Figure 5).

Voter Partisanship

We have shown heterogeneity in the effects of ballot order by candidates' race, as well as by the timing of those elections. Yet our analyses still leave many questions unanswered about how *voters'* characteristics may moderate the effects of ballot order. To investigate this, we next assess whether the partisanship of the electorate conditions the heterogeneity of the effects that we observe by candidates' race and gender.

As a first cut at this analysis, we present bar plots showing the average vote share received by white and non-white candidates in either first or second position in two-candidate races only, for the purposes of simplicity. The two panels of Figure 6 display these numbers in less Democratic counties (left panel) and more Democratic counties (right panel). In less Democratic counties, non-white candidates who are not listed first pay an electoral penalty. They receive an approximately three percentage point lower vote share on average relative to white candidates who are not listed first on the ballot in less Democratic counties. Yet non-white candidates who are listed first on the ballot in these counties appear to garner vote shares on par with white candidates.

We assess this more rigorously by replicating our previous analyses in a regression format using subsets of counties based on roughly quartiles of county-level Democratic presidential vote share. These results are shown in Figure 7, where we present subgroup analyses for CA counties (from least Democratic-leaning in the top-left to most Democratic-leaning in

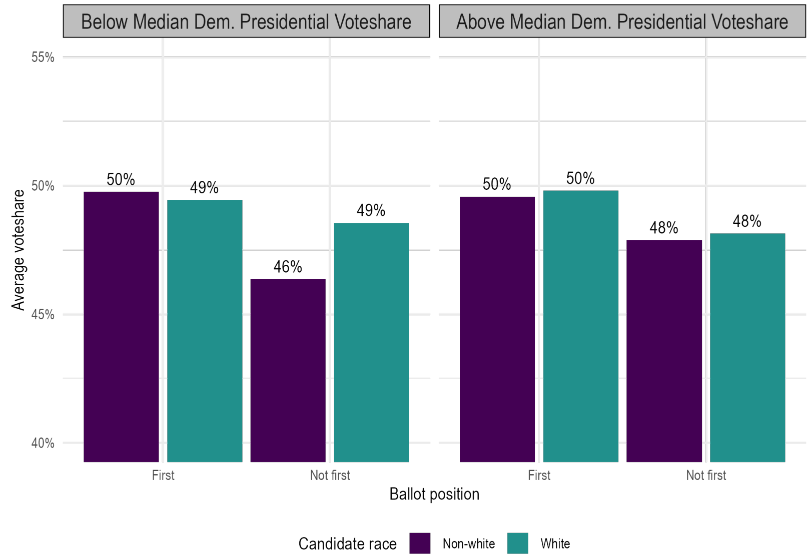

Figure 6: Ballot order, candidate race, and partisanship. Data limited to non-incumbent candidates in races with two candidates only. Vertical axis is truncated.

bottom right).

These results indicate that the greatest degree of heterogeneity in the effects of ballot order by race and gender appears only in more Republican counties. In elections in counties within the lowest quartile of Democratic vote share ($< 50\%$), being listed first on the ballot has much larger effects for non-white candidates (a boost of about 3 percentage points) than for white candidates (about 1 percentage point). Yet in the most Democratic counties—those with Democratic vote share $>72\%$—the differences in the size of ballot order effects by candidates' race is negligible. The middle quartiles are consistent with this pattern, showing less difference as Democratic vote share increases. The shifting effect size is attributable mostly to changes in the non-white ordering premium; while the white boost always hovers around 1 percent (0.8-1.3 percentage points), the non-white boost varies much more (1-3.2 percentage points), consistently decreasing with rising Democratic vote share.

These results are consistent with the theoretical mechanisms that we discussed. Namely, our results suggest that either ideological projection by both Democratic and Republican voters may lead to compression of ballot order effects, or that partisanship may function as a proxy for racial resentment, and that in more Republican counties (where there are

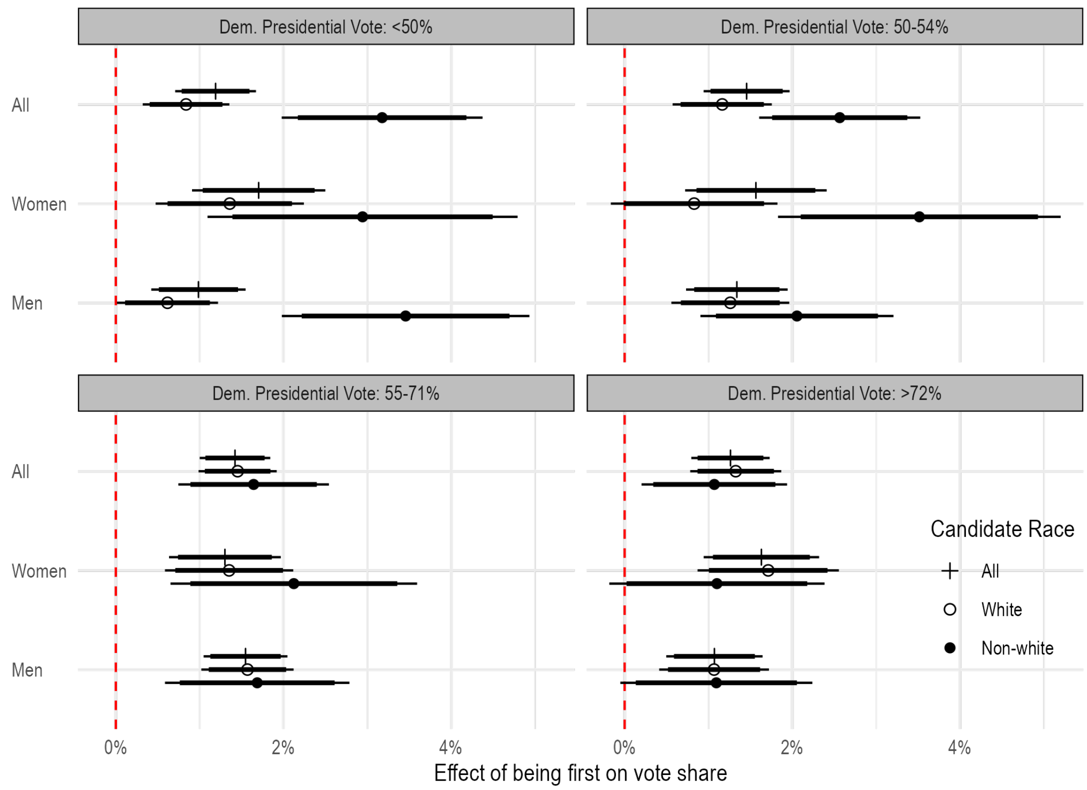

Figure 7: Effect of Ballot Order by Race and Gender and Democratic Presidential Vote Share. Models include fixed effects for the number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

likely higher levels of racial resentment), ballot order can have a greater impact due to lower levels of baseline support for non-white candidates when they are not listed first on the ballot. Further investigation parsing the potentially separate mechanisms at play is a fruitful avenue for future work using micro-level voter data.

District Ethnic Composition

Having examined partisan-based moderation of the heterogeneity in ballot ordering effects by the race and gender of candidates, we next turn to a second possibility: that such heterogeneity is moderated by differences in county-level racial demographics. To examine this possibility, we turn to our data on district ethnic composition. In Appendix F, we display subset analyses by the white population of the county, and for the county-level Hispanic/Latino population in Figure 8 below. Each panel shows the heterogeneity of our ballot order effects by candidate race and gender within a quartile of Hispanic/Latino county-level population share (the least Hispanic/Latino counties in the top left, and the most Hispanic/Latino counties in the bottom right).

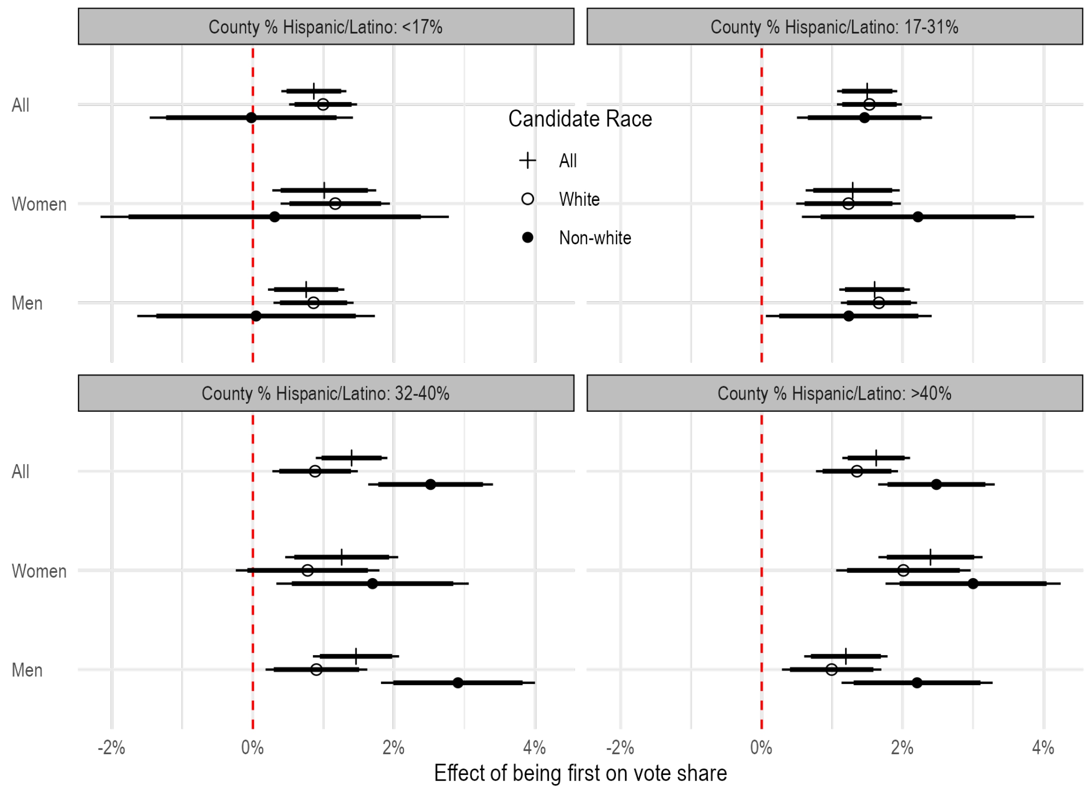

Figure 8: Effect of Ballot Order by Race and Gender and County % Hispanic/Latino. Models include fixed effects for the number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

We find that the heterogeneity of ballot order effects by candidate race is largest in the most Hispanic/Latino counties. The difference in the size of ballot order effects between white and non-white candidates is far larger in counties with above-median proportions of Hispanic or Latino residents—those counties with 32–40% Hispanic/Latino residents and those counties with >40% Hispanic/Latino residents. In the other two quartiles, there is no significant difference between white and non-white effects and, in the lowest quartile counties, the differential is reversed, with non-white candidates receiving zero boost from being listed first, and white candidates receiving the standard boost of about one percentage point (though, again, this difference is not statistically significant, perhaps due to the low number of non-white candidates running in these districts). As with partisanship, the larger difference across races in more Hispanic/Latino counties is due mostly to changes in the effect for non-whites, which increases from 0 to 2.5 percentage points as Hispanic/Latino proportions increase. A more formal analysis of this moderation by county-level Hispanic/Latino population levels using a linear triple-interaction between county %Hispanic/Latino, an indicator for being listed first on the ballot, and an indicator for whether the candidate is non-white, corroborates these subset analyses.[11]

While these findings are exploratory—and they do not provide the same insights that, say, an experiment that measures various kinds of attitudes like linked fate, racial resentment, or symbolic racism might—they suggest that, in keeping with some other scholarship (e.g., McConnaughy, 2010), one driver of our racially heterogeneous ballot position effects may be affinity effects among Hispanic voters.

Discussion and Conclusion

How do seemingly small decisions about election administration influence the outcomes of elections, and how do these administrative procedures interact with voters' biases? Using a natural experiment in California elections, we have examined how the use of a ballot order heuristic varies across types of candidates and types of voters.

We find statistically and substantively significant effects of randomly being listed first on a ballot, corroborating past literature on ballot order. We build on these basic findings by exploring heterogeneity in these effects by the race and gender of candidates. We show that non-white candidates, and especially non-white women, benefit significantly more from being listed first—a random boost in salience—than white candidates. However, we do not find an equally large difference within gender: women benefit slightly, but not significantly

[11] These results are shown in Appendix F.

more from being listed first. These findings may be in keeping with other work, such as that by Lemi (2020) and Rivera-Burgos (2023), which shows that doubly-minoritized individuals can sometimes benefit from unexpected cognitive biases and preferences; but more research is necessary to say whether this is the part of a similar phenomenon or not.

We also find that the size of these effects shifts substantially across information contexts and when we consider other preferences, including ideological alignment and affinity effects. Salience cues appear to have larger effects in local elections held concurrently with presidential elections than in off-cycle elections, especially for women candidates. They also seem to exaggerate the incumbency advantage, but in a way that does not produce heterogeneity by race or gender. Salience cues also seem to help ethnic minority candidates overcome a disadvantage with conservative electorates, but don't change outcomes among progressive electorates. Finally, we see some suggestion that these heterogeneous boosts by race may be produced by affinity effects in heavily Hispanic districts, where no such boost exists in heavily white districts.

While effects of a few percentage points due to being listed first may seem small, they are in line with the main effects found in other work, not only on candidate evaluations, but on election outcomes. First, a meta-analysis by Schwarz and Coppock (2022) of 67 studies estimates a 2 percentage point average marginal effect in favor of women (p. 662). Observational work, which, depending on the context and specification used, often finds a 2-4 percentage point vote share penalty for women: for instance, Fulton (2012) finds that incumbent women fare 3 percentage points worse than incumbent men, ceteris paribus (p. 304). Effects tend to be a similar size by race and ethnicity: Trounstine and Valdini (2008) find the effects of racial district composition and districted (vs. at-large) elections on the probability of electing African American and Latino city councilors range between 1.5-5.5 percentage points (p. 559). Although it is difficult to compare estimates "apples to apples" across so many different contexts, the salience cue estimates we generate here could wipe out many of the disadvantages women and ethnic minorities seem to face at the ballot box. More research will be needed to understand exactly how these effects net out in other contexts.

Indeed, while these effects are interesting and important, it is also critical to note that even in a large state—California accounted for 12% of the total US population in 2024—generalizability is an issue. In part *because* the state is so diverse, the heuristics and stereotypes that voters have about different racial and ethnic groups may be quite different than those held by voters in other parts of the country. For instance, voters in California might assume a Hispanic candidate is left-leaning, but those in Florida might assume the same candidate is right-leaning. Even with the same stereotypes, the context also matters: a Hispanic woman candidate might receive an even bigger boost in New Mexico, and a penalty in Wyoming. We think it is therefore likely even in similar settings (e.g., nonpartisan local races), elsewhere in the country, we might estimate bigger, smaller, or even reversed effects. And of course, where local races are partisan, or higher-salience races where candidates are well-known, we might see even more variation from these estimates. Given that only a handful of states other than California randomize ballot order, such variation may be difficult to study. As the number of states with ballot order randomization has increased dramatically in the past several years, the number of cases available for scholars to study will increase over time. Once this rule has been in place for several election cycles, those states with higher ethnic variation (e.g. New Mexico, New Jersey, and Virginia) might make for ideal comparisons.

That ethnic minority candidates may benefit more from being listed first has important implications for election design that aims to foster better descriptive representation. Ballot order randomization deals with one persistent bias—alphabetic bias. However, the underlying phenomenon remains, which is that people tend to select whomever is listed first at higher rates than those who follow. Moreover, we have shown that ballot ordering effects are more complex than previously understood. Rather than simply satisficing, it appears that many voters respond differently depending on *whose* name is made salient. These effects introduce extra volatility into the odds of historically disadvantaged candidates winning office; while ballot order affects all candidates, non-white candidates may be even more dependent on favorable ordering. Furthermore, as candidates often start their political careers by winning local office, and we find these effects are more pronounced in low-salience and low-information elections, non-white candidates may need to think strategically about the timing of their decision to run for office.

Our findings also raise new questions for ballot order effects in other contexts. Would we see the same large effects for non-white candidates and women randomly listed first in other countries? Or might there be other effects in other contexts that we cannot study here: variation by apparent religion or tribe or caste? We expect that the largest effects should be of characteristics that a) voters care about and have some preferences over and b) they can learn about with fairly little information, as with names. Yet those characteristics will surely differ across contexts—even across states within the same country. Furthermore, what might be the effect of having a substantially better- or worse-informed electorate? Might ballot order effects disappear entirely?

Finally, we note that scholarship on the downstream effects of random-order ballots is relatively small, but suggests that there are meaningful consequences of adopting such laws. In a working paper, Horiuchi and Lange (2019) find a small but significant increase in the likelihood of invalid voting in Australia when the random order is most different from ordering candidates by strength (i.e., in races where the previous highest vote-getter would have been at the top of the ballot if listed by votes received, they are instead listed at the bottom). Lee and Song (2023) find similar results in South Korean elections (some races have random order, and some don't, and invalid votes are higher in the random-order races). Atsusaka (2023) shows that random ballot order affects rankings in ranked choice vote systems, making it a very influential ballot design choice in such an electoral system. And perhaps most relevant, a recent paper by Shi and Singleton (2023) shows, using CEDA data on school boards (a subset of our data), that when an educator is randomly listed first (and thus more likely to win), districts in which they do win significantly increase teacher salaries and decrease openings of charter schools. Random-order ballots therefore appear to have small but meaningful effects on everything from the efficacy of voters to policy outcomes at the local level. As such, we hope that future research will seek to understand both the promises and perils of increased salience and visibility–especially the increased impact of these effects for marginalized groups—and their eventual consequences.

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Supplementary Materials for "Alphabet Soup: Randomized Ballot Order and the Representation of Marginalized Candidates"

A Gender and Race/Ethnicity Coding of Observational Data

Coding gender for the candidates in the CEDA data was relatively straightforward using the 'genderizer' package in R. We supplemented this coding with hand-coding of candidate gender for those candidates where we were unable to generate a probable gender assignment using the automated approach, following Anzia and Bernhard (2022). Taken together, we are able to assign a gender to nearly every candidate.

Coding race and ethnicity for these candidates was more difficult. We improve upon past work by creating joint probabilities that take into account both first names (which are often more indicative for Black individuals) and last names (which are often more indicative for Asian and Hispanic/Latinx individuals).

To code last names, we use the 'wru' package in R, which relies on U.S. Census data to generate continuous probabilities that an individual with a given surname is white, Black, Hispanic, AAPI, Native American, multiracial, or other. We set the 'predict_race' function to take into account that our data is located in California, and within California, localized to the county the individual is running for office in (the probability that the last name "Washington" belongs to a Black individual is higher in Oakland than it is in Marin, for instance). Because both in the Census and even in California whites are the biggest ethno-racial population, this measure will tend to underestimate the likelihood that an individual is non-white.

In our dataset, we show 23,042 unique surnames; out of these, 7,423 were not matched to Census surnames (and so were predicted using a different 'wru' package function).

To code first names, we draw on a dataset from Tzioumis (2018) that uses mortgage data to code the race and ethnicity of first names (Census-derived race and ethnicity data is not available for first names like it is for surnames). This dataset provides continuous probabilities for six racial categories: white, Black, Asian, Native Hawaiian or Pacific Islander, American Indian or Alaska Native, and other; and two ethnic categories (Hispanic/Latino or non-Hispanic). Unlike the 'wru' package, however, it does not have a function to impute race or ethnicity for names that are not exact matches. This means that we have missing values for some first names, which in turn means their racial or ethnic coding is predicted from surnames alone.

In our dataset, we show 7,660 unique first names, 2,628 of which matched to the mortgages dataset (which contains 4,251 unique first names). After matching first names, we condensed the racial and ethnic categories in the mortgage data to match those in the Census surname data where possible (e.g., Asian and Native Hawaiian or Pacific Islander were joined to be equivalent to the AAPI category).

To create joint probabilities, we tried a number of simple procedures. For instance, we tried averaging the first and last name probabilities to create new continuous averages for each race/ethnicity variable. We ended up settling on the following formula: First, we categorized first names and surnames based on whichever racial or ethnic category had the highest predicted probability (e.g., "Alison" is classified as white with a roughly 96% probability of being white, while "Alka" is coded as AAPI with a roughly 81% probability of being AAPI). Next, ties were extremely rare (the probabilities generated extend to several decimal points) and were broken in favor of the non-white category (if tied between white and a non-white category) and randomly broken if tied between two non-white categories. We decided to break non-randomly for non-white categories because the vast majority of names have white as the highest probability (and thus make us likely to overestimate the number of white candidates).

This process ensures each person receives one ethnoracial coding for their first name (if non-missing) and one for their last name (if non-missing). To create a single ethnoracial coding, we then take several steps. First, we treated codings as unambiguous if both names fell in the same category: e.g., "Jamal Ojukwu" would have dual codings as Black and therefore be classified overall as Black. The vast majority of individuals have a match, since white is the most common classification for both first and last names. Individuals with no first name coding defaulted to the surname coding.

Next, we focused on codings that were ambiguous (e.g., "Jamal Jones"). Candidates with a high probability of a Native surname and any first name were condensed into an Other/Multiracial category. Among the remainder, we then took the coding of the first name if the coding of the first name matched the second-highest probability last name (Black, in this case). Next, we took the coding of the surname if it matched the second-highest probability coding for the first name (e.g., "Sunny Chang" gets coded as AAPI). We then incorporate gender: women who have not yet been classified are coded as their first name's category, based on the fact that women are more likely to take their partner's names after marriage than men and so may have a higher rate of surname classification errors. We then classify the men who remain uncoded based on their surname.

B Effects as a Function of Number of Candidates

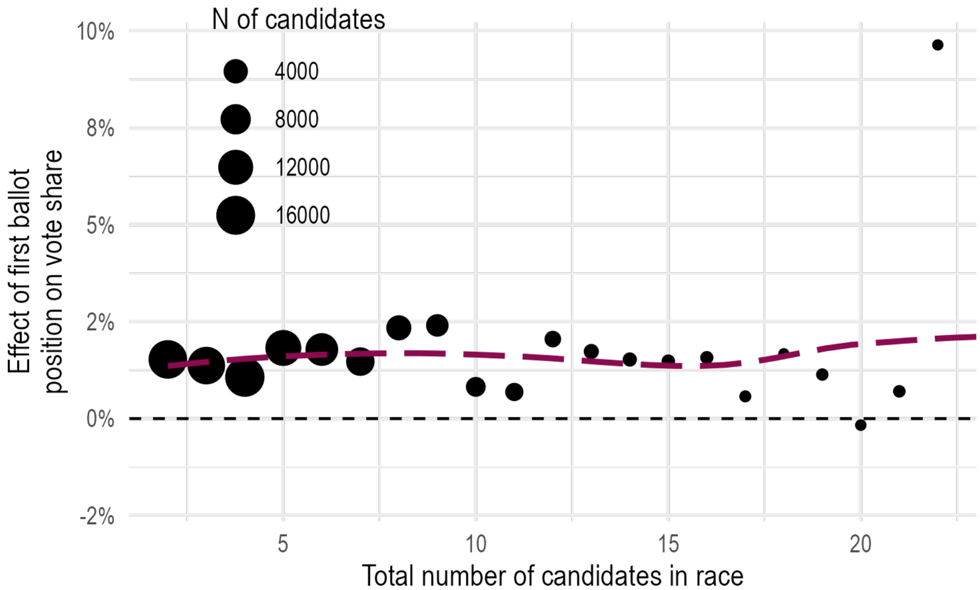

Figure A1: Size of Ballot Position Effects by Number of Candidates. Model includes no other controls.

To disentangle this basic feature of electoral competitiveness from the effect of ballot position, we examine how ballot position influences candidates' success across races with different total numbers of candidates in the race. Figure A1 plots the difference in vote share between the candidate appearing first on the ballot and the average vote share received by all other candidates in the race, as a function of the total number of candidates in the race. We see fairly consistent effects of being randomly listed first across elections of all sizes.

C Heterogeneity of Effects by Office of Election

Our analyses in the main body of the manuscript control for the office of the election using fixed effects. We do so because we recognize that the effects of ballot position may differ in races for different types of offices, and also that race and gender differentials may operate differently for these different types of offices. In this section, we present analyses separately analyzing the effect of ballot position among these different offices.

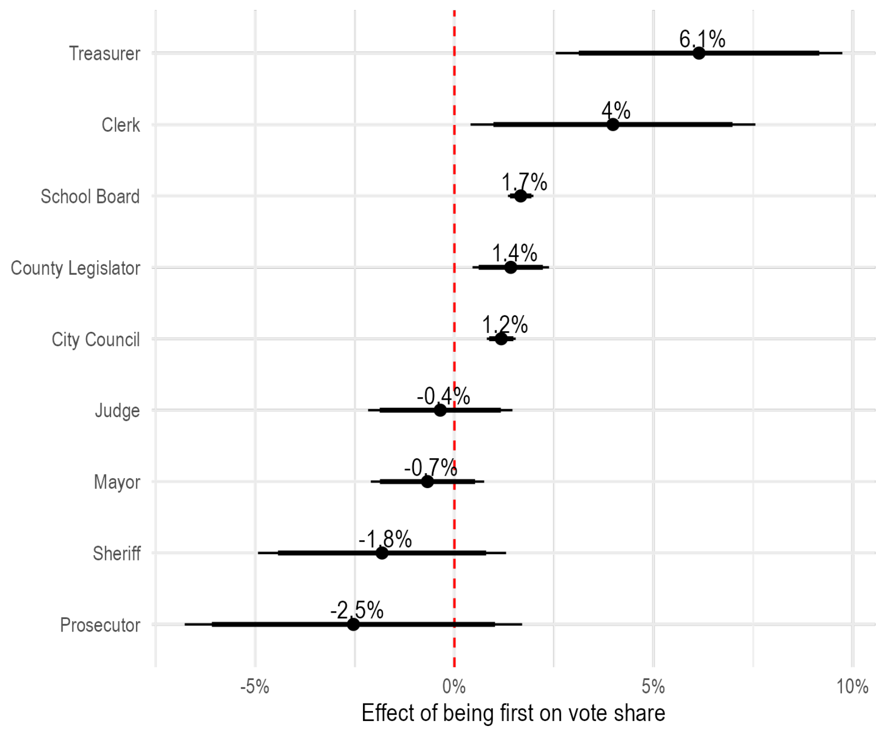

Figure A2: Effect of Ballot Order by Office. Models include fixed effects for number of candidates, year, and county. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

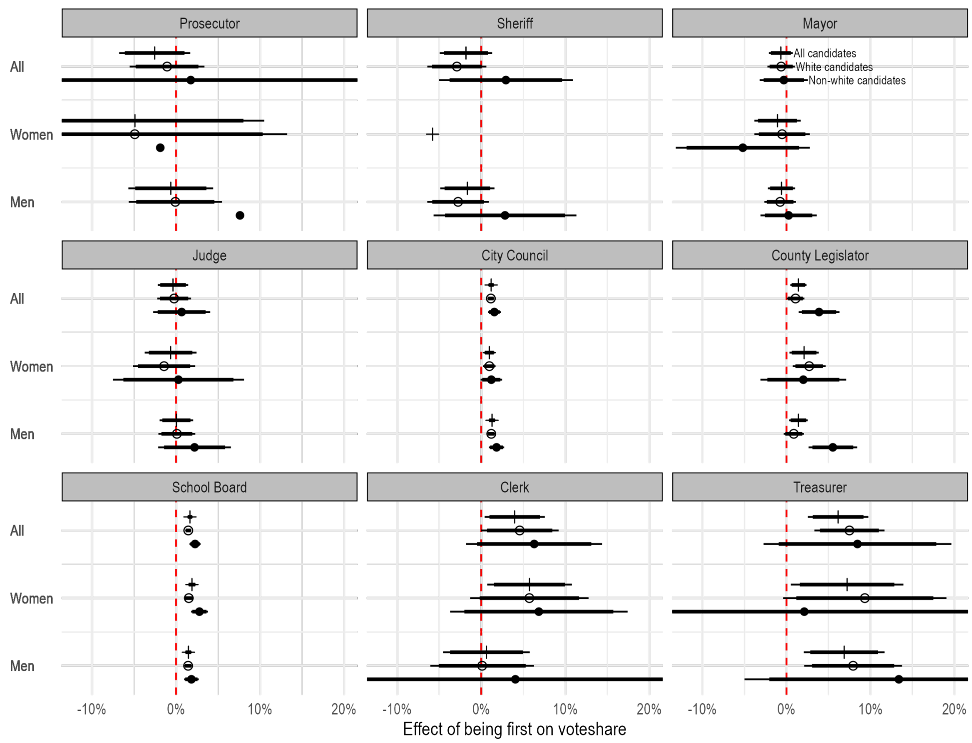

Figure A3: Effect of Ballot Position by Candidate Race and Gender and by Office. Models use standard errors clustered by contest, include fixed effects for number of candidates, year, county, and office, and are limited to non-incumbent candidates only. Thicker bars depict 90% confidence intervals and thinner bars, 95% confidence intervals.

D Heterogeneity of Effects by Time Period

The effects of ballot position that we explore in this paper are likely to vary in size due to a number of factors. One of the theoretical mechanisms behind the ballot position effects we observe is the low levels of attention that voters pay to electoral choices, especially at the local level. Voter levels of attention, of course, might vary by other features of the electoral context. While we explore whether election timing is one of these features in the main body of the manuscript, another contextual feature of the choice environment that might influence the attention voters pay to their choices and thereby the size of the ballot position effects is whether they are voting in-person (with limited time and attention) rather than by mail or early in-person (with, presumably, more time and attention). Voting early by mail gives voters the chance to spend a longer period of time making choices, perhaps consulting a voter guide, researching candidates on the internet, or consulting friends and family both before and while casting their ballots (Kleinberg and Lau, 2021; Reedy, Gastil, and Moy, 2016; Wolak and Stapleton, 2025).[12]

The massive shift in the United States toward vote-by-mail over the last two decades may have potentially changed the degree to which candidates' ballot order matters given that voters now have more time to make an effortful vote choice. While we do not have data detailed enough to investigate the difference in ballot order effects between in-person voters and voters who utilized mail-in-ballots, the State of California does publish rates of in-person and by-mail voting in statewide elections.[13] Since 2012, the state has had rates of vote-by-mail that are greater than 50%. We use this year as a rough dividing line between a later period that has featured high levels of vote-by-mail and an earlier period that featured lower levels of vote-by-mail. We divide our sample of elections between these two periods and conduct subset analyses to replicate our main analyses, and present these results in

[12] Though see, e.g., Hopkins et al. (2021) for evidence that vote-by-mail can also increase rates of problematic ballots that could be rejected by election administrators or Meredith and Malhotra (2011) for evidence that vote-by-mail has higher rates of ballots with errors.

[13] See `https://www.sos.ca.gov/elections/historical-absentee`.

Figure A4 as well as divided by race and gender of candidates in Figure A5. These analyses show little differences in the size of the effects of ballot position between the pre-2012 period and the period since 2012.

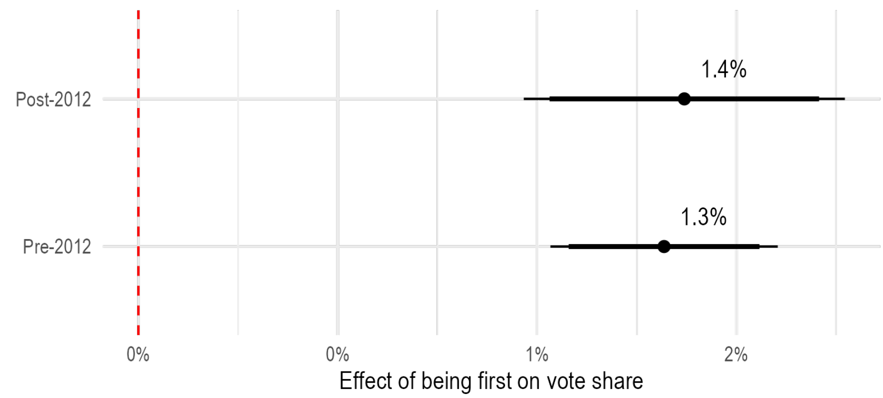

Figure A4: Effect of Ballot Order by Time Period. Models include fixed effects for number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

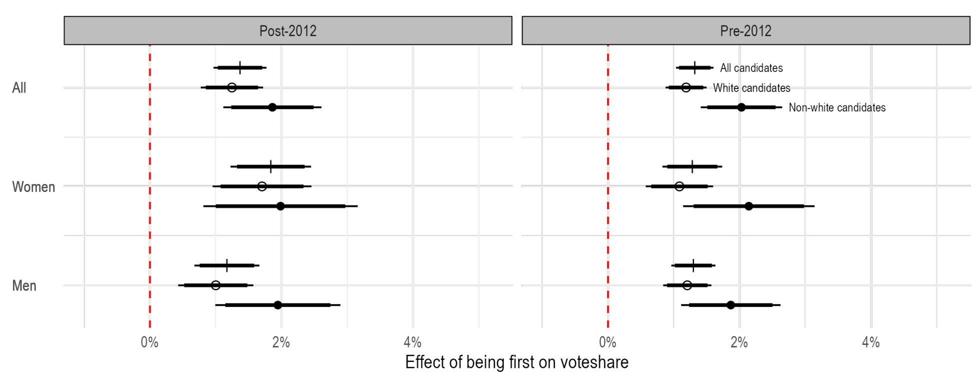

Figure A5: Effect of Ballot Order by Time Period. Models include fixed effects for number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

E Analyses Using Alternative Outcome of Win Probability

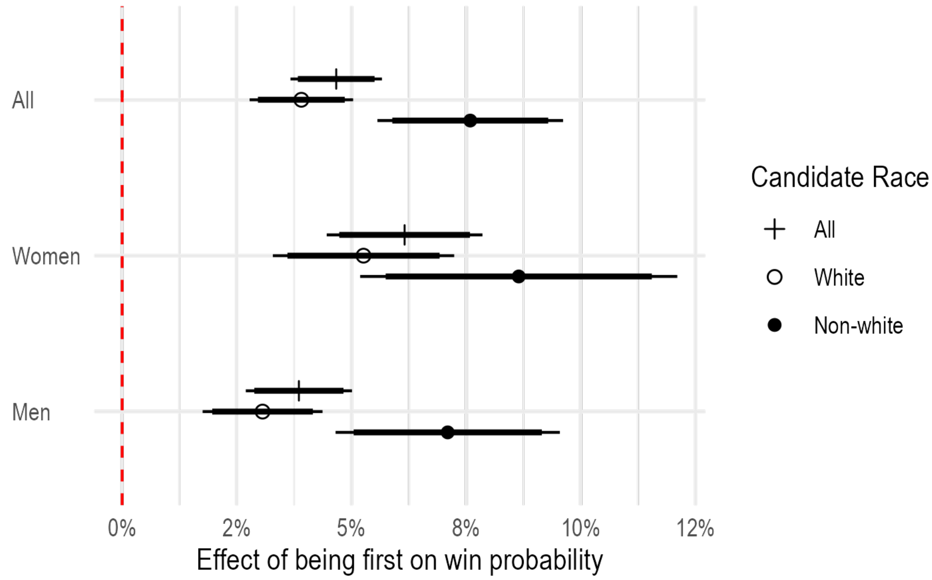

Figure A6: Effect of Ballot Position on Win Probability by Candidate Race and Gender. Models use standard errors clustered by contest, include fixed effects for number of candidates, year, county, and office, and are limited to non-incumbent candidates only. Thicker bars depict 90% confidence intervals and thinner bars, 95% confidence intervals.

F Heterogeneity of Effects by County Racial Demographics

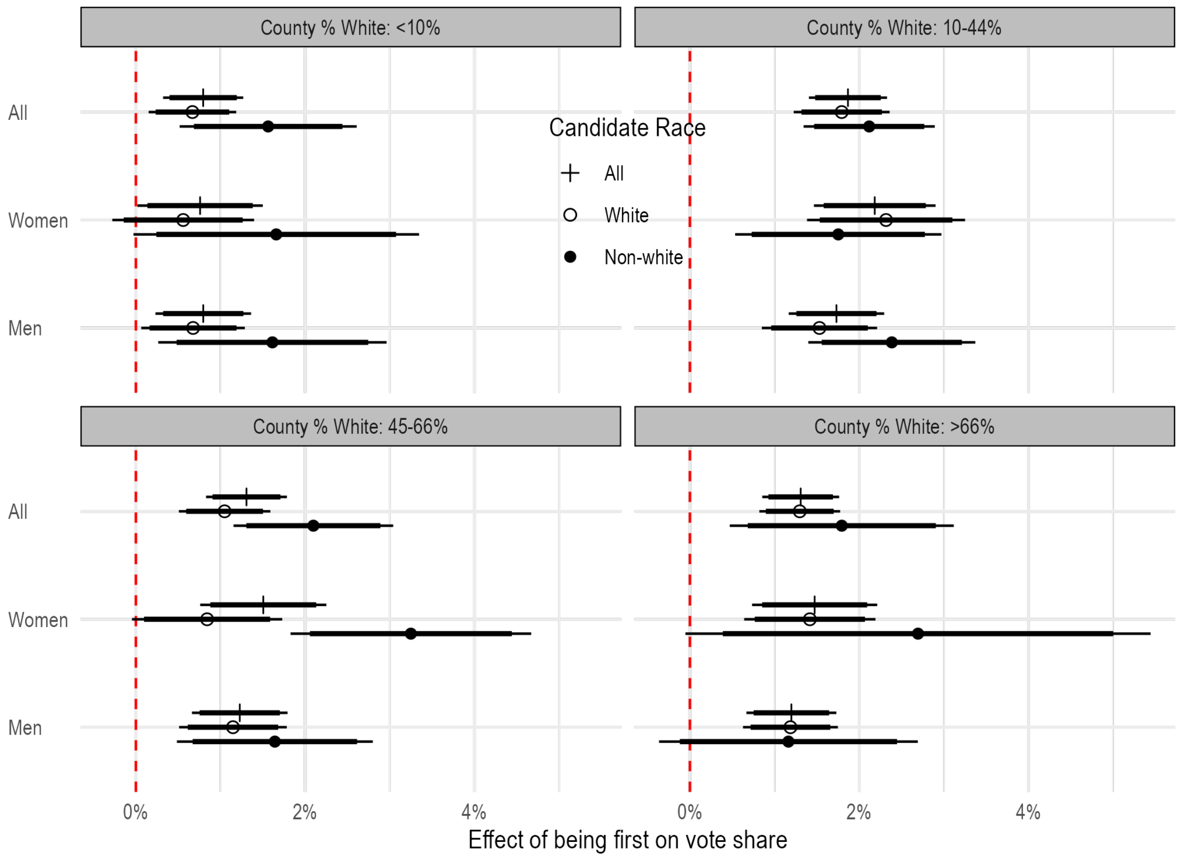

Figure A7: Effect of Ballot Order by Race and Gender and County % White. Models include fixed effects for number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

Table A1: Effect of Ballot Order by Race and Gender and County % White

Dependent Variable:	Voteshare for candidate
Model:	(1)
Variables	
First on ballot	0.018***
	(0.004)
White candidate	0.003
	(0.002)
County % white	0.004
	(0.006)
First on ballot × White candidate	-0.007
	(0.005)
First on ballot × County % white	0.004
	(0.009)
White candidate × County % white	-0.004
	(0.004)
First on ballot × White candidate × County % white	-0.002
	(0.010)
Fixed-effects	
Number of candidates	Yes
Year	Yes
County	Yes
Office of election	Yes
Fit statistics	
Observations	58,792
R^2	0.599
Within R^2	0.004

Clustered (new_raceid) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*

Table A2: Effect of Ballot Order by Race and Gender and County % Hispanic/Latino

Dependent Variable:	Voteshare for candidate
Model:	(1)
Variables	
First on ballot	-0.0002
	(0.006)
White candidate	0.002
	(0.003)
County % Hispanic/Latino	-0.013
	(0.009)
First on ballot × White candidate	0.012*
	(0.006)
First on ballot × County % Hispanic/Latino	0.052***
	(0.015)
White candidate × County % Hispanic/Latino	-0.004
	(0.007)
First on ballot × White candidate × County % Hispanic/Latino	-0.051***
	(0.016)
Fixed-effects	
Number of candidates	Yes
Year	Yes
County	Yes
Office of election	Yes
Fit statistics	
Observations	58,792
R^2	0.599
Within R^2	0.004

Clustered (new_raceid) standard-errors in parentheses
*Signif. Codes: ***: 0.01, **: 0.05, *: 0.1*

G Effects of Other Ballot Positions

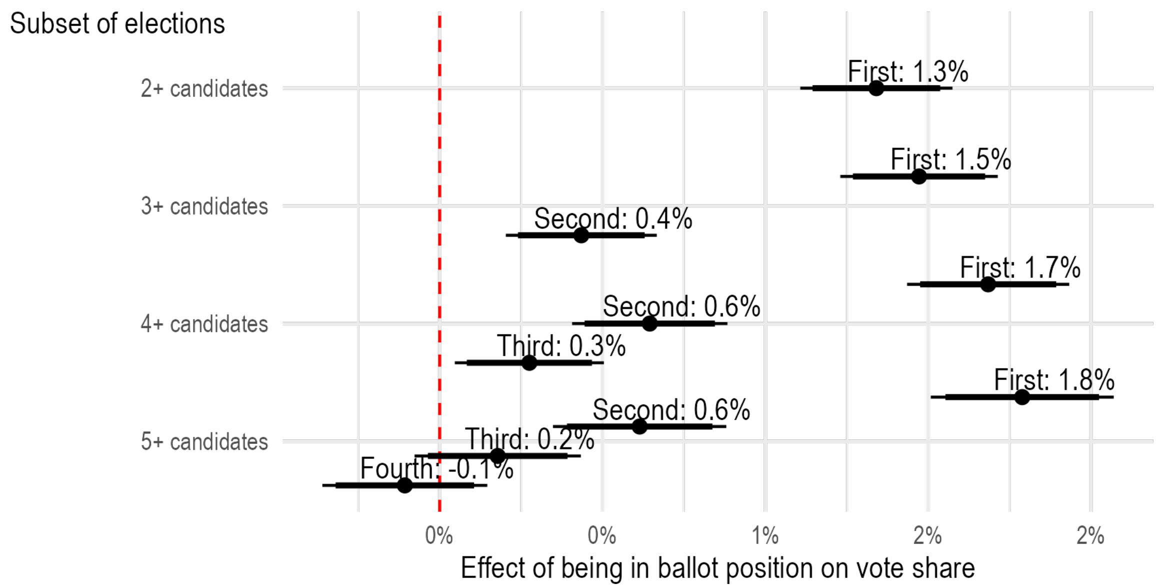

Figure A8: Effect of Ballot Order for All Ballot Positions. Models include fixed effects for number of candidates, year, county, and office. Only non-incumbents included. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.

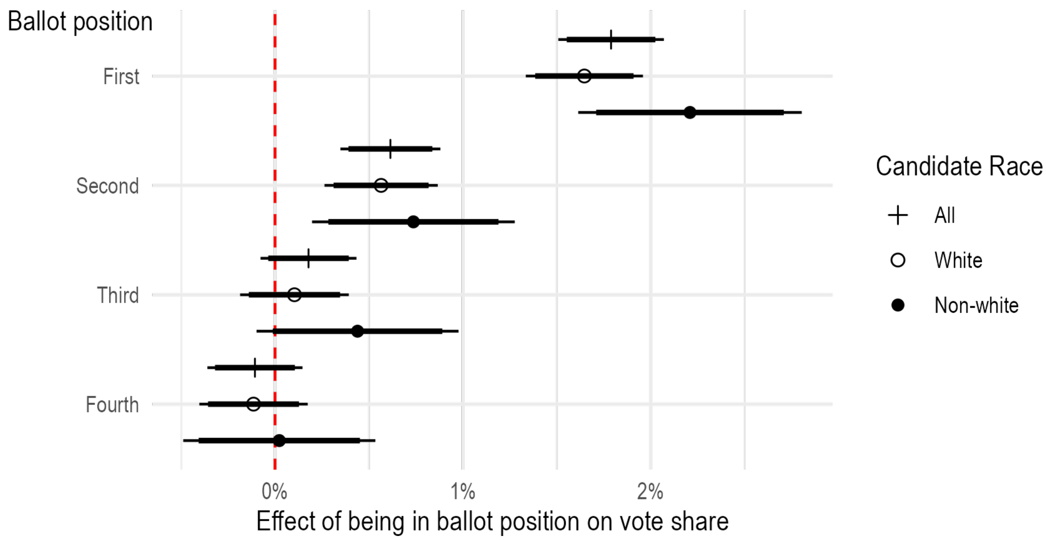

Figure A9: Effect of Ballot Order by Race and Gender for All Ballot Positions. Models include fixed effects for number of candidates, year, county, and office. Only non-incumbents included, and only races with 5+ candidates running. Bars depict 90% (thick lines) and 95% (thin lines) confidence intervals, and standard errors in all models are clustered at the unique contest-level.