Congressional Agency Control: The Impact of Statutory Partisan Requirements on Regulation

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Abstract

Presidential and congressional influence over regulatory agencies forms a central topic in the study of bureaucratic politics and administrative law. One mechanism of congressional control consists of formal constraints on the president’s appointment power: enabling statutes for independent regulatory commissions often cap the number of same-party (“party-line”) commissioners a president may appoint, forcing presidents to appoint “cross-party” commissioners. This paper provides the first systematic empirical evidence on the impact of partisan requirements. I present evidence from a new dataset of published adjudications and rulemakings in the Federal Communications Commission from 1965–2006. This dataset is the largest and most complete ever assembled, documenting 94,693 votes by 46 commissioners in 17,879 adjudications and rulemakings. Using a Bayesian multilevel ideal point model for mixed ordinal votes, I find that partisan requirements may have considerable effects on substantive policy outcomes. The evidence squarely rejects theories of presidential dominance in the appointments game. But effects are in part unanticipated: evidence suggests that after 1980, cross-party appointees are even more extreme than party-line appointees, pointing to a sharp rupture in Senatorial deference in the 1980s.

1 Introduction

Independent regulatory commissions affect a staggering array of day-to-day activities. Commissions draft binding rules, adjudicate contentious disputes, and set the terms of competition in a wide variety of economic and political matters. Administrative law and the study of bureaucratic policymaking are concerned most basically with the accountability of such agencies to the judiciary, the executive, and the legislature. One central mechanism of perceived legislative control consists of statutory partisan requirements, which place formal constraints on how many commissioners of the same party the president may appoint to independent regulatory commissions. The Federal Communications Commission (FCC), for example, currently

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consists of five members, but its enabling act limits the number of commissioners from the same party to three. These partisan requirements are lauded as “critical characteristics” of institutional design, as “hallmarks of independence” (Office of Legal Counsel 2003; May 2006 p. 444 n.83), and exist for countless commissions at state and federal levels. Appointments to the Securities and Exchange Commission, the Federal Trade Commission (FTC), the Federal Energy Regulatory Commission, the Federal Election Commission (FEC), and most state redistricting commissions, to name a few, are constrained in this fashion.

Yet do these requirements achieve anything at all? Was the Congress rational in its design of agencies by expressly enacting partisan requirements?

The effect of partisan requirements has substantial implications for administrative law and the study of bureaucratic behavior. Scholars and textbooks treat these requirements as readily justified, effective instruments of congressional control on executive appointment power (e.g., Lewis 2004, p. 47-48; Mackenzie 1981, p. 195-99; Mashaw et al. 2003, p. 196; Strauss 2002, p. 96; Sunstein 2000, p. 103; Vermeule 2004). The U.S. Senate Committee on Government Operations (1977, p. 31) similarly concluded that the requirements are “justified as an important restraint on the President.” Some even advocate the expansion of partisan requirements beyond the territory of regulatory commissions. For example, Tiller and Cross (1999) extols the virtues of partisan requirements “to protect against ideological excess by one party” and advocates partisan balance in all federal circuit court panels. A small minority contends that partisan requirements unconstitutionally violate presidential appointment power (e.g., Office of Legal Counsel 1989; Rappaport 2001; Raskin 2000). These views share the assumption that partisan requirements substantively affect appointments and thereby policy.

Others dismiss this assumption, arguing instead that partisan requirements only nominally constrain appointments (e.g., Zarkin and Zarkin 2006, p. 54). The criticism takes two forms. First, in the optimism of the New Deal period, some believed that partisanship, even as an empirical matter, had no effect on commission behavior (Mansfield 1932). Cushman (1941, p. 752-753) thereby concluded that the requirements are “neither sound nor relevant” and “obsolete and unnecessary.” Second, a more recent criticism – one implicitly rejecting the New Deal assertion of the irrelevance of partisanship – charges that partisan requirements demand only that presidents appoint commissioners differing formally in partisanship but who are otherwise identical in viewpoint to the president: cross-party commissioners come in sheep’s clothing (e.g., Krasnow et al. 1982, p. 39-41, 67-68; Lichty 1961, p. 25-26; Mackenzie 1997, p. 31; Morrison 1988; Williams Jr. 1976). Anecdotal evidence abounds. When a Democratic FCC seat opened up several months into President Nixon’s second term, public interest groups charged that the nominee James Quello – a “nominal Demo-
“crat” and broadcaster who had actively lobbied for the industry for many years – was hopelessly out of touch with consumer interests, and assailed him as a Nixon loyalist unfit for the Democratic seat. Adding fuel to the fire was the revelation that Quello had contributed $1,100 to Nixon’s reelection campaign (Graham and Kramer 1976, p. 357–367). Similarly, two Reagan appointees on the FTC, the (self-described) independent Azcuenaga and Democrat Douglas, reportedly formed a “solid conservative majority” with Republican Miller. The FTC bloc emerged despite the fact that there were already two Republicans (one a Reagan appointee) serving on the commission, leading one Congressmember to decry why Reagan independents weren’t nominated to Republican seats.  

The view that partisan requirements may achieve nothing at all has found its way into case law as well. The D.C. Circuit found injusticiable a constitutional challenge to partisan requirements of the FEC because “it is impossible to determine in this case whether the statute actually limited the President’s appointment power.” Similarly, a 1976 Senate Commerce Committee Report, qualitatively studying 51 appointees to the FTC and FCC from 1952–1973, found that no president initially appointed “bonafide, honest-to-God members of the other party.” Even worse than finding cross-party appointees to arrive in sheep’s clothing, the report found that party-line appointees are typically “abler” than cross-party appointees. While partisan requirements had none of the intended effects of achieving substantive ideological heterogeneity, the report found that they had the additional unintended consequence of decreasing the quality of appointees. It thereby concluded that “there is good reason for abolishing [partisan requirements] altogether” (Graham and Kramer 1976).

At heart of the debate is an assumption about the empirical effects of partisan requirements. Yet beyond anecdotal evidence of commissioners such as Quello, Azcuenaga, and Douglas, no systematic empirical evidence exists as to whether these partisan requirements matter, and, if so, how. This paper presents a first effort to systematically analyze the effect of partisan requirements. I first examine tentative evidence across independent regulatory commissions. I then compile a new dataset of 94,693 FCC votes by 46 commissioners in 17,879 published adjudications and rulemakings from 1965–2006. This represents by far the largest and most comprehensive database of commissioner voting behavior to date. I use a Bayesian multilevel ideal point model of mixed ordinal votes to assess whether commissioner partisanship explains voting behavior.

The findings are threefold. First, I find weak evidence that partisan requirements increase the raw num-

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3 FEC v. NRA Political Victory Fund, 6 F.3d 821, 824 (D.C. Cir. 1993) (emphasis in original); See also National Committee of Reform Party of U.S. v. Democratic National Committee, 168 F.3d 360, 365–67 (9th Cir. 1999) (refusing to reach merits on challenge to FEC partisan requirement for lack of redressability).
ber of cross-party appointees. Although virtually all FCC cross-party appointees over the past 40 years fill seats mandated by partisan requirements, the cross-party appointment rate is only moderately higher than in the few commissions without partisan requirements. This could be explained by the fact that alternative political channels exist for interest groups and the Congress to secure politically-favorable appointees. Second, the effect of commissioner ideology on voting is profound. Commissioner partisan affiliation exhibits robust and large predictive power over votes, even holding constant the party of the appointing president. This strongly corroborates that partisan requirements genuinely constrain presidents and rejects the notion that expertise exclusively drives decisionmaking. Lastly, I find that cross-party appointees may be more extreme than party-line appointees in the post-1980 period. Republican presidents appear to appoint Democrats who are even more liberal than Democrats appointed by Democratic presidents (and vice versa). The primary explanation of these findings appears to be congressional influence in the selection and oversight of regulatory commissions, and a sharp rupture in Senatorial deference in the 1980s.

The paper proceeds as follows. Section 2 discusses this study’s contribution to research on bureaucratic politics, ideal points, and the FCC. Section 3 provides an overview of data across commissions and specific to the FCC. Section 4 outlines a Bayesian statistical model to learn about commissioner policy preferences, results of which are presented in Section 5. Section 6 presents a series of sensitivity analyses, and Section 7 concludes.

2 Extant Research

This study contributes to extant research in three ways. First, at the most general level this study contributes to research on bureaucratic politics and administrative law (Moe, 1989). Partisan requirements are frequently theorized about in this literature. Lewis (2004) proposes and tests a theory of the institutional determinants of agency design, including partisan requirements. Partisan requirements, from this view, are tools of congressional agency design to ensure politically-favorable outcomes. Sunstein (2005, p. 1021) theorizes that partisan requirements may be “an effort to protect against . . . deliberative pathologies that are likely to result if deliberations are restricted to like-minded people.” Yet not a single study directly assesses the effects of partisan requirements through systematic empirical investigation. This study examines the evidence for how, whether, and to what degree agency decisionmaking can be understood in political terms, and to what degree the Congress, using formal statutory legal constraints, can affect presidential appointment power and agency behavior.

More specifically, the FCC provides a direct test case for presidential and legislative dominance theo-
ries as applied to appointments to administrative agencies. Presidential dominance theories maintain that the executive retains substantial if not unfettered control over agencies, even formally independent regulatory commissions (Moe, 1982; Goodsell and Gayo, 1970). Because presidents have the power to appoint and promote commissioners, to name the chairperson, to recommend agency budgets, to exercise litigating authority via the Department of Justice, and to otherwise intervene in regulatory affairs, scholars have hypothesized that presidents retain dominant influence in agency policymaking. Given the conditions of most independent regulatory commissions, an observable implication from the presidential dominance perspective is that cross-party commissioners don sheep’s clothing: due to the prominent role of the executive in the appointments process, and the limited congressional role, commissioner votes should be explained primarily by the ideology of the appointing president, not the partisan affiliation of the commissioner.

Legislative dominance scholars, on the other hand, stress that due to the Senate’s advice and consent power in appointments and congressional fiscal, oversight, and reversal power, agency behavior is determined largely by the Congress (McCubbins, 1985; Krasnow et al., 1982; Weingast and Moran, 1983). Agencies are creatures of Congress. Wiley (1988), a former FCC Chairman, for example, asserts that the most powerful political influence stems from Congress in the form of appropriations and oversight committees (particularly the Senate Committee on Commerce, which oversees FCC appointments) and individual Congressmembers. If congressional dominance is right, and the Congress successfully constrains executive appointments by statute and advice and consent, we should observe that partisan requirements have teeth: the party of the commissioner should explain voting heterogeneity, even holding constant the party of the president.

Of course, presidential and legislative dominance theories need not be mutually exclusive. Both the president and Congress may play a role in the appointments game (Hammond and Hill, 1993; Snyder and Weingast, 2000). The observable implication should be that we find both presidential and commissioner partisan affiliation to have predictive power over appointments.

Second, this study contributes to research seeking to measure policy preferences (ideal points) of bureaucratic actors and institutions (e.g., Bailey and Chang, 2001). Moe (1985) and Snyder and Weingast (2000) measure commissioner-specific ideology estimates for the National Labor Relations Board (NLRB) based on the ratio of pro-business decisions in unfair labor actions starting in 1949. Nixon (2004) uses common space scores of Poole (1983), available for elected Congressmembers, to study policy preferences of officials that have served both in appointed and elected positions, but not those serving on multi-member independent regulatory commissions. This paper enriches this research in several ways. I illustrate methods that use actual votes by commissioners to estimate policy preferences without reading and hand-coding
substantive outcomes, a costly and error-prone process. This approach, derived from item-response theory, is well-known to accord with a stochastic utility model of voting (Clinton et al., 2004; Krehbiel and Rivers, 1988), but to date has not been employed to study commission voting behavior. This study thereby examines a key substantive output of agencies, which to date has been difficult to incorporate systematically: agency votes on adjudications and rulemakings (but cf. Moe, 1985; Snyder and Weingast, 2000).

Estimates of commissioner policy preferences and valid uncertainty intervals (Bayesian analogues of confidence intervals) are provided for all commissioners serving from 1965–2006. This broad scope is possible because of the staggered nature of commission terms, permitting us to draw valid inferences about ideology across commissioners and across time, and avoiding the potential of endogenously filed cases (Moe, 1985). This paper also extends dichotomous ideal point models to incorporate more substantive voting information characteristic of quasi-judicial votes. The four voting alternatives available to commissioners (e.g., majority, concurrence, partial dissent, full dissent) convey more information than dichotomous legislative roll call votes. By and large, previous studies have not incorporated such alternatives, forcing scholars to discard substantial amounts of information.

Third, this study also speaks directly to the rich literature on the FCC (e.g., Krasnow et al., 1982). While prior research has examined partisan determinants of FCC output in certain limited contexts and in different time periods, the evidence as to the effects of partisan requirements is mixed. Cohen (1986) analyzes a sample of 10 cases per year from 1955-74, finding that when controlling for presidential affiliation, the party of the commissioner has no explanatory power over votes. This finding would suggest that partisan requirements do not constrain presidents. Gormley (1979) examines some 300 cases decided from 1974-76 to evaluate the revolving door hypothesis, and finds that commissioner partisanship explains a higher proportion of votes than prior employment. Despite the fact that all commissioners were appointed by a Republican president, Gormley concludes that “regulatory agencies could change dramatically if presidents were permitted to appoint an unlimited number of members of their own party to regulatory agencies” (p. 681). Canon (1969) examines 665 non-unanimous FCC votes from 1963–67, finding that commissioner and presidential partisan affiliation predict voting patterns, but only in the areas of programming standards and licensing criteria. Nagel and Lubin (1964) studies 100 commissioners from a number of commissions including the FCC from 1936, 1946, and 1956. Assigning each commissioner a score proportional to the number of times she voted in a liberal direction, Nagel and Lubin finds (albeit without reporting uncertainty estimates) that commissioner and presidential partisan affiliations predict liberalism.

Although these FCC studies speak to some degree about the effects of partisan requirements, they pro-
vide conflicting evidence, surely in part because they were not designed with such an assessment in mind. For example, Cohen uses partisanship only as a control with other covariates to examine the impact of industry employment on commissioners, which may confound inferences about partisanship. While Gormley concludes that partisan requirements constrain, Gormley’s data includes only data from a single president, thereby making it hard to draw conclusions about whether the variation between commissioners appointed by the same president is drowned out by variation across different presidents. A common limitation of these studies is that all are limited to a subset of cases, commissioners, and presidential administrations. No study examines cases past 1976. I build on this research by compiling the most comprehensive dataset on FCC voting behavior and commissioners to date and employing modern ideal point methods to fully incorporate voting information.

3 Data

Subsection 3.1 presents data across commissions. As a first cut, I examine whether partisan requirements are correlated with higher cross-party appointment rates. To be sure, cross-commission information is severely limited, and I refrain from drawing any strong inferences from this tentative evidence. Nonetheless, this information is crucial to clarify the conditions for and limits of what we can learn about the impact of partisan requirements. Subsection 3.2 introduces, and the rest of the paper focuses on, a comprehensive FCC dataset to more systematically analyze intra-commission implications of partisan requirements.

3.1 Across Independent Regulatory Commissions

Partisan requirements are conventionally seen to have one principal effect: increasing the total number of cross-party commissioners beyond the number that would otherwise be appointed. But even this effect is not a foregone conclusion. Interest groups, parties, and the Congress may pressure the president to appoint cross-party commissioners even in the absence of partisan requirements, and presidents may independently value the appearance of bipartisanship (see, e.g., Mackenzie [1981] p. 23-24). Indeed, it is even conceivable that formal requirements undercut the ability of vested parties to lobby for more cross-party appointments. If so, formal requirements could ironically cap the number of cross-party appointees or merely reduce its variance over time.

To gain a preliminary understanding of the link between statutory requirements and the frequency of cross-party appointees, Table 1 presents information on major (federal) independent regulatory commissions. If partisan requirements are binding, we should see that the cross-party rate, the number of cross-
<table>
<thead>
<tr>
<th>Agency</th>
<th>Years Active</th>
<th>Number of Seats</th>
<th>Partisan Requirement</th>
<th>Term Length (years)</th>
<th>Number of Commissioners</th>
<th>Cross-Party Rate</th>
<th>But-For Rate</th>
<th>Dissent/Concurrence Rate</th>
<th>Constraint Ceiling (p-value)</th>
</tr>
</thead>
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<tr>
<td>Federal Trade Commission (FTC)</td>
<td>1914–</td>
<td>5</td>
<td>3</td>
<td>103</td>
<td>4.1</td>
<td>0.35</td>
<td>0.77</td>
<td>12</td>
<td>0.00</td>
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<tr>
<td>Securities and Exchange Commission (SEC)</td>
<td>1934–</td>
<td>5</td>
<td>3</td>
<td>83</td>
<td>2.5</td>
<td>0.30</td>
<td>0.88</td>
<td>6</td>
<td>0.00</td>
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<tr>
<td>Federal Communications Commission (FCC)</td>
<td>1934–83</td>
<td>7</td>
<td>4</td>
<td>58</td>
<td>4.2</td>
<td>0.33</td>
<td>0.68</td>
<td>172</td>
<td>0.00</td>
</tr>
<tr>
<td></td>
<td>1983–86</td>
<td>5</td>
<td>3</td>
<td>19</td>
<td>3.3</td>
<td>0.37</td>
<td>1.00</td>
<td>56</td>
<td></td>
</tr>
<tr>
<td></td>
<td>1979–92</td>
<td>5</td>
<td>3</td>
<td>20</td>
<td>2.7</td>
<td>0.35</td>
<td>0.71</td>
<td>67</td>
<td>0.10</td>
</tr>
<tr>
<td>Federal Energy Regulatory Commission (FERC)</td>
<td>1979–92</td>
<td>5</td>
<td>3</td>
<td>11</td>
<td>3.8</td>
<td>0.45</td>
<td>0.60</td>
<td>68</td>
<td></td>
</tr>
<tr>
<td>Federal Election Commission (FEC)</td>
<td>1975–</td>
<td>6</td>
<td>3</td>
<td>18</td>
<td>5.1</td>
<td>0.56</td>
<td>1.00</td>
<td>4</td>
<td>NA</td>
</tr>
<tr>
<td>Occupational Health &amp; Safety Review Commission (OSHRC)</td>
<td>1971–</td>
<td>3</td>
<td>None</td>
<td>22</td>
<td>3.4</td>
<td>0.14-0.23</td>
<td>NA</td>
<td>62</td>
<td>NA</td>
</tr>
<tr>
<td>Atomic Energy Commission (AEC)</td>
<td>1946–74</td>
<td>5</td>
<td>None</td>
<td>36</td>
<td>2.4</td>
<td>0.44</td>
<td>NA</td>
<td>?</td>
<td>NA</td>
</tr>
<tr>
<td>Nuclear Regulatory Commission (NRC)</td>
<td>1974–</td>
<td>5</td>
<td>3</td>
<td>25</td>
<td>4.1</td>
<td>0.44</td>
<td>0.45</td>
<td>3</td>
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<tr>
<td>National Labor Relations Board (NLRB)</td>
<td>1935–47</td>
<td>3</td>
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<td>10</td>
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<td>NA</td>
<td>58</td>
<td>NA</td>
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<tr>
<td></td>
<td>1947–</td>
<td>5</td>
<td>None</td>
<td>51</td>
<td>3.3</td>
<td>0.29</td>
<td>NA</td>
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</table>

Table 1: Characteristics of Selected Federal Agencies. Years Active denotes the years a commission was active subject to particular composition requirements; Number of Commissioners denotes the number of commissioners who served during the time period; Partisan Requirement indicates the maximum number of commissioners from the same party, if applicable; Term length indicates the length of an appointment spell; Average Service Spell indicates the average length of time served by a commissioner (per appointment); Cross-Party Rate indicates the proportion of commissioners who are cross-party appointees (on commissioner basis); But-for Rate indicates the proportion of cross-party commissioners that were clearly subject to partisan requirements as described in text; and Dissent/Concurrence Rate indicates the number of cases with any dissent or concurrence divided by the number of commissioners (calculated only for post-1965 period for FCC); Constraint ceiling calculates probability of presidential and commissioner partisanship independence (outlined in Appendix A), pooling all appointments. Last five columns are presented for pooled 1983–2006 period for FCC. OSHRC data provides bounds, because of lack of data on partisan affiliation for two commissioners. For FEC change, see Buckley v. Valeo, 424 U.S. 1 (1976). Other sources: 47 U.S.C. § 154 (FCC); 15 U.S.C. § 78d (SEC); 15 U.S.C. § 41 (FTC); 42 U.S.C. § 7171 (FERC); 2 U.S.C. § 437c (FEC); Atomic Energy Act of 1954 § 22, Public Law 83-703, 68 Stat. 919 (AEC); 42 U.S.C. § 5841 (NRC); 29 U.S.C. § 153 (NLRB). “?” indicates data unavailable, and “NA” indicates data not applicable.
party commissioners divided by the total number of commissioners who have served, is close to the statutory minimum (partisan requirement) divided by the total number of commissioners. The numbers are close in each instance: the FTC has a minimum requirement of two out of five non-majority party commissioners and the observed cross-party rate is 0.34; the FEC has a minimum requirement of three of six and the observed cross-party rate is 0.56. (Given that the cross-party rate is a proportion of commissioners, we do not expect these numbers to be identical.)

Of course the raw cross-party rate is merely a measure of whether presidents, who typically run up against statutory requirements, in fact obey the law. To assess the effect of partisan requirements, we might turn to three major agencies without partisan requirements: the NLRB (enacted around the same time as the FCC), the Occupational Health & Safety Review Commission (OSHRC), and the Atomic Energy Commission (AEC), which was reorganized into the Nuclear Regulatory Commission (NRC) in 1974, at which point partisan requirements were added. The latter could be seen as a natural experiment, and it suggests that there was no effect on the raw number of cross-party commissioners: the cross-party rate was 0.44 both for the AEC and the NRC. But of course the 1974 reorganization also dramatically changed other key aspects of the NRC, and partisanship may play a different role given AEC/NRC defense-related functions. Indeed Truman appointed only Republicans to the AEC, so the comparison may not provide any information about the effect of partisan requirements on the FCC. If the NLRB is seen as a control, having been enacted at roughly the same time as the FCC, we might infer that there is some bite to partisan requirements: separating periods based on total seats, the NLRB cross-party rate is 0.2 from 1935–47 and 0.29 from 1947–2006, compared to 0.33 from 1934–83 and 0.37 from 1983–2006 for the FCC. At the very least, partisan requirements appeared to have had an effect in the immediate post New Deal period: Roosevelt, for example, appointed no Republicans to the NLRB. Pooling across all years and using appointments as the unit of analysis, a one-tailed Fisher’s exact test yields a p-value of 0.09 that cross-party appointment rates are higher for the FCC. But conditioning on the total number of seats by comparing only post-1986 FCC and post-1947 NLRB appointments yields no statistically significant difference. This could be driven by small sample size or the fact that partisan requirements affected commission composition only in the immediate New Deal period. The OSHRC has a cross-party rate of 0.14–0.23 (due to uncertainty in the partisan affiliation of two commissioners), also suggesting that partisan requirements increase cross-party appointments (p-value bounds are (0.02, 0.14)).

These comparisons across commissions are of course hard to take as serious estimates of the effect of partisan requirements on cross-party appointments. First, due to low variation and small sample size of

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5 Even if requirements are not binding, the cross-party rate could still be close to the statutory minimum.
data across commissions, it would be premature to draw any strong inferences about the effects of partisan requirements using commissions as the units of analysis. Second, if the Congress is optimizing in any sense of the word, it would enact partisan requirements precisely for those commissions in which it anticipates having less oversight. The NLRB, for example, is a heavily politicized agency, and perhaps Congress knew that it would exercise continual oversight, thereby eviscerating the need for more formal partisan requirements. Because of more difficulties in overseeing the work of the FCC, FTC, and SEC, perhaps the Congress looked to partisan requirements as a formal check. Third, it may also be the case that partisanship is an uninformative signal of performance for some policy areas, leading to the rejection of partisan requirements, and an informative signal for others, where partisan requirements are adopted. Lastly, even if partisan requirements don’t affect the raw number of cross-party commissioners, they may still affect the type of commissioners. For these reasons, the analysis of the FCC proceeds below without drawing strong inferences from the cross-commission data.

Another measure of constraint of partisan requirements is how many commissioners historically would not have been appointed but for the partisan requirement (“but-for” commissioners). Although there is some uncertainty over when appointments are in fact constrained by partisan requirements, one natural definition is to code cross-party appointees as but-for commissioners if appointed when the number of active commissioners from the president’s party is at a statutory maximum. For example, of 15 cross-party commissioners appointed to the FCC, 11 are clearly constrained under these criteria. Even the remaining 4 (Quello, Copps, Fogarty, and Hyde) are plausibly but-for commissioners. If it were sufficiently clear that some subset of cross-party commissioners is not induced by partisan requirements, we might have additional leverage to investigate the differences across constrained and unconstrained cross-party commissioners. With this conservative measure in hand, Table presents the but-for rate across commissions. With the exception of the NRC, on which many independents served, they are all quite high (ranging from 0.73 to 1.00), suggesting that partisan requirements may have some teeth.

In the end, however, inferences across commissions remain fragile. Perhaps the most telling sign is that

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6 Even if partisan requirements affect neither the number nor the type of commissioners, the analysis below still offers the first application of ideal point methods to the investigation into policy preferences at the FCC.
7 Other coding criteria exist, but were not particularly reliable. Using the nomination, in lieu of confirmation, presents the problem that some commissioners are nominated before or at the same time that a party-line appointee is nominated. Collecting information about informal announcements of a commissioner’s intent to retire is difficult and likely fails to come close to the information set of the White House personnel office.
8 Quello was nominated by Nixon to fill the one empty Democratic seat, but because his appointment process dragged on, with the longest Senate hearing to date on any FCC nominee, he was confirmed at a time when only 3 of 7 FCC commissioners were still actively serving on the commission. Copps (a Democrat) and Abernathy and Martin (both Republicans) are confirmed on the same date, but Copps and Abernathy started their service one week earlier. Fogarty, a Democrat and Ford appointee, was nominated for a Democratic seat and was confirmed the same day as White (a Republican), but because he started serving 6 days before White he is not measured as a but-for commissioner. Hyde is less clear due to simultaneous departures from the FCC.
presidents virtually never appoint (a) fewer party-line commissioners than possible (with the exception of a handful of independents), or (b) more cross-party appointments than statutorily required to commissions with partisan requirements. Randomization inference (outlined in Appendix A) strongly rejects the hypothesis that commissioner partisanship is independent of presidential partisanship when partisan requirements don’t constrain an FCC appointment ($p$-value $\approx 0.00$). The last column in Table 1 finds the same across other agencies, save for the FEC, where there are virtually no unconstrained appointments. Presidents don’t do more or less than required by statute. Partisan requirements may very well have played a historical role in engendering this norm, and the question then becomes whether cross-party appointees in fact have any substantive impact on decided cases.

3.2 The FCC

For the remainder of the paper, I focus on the FCC for three chief reasons. First, as the penultimate column of Table 1 reveals, the FCC has the highest dissent rate of all commissions, making it ideal to distinguish between commissioners from voting records. That said, the low dissent rate on other commissions suggests that a considerable dimension of agency output may not be captured by focusing on votes alone. For example, the product of agency deliberation may be subject only to a formal vote at the tail-end of a long deliberation process, during which policy preferences of each of the commissioners are incorporated into the proposed rule or order. The vote data may thereby underestimate dissensus on commissions. Although each of these dimensions of agency decisionmaking are worth studying independently, I focus here on votes on substantive rulemakings and adjudications, which is surely at least one crucial aspect of agency behavior.

Second, the FCC has a partisan requirement. While there is no direct “control” commission, the strategy of this research design is to test observable within-agency implications of extant theories. As discussed in Subsection 3.1, there may be no genuine control commission that doesn’t also differ dramatically in a large variety of other criteria (observed and unobserved). From one perspective, all cross-party commissioners are affected by the partisan requirements, but what we desire to test is whether cross-party commissioners differ from party-line commissioners: e.g., do Democrats appointed by Republicans differ from Democrats appointed by Democrats? This design thereby capitalizes on variation of who appoints whom.

Third, the FCC is of inherent substantive interest to the understanding of independent regulatory agencies. As a New Deal agency, the FCC has experienced several waves of technological revolution, from microwave challenges to AT&T’s long-distance monopoly to IP-enabled services threatening the infras-

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9 The $p$-value for the FERC is 0.10, but because the FERC only existed for 5 presidential administrations (3 Republican), the randomization distribution consists of only 10 \(= \binom{5}{3} \) test statistics, such that the power of the test is low. The Federal Power Commission, which was reorganized into the FERC in 1977, also had partisan requirements.
structure of conventional common carrier regulations. The FCC has also not been insulated from political struggle, particularly in the areas of content regulation, antidiscrimination, and the fairness doctrine. To the degree that the FCC differs from other agencies, inferences here are “in-sample” (i.e., limited to the FCC).

Turning to the FCC data, Figure 1 presents FCC commissioner appointments, with years on the y-axis and seats on the x-axis. Commissioner partisan affiliation is denoted by the solid color of the bars (blue for Democrat, red for Republican, and purple for other), the length of which denote the length of service. Bars outlined in black represent but-for commissioners; hence, not all cross-party commissioners are outlined in black. The larger black boxes (filled with white) represent the statutory term length of the seat, and the vertical lines indicate who served as chair. Background colors denote the partisanship of the appointing president, with white lines separating unique administrations. Stewart, for example, depicted in the third column on the top, was a Democratic commissioner appointed by Roosevelt, serving from 1934–37 in a three-year seat (seats were initially 1-7 years in length to stagger appointments). Payne, Case, and Brown, were Republican but-for appointees and Sykes served as chair for the first (FDR-appointed) commission. When a colored bar extends beyond the term box, the commissioner is a “holdover” serving until another nominee is confirmed, such as Quello’s service during the Clinton administration prior to Kennard’s confirmation. As can be seen from Figure 1, the FCC organic act requires terms to be staggered.

The total number of FCC commissioners was changed in 1983 from 7 to 5, and in 1986 the term length was reduced from 7 to 5 years. As the seventh column of Table 1 makes clear, commissioners on average serve only 3-4 years, thereby providing each president a handful of appointments. No third-party or independent appointees serve after the Eisenhower administration.

To test for partisan requirement effects, I compiled a new dataset of all votes published in FCC rulemakings and adjudications available in Westlaw. Appendix B documents details on case selection. Although the dataset includes both orders in adjudications and rules in rulemakings, for simplicity I will refer to all actions as “cases.” From 1965–2006, a total of 46 commissioners served in the FCC, deciding a total of 17,879 cases, casting a total of 94,293 votes. The sample for this study are 5,417 non-unanimous cases, in which 30,185 votes were cast. Figure 2 plots basic summary statistics of the sample. The first panel presents the total yearly count of cases, which spikes in the mid-1970s, reflecting the FCC’s turn in 1965 toward formal rate investigations and increased filings by public interest groups. The second panel plots the proportion of unanimously-decided cases, which peaks in the late 1980s. The last three panels plot the proportion of cases with a concurrence, partial dissent, or full dissent. One key measurement challenge is that despite the fact that total number of cases in the dataset is large, ideal points are conventionally

---

10 The appointment dates prior to 2001 were cross-validated with David C. Nixon’s valuable Independent Regulatory Commissioner Database (available at http://www2.gsu.edu/ wwwirc/).
Figure 1: Summary of FCC Commissioners Appointments, 1934–2006. Each colored bar indicates the term served by a commissioner, denoted by red for Republicans, blue for Democrats, and purple for independents. The border is black if the commissioner was a but-for commissioner, subject to partisan requirements on the day of confirmation. The larger box indicates the statutory term of the seat. Background colors denote partisanship of the appointing president, with each administration separated by white horizontal lines, and narrow lines indicate the chair of the commission.
identified using only cases with dissents. This forces researchers to discard large amounts of information – i.e., dropping 89% of all FCC cases and 65% of non-unanimous FCC cases. The extension of ideal point methods, discussed below in Section 4, addresses this directly, by incorporating more voting information.

Each commissioner’s propensity to deviate from the majority (as measured by the proportion of cases where a commissioner files a concurrence or dissent) differs. The lowest deviation rate is 1.2% for Sikes; the maximum is 29.5% for Johnson; and the average deviation rate is 7%. The fact that Sikes’s rate is so low could merely mean that he was the median commissioner during his terms, whereas Johnson may have been an outlier. To systematically analyze this data, we require statistical methods to measure commissioner preferences.

4 A Multilevel Ideal Point Model of Mixed Ordinal Votes

I use a multilevel model for the FCC data: at the case level, ordinal votes cast by each commissioner are fit with an ordered probit model, with the only complication that explanatory parameters are unobserved and treated as missing data; at the commissioner level, a simple linear model with commissioner and presidential partisan affiliation as covariates is used. This model provides direct uncertainty of ideal points, naturally fixes the policy dimension using prior information, directly tests observable implications of theories of partisan requirements, and incorporates more information about heterogeneity in quasi-judicial votes than dichotomous ideal point models. It is a straightforward generalization of Bafumi et al. (2005); Clinton et al. (2004); Treier and Jackman (2003); and Hathaway and Ho (2006).

4.1 Case Level

For each non-unanimous case indexed by $i = \{1, ..., N\}$, a subset of commissioners indexed by $c = \{1, ..., C\}$ participates. Active commissioners do not necessarily participate in all cases, so the total number of votes
are not necessarily a multiple of $N$. Votes by commissioner $c$ on case $i$ are denoted by:

$$Y_{ic} = \begin{cases} 
1 & \text{if } c \text{ voted for the majority} \\
2 & \text{if } c \text{ concurred} \\
3 & \text{if } c \text{ partially dissented} \\
4 & \text{if } c \text{ dissented in toto}
\end{cases}$$

The observation mechanism is a simple ordered probit:

$$P(Y_{ic} = k) = \Phi(\gamma_{k+1}^{i} - \mu_{ic}) - \Phi(\gamma_{k}^{i} - \mu_{ic}),$$

(1)

where $\gamma_{k}^{i}$ denotes the $k$th cutpoint for case $i$, $k \in \{1, 2, 3, 4\}$, and $\Phi()$ is the standard normal CDF. As is conventional, for identification $\gamma_1^i \equiv -\infty$, $\gamma_2^i \equiv 0$, and $\gamma_4^i \equiv \infty$. To illustrate, if $\gamma_3^i = 2$ and $\mu_{ic} = 0.1$, the probability of a concurrence is $\Phi(2 - 0.1) - \Phi(0 - 0.1) \approx 0.51$. The probability of disagreement increases generally in $\mu_{ic}$: for example, if $\mu_{ic}$ increases to 1, the probability of a concurrence becomes 0.62. (An increase in $\mu_{ic}$ might also decrease the probability of a concurrence, while increasing disagreement by dissent.) The systematic component $\mu_{ic}$ is a linear transformation of $\phi_c$:

$$\mu_{ic} = \Lambda_{i1} + \Lambda_{i2}\phi_c.$$ 

(2)

The “ideal points” $\phi_c$ can be interpreted as policy preferences in unidimensional policy space (typically, liberal-conservative – see Subsection 4.2). $\Lambda_{i1}$ models the likelihood that case $i$ will generate disagreement between the commissioners. It corresponds to an “item-difficulty” parameter in educational testing, accounting for how many test-takers answer a question incorrectly. $\Lambda_{i2}$ models to what degree disagreement in case $i$ is driven by latent ideal points $\phi$ of the commissioners. It corresponds to an “item discrimination” parameter in educational testing, where it indicates how much a test question discriminates by test taker ability. For example, if $\Lambda_{i2}$ is positive commissioners with higher values of $\phi$ will be more likely to disagree. If $\Lambda_{i1}$ is positive but $\Lambda_{i2} = 0$, commissioner disagreement is orthogonal to underlying policy preferences (hence the derivation from a random utility model).

One key strength of this approach is that it does not rely on coder-contingent “directional coding,” where individual coders have to specify whether a particular decision was “liberal” or “conservative.” Such data collection can be error-prone and costly, thereby confining researchers to only a limited sample of cases. The model directly estimates the latent direction that each vote signifies from the data.

One remaining complication is that we don’t observe all four types of votes in all cases – we have a mixture of di-, tri-, and quadrichotomous votes. All elements of $\gamma^i$ are not identified (or identified entirely by the priors) when $i$’s number of unique votes is less than four. When $i$’s votes are dichotomous, no cutpoints need be estimated and the observation process is equivalent to a standard dichotomous probit.
model where:

\[
Y_{ic} \in \begin{cases} 
\{1, 2\} \forall c, \text{then } \gamma_3^i \equiv \gamma_4^i \equiv \infty \\
\{1, 3\} \forall c, \text{then } \gamma_3^i \equiv 0 \text{ and } \gamma_4^i \equiv \infty \\
\{1, 4\} \forall c, \text{then } \gamma_3^i \equiv \gamma_4^i \equiv 0.
\end{cases}
\]

When \(i\) is trichotomous, \(\gamma_3^i\) is estimated, and the observation process is equivalent to a trichotomous ordered probit, where:

\[
Y_{ic} \in \begin{cases} 
\{1, 2, 3\} \forall c, \text{then } \gamma_4^i \equiv \infty \\
\{1, 2, 4\} \forall c, \text{then } \gamma_3^i \equiv \gamma_4^i \\
\{1, 3, 4\} \forall c, \text{then } \gamma_3^i \equiv 0.
\end{cases}
\]

When four unique types of votes are observed, \(0 < \gamma_3^i < \gamma_4^i\). The model can be interpreted as an Bayesian factor analysis model with mixed ordinal indicators (Quinn, 2004). This clarifies that the ordered probit generalizes conventional ideal point methods designed for dichotomous votes (aye/nay). While a dichotomous approach may be appropriate for legislative roll calls, it ignores significant information in the study of judicial (or quasi-judicial) decisionmaking.

I use diffuse priors for all remaining case level parameters:

\[
\Lambda_i \sim N(0, 4) \\
\gamma_3^{i, \#(Y_i) \in \{3, 4\}} \sim LN(1, 1) \\
\gamma_4^{i, \#(Y_i) = 4} - \gamma_3^i \sim LN(1, 1),
\]

where \(LN\) indicates the log-normal distribution, \(\#(Y_i)\) indicates the number of unique votes in \(Y_i\), and the last line differences cutpoints to ensure that they monotonically increase.

### 4.2 Commissioner Level

So far ideal points are not identified, as the latent ideological dimension is not scale or rotation-invariant. Each of the parameters in Equation 1 can be multiplied by \(-5\), resulting in equivalent fit of the model. The dimension still lacks substantive meaning. To address this identification problem and to test directly propositions about the effects of partisan requirements, I incorporate information at the commissioner level (see Bafumi et al., 2005). Specifically, we observe commissioner \(c\)’s partisan affiliation \((A_c)\) which equals 1 if Republican and -1 if Democratic. (No independents are appointed from 1965–2006.) In addition, we observe the party of \(c\)’s initial appointing president \((P_c)\) which equals 1 if Republican and -1 if Democratic. Subsection 6.3 investigates sensitivity to the use of this measure. Commissioner ideal points are modeled

\[11\] Note that the model conditions on the types of votes in each case – for example, it does not account for uncertainty in the issuance of a partial or full dissent when only one type is observed.
Table 2: Interpretation of Commissioner-Level Parameters. This Table spells out observable implications from extant theories. The last two rows can be interpreted as versions of presidential-legislative bargaining. Strictly speaking there are four other possible combinations omitted from the Table: (a) $\delta_2 < 0$ and $\delta_3 = 0$; (b) $\delta_2 = 0$ and $\delta_3 < 0$; (c) $\delta_2 < 0$ and $\delta_3 < 0$; and (d) $\delta_2 < 0$ and $\delta_3 > 0$. While these might be termed “legislative ignorance,” “presidential ignorance,” “ignorant bargaining,” and “ignorant cross-party extremes,” respectively, they are very likely to be artifacts of a reversal of the dimensional axis. Validating the dimension with a substantive understanding of the cases and commissioners is crucial for this data-drive identification approach.

Table 2

<table>
<thead>
<tr>
<th>Theoretical Approach</th>
<th>Commissioner Party ($\delta_2$)</th>
<th>Presidential Party ($\delta_3$)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Pure nonpartisanship</td>
<td>Zero</td>
<td>Zero</td>
</tr>
<tr>
<td>2. Presidential dominance</td>
<td>Zero</td>
<td>Positive</td>
</tr>
<tr>
<td>3. Legislative dominance</td>
<td>Positive</td>
<td>Zero</td>
</tr>
<tr>
<td>4. Presidential-legislative bargaining</td>
<td>Positive</td>
<td>Positive</td>
</tr>
<tr>
<td>5. Cross-party extremists</td>
<td>Positive</td>
<td>Negative</td>
</tr>
</tbody>
</table>

Table 2 summarizes observable implications and associated theories. Row 1 indicates a purely nonpartisan commission, where partisanship does not explain voting behavior at all. Row 2 presents observable implications for presidential dominance: if partisan requirements don’t constrain the president, commissioner’s partisan affiliation should have no predictive power over voting behavior holding constant the party of the appointing president. Legislative dominance (row 3) implies that presidential party has no explanatory power holding constant commissioner affiliation. Presidential-legislative bargaining models (row 4) would suggest that both commissioner and presidential partisan affiliation are positive: for exam-

as:

$$\phi_c \sim N(\delta_1 + \delta_2 A_c + \delta_3 P_c, 1)$$

$$\delta_{m,m \in \{1,2\}} \sim N(0, 25)$$

$$\delta_3 \sim N_+(0, 25),$$

where $N_+$ indicates a truncated normal. This constrains $\delta_3$ to be positive, fixing the latent dimension so that higher positive values of $\phi$ indicate more conservative ideal points (associated with Republican appointers). The dimension can thereby be interpreted as “right” and “left” in terms of regulatory preferences. We can see whether this constraint is reasonable by plotting its posterior distribution and seeing whether appreciable posterior mass is close to 0. If it is not, we can constrain $\delta_2$ to be positive, conducting the same check (Bafumi et al., 2005). In the FCC data, the posterior mass of $\delta_2$ is so far from the origin that it provides a reasonable way to identify the model.

12 In Table 2 “positive” indicates that the posterior mass of a coefficient is positive in the sense that there is no posterior mass that runs up against the origin, even if constrained to be positive for identification.
ple, Democratic commissioners vote more liberally than Republican commissioners, even holding constant
the appointing president’s party; and cross-party commissioners are moderates. The last row depicts the
seemingly-counterintuitive possibility that plays prominently in the empirical results, and has not been con-
templated in the literature\[13\] if $\delta_2 > 0$ and $\delta_3 < 0$, cross-party appointees are extremists, exhibiting voting
patterns even more extreme than party-line appointees.

4.3 Estimation

With the probability model identified, we can now specify the full posterior as:

$$P(\Lambda, \phi, \delta, \gamma | Y, X) \propto P(Y | \Lambda, \gamma, \phi)P(\phi | X, \delta)P(\delta, \Lambda, \gamma),$$ \hspace{1cm} (4)

where $X$ contains $A$ and $P$. The total number of parameters is a function of the number of cases ($N$),
the number of unique votes on each case ($\sum N_i \#(Y_i)$), the number of commissioners ($C$), and the number
of commissioner covariates ($J$), namely $C + \sum_N \#(Y_i) + J$. With 46 commissioners voting on 5,417 cases
from 1965-2006, fitting the model involves estimating 11,986 parameters. Bayesian methods using Markov
Chain Monte Carlo simulation are ideally suited for this kind of a high dimensional model. I use WinBUGS
(Spiegelhalter et al., 2004) and R2WinBUGS (Sturtz et al., 2005) to draw 150,000 samples of parameters
from Equation 4. Standard diagnostics ($\hat{R}$, ACF plots, etc.) indicate convergence.

5 Results

This section presents results from ideal point estimates, validating the score with conventional qualitative
accounts of commissioner ideology and cases. Posterior predictive checks suggest good model fit. Cross-
party commissioners appear to be more extreme than party-line appointees after 1980. This seemingly-
counterintuitive result is consistent with further (independently-derived) tests of whether Senatorial defer-
ence changed in the 1980s. Lastly, I calculate substantive policy effects by simulating the empirical and
counterfactual median commissioner over time. Partisan requirements appear to smoothen policy outcomes
temporally.

5.1 Ideal Point Estimates

Figure[3] presents bivariate slices of the posterior for pairs of commissioners serving in one natural commis-
sion (a combination of commissioners in active service at the same time) in 1977. Commissioners are sorted

\[13\] Compare Nagel and Lubin (1964) (finding, but discounting, that Democratic commissioners appointed by Republicans may
be more liberal than Democratic commissioners appointed by Democrats).
by median ideal points, such that the left-top presents the most liberal commissioner and the right-bottom presents the most conservative commissioner. Democrats are denoted in blue and Republicans in red. Fogarty, for example, is a Democratic Ford appointee, who is the most liberal commissioner in this natural commission. Panels below the diagonal presents scatter plots that hold fixed the axes ranges for all pairs. Moving down the rows from the top-left cell, for example, the marginal distribution on the \( y \)-axis remains fixed for Fogarty, but the distribution for each of the matched commissioners moves to the right. Because these scatter plots never intersect the 45-degree line, the posterior probability that Fogarty is more liberal than even the second-most-liberal Hooks is very high (each pairwise single-tailed posterior \( p \)-value \( \approx 0 \)).

Panels above the diagonal present contour plots, permitting the axes ranges to float. The larger the contour plot, the more closely related two commissioners are in terms of ideology. From this we can easily see that Quello was the most conservative Democrat, voting frequently with the four Republicans. The posterior sample easily permits us to calculate other quantities of interest: for example, there is a 10% probability that Quello is the median commissioner (White and Wiley, the Chairman, have 59% and 30% probabilities of being the median commissioner). While Quello may be an example of a successful effort by Nixon to appoint a conservative Democrat, anecdotes do not make for generalizable patterns. After all, the Senate can reject such nominees: at the same time when Quello’s nomination was pending, Nixon also nominated Holcomb, a Texas Democrat who had previously declared in a letter to Nixon his “total commitment to President Nixon for re-election,” but Holcomb’s nomination died in the Senate.

To determine whether cross-party commissioners, like Quello, come in sheep’s clothing, Figure 4 summarizes ideal point results for all 46 commissioners serving in our observation period. The posterior medians are denoted by the dots (the colors of which represent the party of the appointing president), and the intervals denote 95% posterior intervals (the colors of which represent the party of the commissioner). As illustrated by Figure 3, marginal posterior distributions are roughly symmetric. The uncertainty intervals are quite narrow – we have sufficient information in the cases to estimate ideal points with relative precision. The intervals also take into account estimation uncertainty. For example, Sharp served only 9 months, as his seat was eliminated in the 1983 reduction to five seats, and his interval, accordingly, is very wide. From Figure 4 it becomes clear that most partisans are in fact genuine: Democrats and Republicans separate quite cleanly.

These ideal points make substantive sense. We can validate this by examining cases and commissioners. In terms of cases, consider the FCC’s highly contested and politicized 2002 biennial review of broadcast ownership concentration rules, specifying how many stations and how much of the market could be owned

by a single entity.\textsuperscript{15} Upon the FCC’s notice, some three quarters of a million citizens and organizations, from churches to labor unions to the NRA to NOW, and 150 Congressmembers – Republican and Democrat – flooded the FCC with public comments. The commission split 3-2 along straight partisan lines of the commissioners (but not their appointing presidents) to relax concentration restrictions. Powell, Abernathy, and Martin voted to approve the measure, each issuing separate concurrences, while Copps and Adelstein, both W. Bush appointees, dissented. Adelstein lambasted the majority, charging that “[t]his is the most sweeping and destructive rollback of consumer protection rules in the history of American broadcasting. . . . [it] shatters most of the last vestiges of the consumer protections that weren’t eliminated in the 1980’s . . . [and] leav[es] the FCC a toothless tiger.”\textsuperscript{15} Copps similarly charged: “[T]oday the FCC empowers America’s new Media Elite with unacceptable levels of influence over the media on which our society and our democracy so heavily depend.”\textsuperscript{17} The model incorporates this split directly. The item discrimination

\textsuperscript{17} Statement of Commissioner Michael J. Copps (June 2, 2003) (dissenting from Biennial Review).
Figure 4: Ideal point estimates for all commissioners. The x-axis represents the policy dimension, which can be interpreted as going from liberal (negative) to conservative (positive). Red (blue) intervals and names denote Republican (Democratic) affiliation of the commissioner. Red (blue) median dots indicate Republican (Democratic) appointing president. Thicker intervals denote but-for commissioners, and bolded names indicate cross-party commissioners.
parameter $\Lambda_2$ for the *Biennial Review* is strongly negative: the posterior probability that it is negative is 100%, indicating that the *Biennial Review* was strongly polarizing, and that votes are strongly associated with the latent policy dimension. Copps and Adelstein, two of the five most liberal commissioners, were much more likely to disagree with the majority.

In *Alabama Educational Television Commission*\(^{18}\) the commission, in a 4-2 order, denied license renewal to certain Alabama TV stations for racially discriminatory practices. Two commissioners, Lee and Reid, dissented. All of the members of the majority – Hooks, Quello, Robinson, and Washburn – are estimated to be substantially more liberal than Reid and almost all are estimated to be more liberal than Lee (though Washburn’s ideal point interval overlaps substantially with Lee’s).

Another illustrative case is *Vonage*, in which the FCC found state telecommunications provisions for IP-enabled phone services to be preempted by federal law.\(^{19}\) Martin, who is estimated to be slightly more conservative than all others on the panel, was in the majority; Powell and Abernathy, who are both more liberal than Martin, concurred; and Copps and Adelstein, the most liberal panel members, also concurred. The posterior probability that the item discrimination parameter is less than 0 is roughly 98%, so we estimate that concurrences were driven by liberal concerns (e.g., about federal preemption with deregulatory effects).

News reports and popular perceptions of the commissioners are also resoundingly consistent with our estimates. Furchtgott-Roth, who is estimated as the most conservative commissioner by far, is a conservative free market economist, termed “Dr. Dissent” \(^{20}\) Zarkin and Zarkin, p. 174–75), who wrote an AEI-published book arguing that independent regulatory agencies violate separation of powers.\(^{20}\) In a *National Review* column, he wrote that “[p]ractically all that ails the American telecommunications sector can be cured with stronger property and contract rights, not with industrial policy.”\(^{21}\) The second-most conservative commissioner, Reid, is described as “slightly to the right of Marie Antoinette.”\(^{22}\) Although imprecisely estimated as quite conservative, McDowell, sponsored by Senator George Allen, is also on the record as stating that “[m]ost industry conflicts can be resolved by market forces, new technologies and innovation.”\(^{23}\)

On the other end of the ideological spectrum, Hundt, estimated as amongst the four most liberal commissioners, supported initiatives like 3-hour mandatory children’s television, a ban on liquor ads, and free airtime to political candidates \(^{24}\) Zarkin and Zarkin, 2006. One commentator noted: “thank goodness . . . Hundt

\(^{18}\) 50 F.C.C. 2d 461 (1975).
\(^{19}\) 19 F.C.C. Rcd. 22,404 (2004).
is gone. He managed to make enemies in nearly every industry that the FCC regulates.\footnote{Jeffrey Krauss, \textit{Here Comes the FCC}, CED Magazine, Jan. 1, 1998.} Copps is described as a liberal Democrat, the “conscience of the FCC,” and prior to Adelstein’s appointment as a “one-man band banging the drum for a national debate” on media concentration rules.\footnote{Catherine Yang, \textit{The FCC’s Loner Is No Longer So Lonely}, Bush Week, at 78, March 24, 2003.} Adelstein is described as the “FCC anti-deregulation cheerleader.”\footnote{Bill McConnell, \textit{The FCC Anti-Deregulation Cheerleader}, Broadcasting & Cable, Aug. 25, 2003.} Lastly, Tristani is commonly seen as “pro-consumer”\footnote{See supra note \ref{note24}.} strongly supportive of universal service, and opposed to media consolidation.

While the ideal points perform remarkably well, there are slight surprises in comparison to popular accounts. One might wonder, for example, why Johnson – who wrote a blistering 85-page dissent to a proposed merger between ABC and ITT in 1966, is on public record with caustic remarks against industry groups, and was described by one industry magazine as “a curmudgeon, a burr, . . . a phony” and “hated”\footnote{Leonard Zeidenberg, \textit{Seven Years and Five Months: A Look Back at the Tenure of Nick Johnson}, Broadcasting, at 20, Dec. 10, 1973.} – is not estimated to be more liberal. Other surprises may be Dennis, a Reagan appointee who initially supported repeal of the fairness doctrine, Duggan, who is described as relatively moderate, Jones, who favored deregulation unless there was a clear rule, and Patrick, who is conventionally described as an ardent free-market supporter.

But these seemingly surprising estimates demonstrate the virtue of statistical methods in making transparent the measurement process: while it is true, for example, that Johnson was outspoken and flamboyant – the only commissioner to ever appear on the cover of \textit{Rolling Stone} – as a matter of voting he was also seen as an ally with Cox, who is not similarly covered in press accounts. Similarly, Duggan, who would later become the head of PBS, supported universal service and the fining of TV stations for violating advertising limitations set by the Children’s Television Act (\textit{Zarkin and Zarkin} 2006, p. 169). While Tristani initially opposed efforts to strengthen equal opportunities for women and racial minorities she soon changed her mind during her tenure on the commission. Lastly, Dennis’s voting record belies the initial descriptions of her conservatism\footnote{See, e.g., \textit{FCC’s Patricia Diaz Dennis: Following Her Own Drummer}, Broadcasting, at 87, March 9, 1987; \textit{Life in the Slow Lane at the FCC}, Broadcasting, at 75, Feb. 22, 1988; David Lauter & Cathleen Decker, \textit{Bush Health Post Choice under Fire: Anti-Abortionists Angry; Selection May Come Today}, Los Angeles Times, Dec. 21, 1988; Weisenberger, supra note \ref{note22}.} Press accounts, although they have proven fruitful in exogenously coding preferences in some contexts (\textit{Segal and Cover} 1989), still have weaknesses of comparability across time and commissioners, perceptual bias, and qualitative inconsistencies. Ideal point methods provide an alternative, transparent outcome-based measure of policy preferences that circumvents many of these challenges.

On the other hand, these surprises may also suggest ways to refine the estimation method. For example, one criticism of ideal point methods may be that votes are weighted equally. One might argue that Johnson’s
Figure 5: Posterior predictive check of model. The first panel presents the actual data from 1978-1981 (with cases and commissioners sorted from conservative to liberal according to posterior median $\Lambda_2$ and $\phi_c$), and the other panels present simulated datasets.

85-page dissent should be weighted more heavily to account for (a) the intensity of preferences, and/or (b) the importance of an issue. But of course all of the cases from which ideal points are identified are non-unanimous: someone disagreed with respect to some issue. The item discrimination parameter does not in fact weight ideology equally in cases. Nonetheless, incorporating more information along these lines may prove fruitful.

5.2 Posterior Predictive Checks

Another way to validate the model is to check model fit. Figure 5 plots the observed data on the left panel and three datasets generated from the model in the right panels. In each panel, the $x$-axis represents commissioners sorted from liberal (left) to conservative (right), and the $y$-axis presents cases sorted by conservative (top) to liberal (bottom) outcome. The color of each cell in the data matrix represents how one commissioner voted in the case. The majority of the matrix consists of white cells, because only a handful of the 46 commissioners serve on any given case. Although there is slightly more noise in the replication datasets due to mispredictions, the simulated datasets look similar to the observed dataset. The model classifies roughly 71% of all cases correctly, so model fit appears good.

5.3 Cross-Party Extremism and Senatorial Deference after 1980

Figure 4 demonstrates that we can learn a considerable amount about commissioners from voting records: there is clear separation by partisan affiliation of the commissioner. Cross-party appointees don’t don sheep’s clothing. One striking – and seemingly paradoxical – finding is that cross-party appointees are at the outer extremes of the policy dimension: Copps, Dennis, Duggan, and Adelstein are Republican-appointed
Democrats who are four of the five most liberal commissioners. And Furchtgott-Roth, a Democratic-appointed Republican is by far the most conservative commissioner. To summarize this trend, Figure 6 plots the marginal posterior distributions of $\delta_2$ in the left panel and $\delta_3$ in the right panel. We can see that the party of the commissioner is strongly positive, but that almost all of the posterior mass of $\delta_3$ is negative. Indeed, the posterior probability that $\delta_3 < 0$ is 99%.

As described in Subsection 4.2, fixing the dimension required some fine-tuning. One might ask why the opposite signs of $\delta_2$ and $\delta_3$ don’t mean that the policy dimension is reversed: but the qualitative checks in Subsection 5.1 demonstrate that the dimension identified by $\delta_2$ clearly coheres with the interpretation that commissioners with higher values of $\phi$ are more conservative. Presidential affiliation is such a poor predictor of ideal points that without controlling for commissioner affiliation, its marginal posterior runs into the origin. This can be seen by the asymmetry in the intervals (and histogram plotted below interval) in row 1 of Figure 7. As row 2 shows, using only commissioner partisan affiliation in Equation 3 leads to a marginal posterior far from the origin. (Other results of Figure 7 are discussed below and in Section 6.) Figure 8 plots ideal points sorted by president, and ordered by magnitude within each presidency. For each presidential set, save for Carter’s, Republicans are more conservative than Democrats, as can be clearly seen in the right panel, which limits the y-axis range. This provides resounding evidence against presidential dominance: partisanship requirements constrain. Figure 8 also highlights cross-party extremism. George W. Bush appointed Copps, who is even more liberal than the most liberal Democrat appointed by Clinton. And, in turn, Clinton appointed Furchtgott-Roth, “by far the most conservative commissioner on the Federal

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30 Carter’s results are consistent with the common account that his selection process was decentralized and depoliticized, without much consultation of party officials. (Bonafede 1987a, p. 44-47; Mackenzie 1981, p. 62-69, 196-97).
Figure 7: Robustness of Results. This figure plots posterior 95% and 80% intervals and medians of the parameters on the partisan affiliation of commissioners and appointing presidents in the left and right panels, respectively. Row 1 (from the top) shows that even when omitting commissioner partisan affiliation, the posterior mass of presidential parameter (which is constrained to be positive to identify the ideological dimension) runs up against the origin (see histogram). Row 2 shows that controlling for commissioner affiliation alone leads to strong and robust effects. Rows 3-5 dichotomize voting outcomes to assess sensitivity to ordering assumptions. Row 6 relaxes stationarity assumptions for the longest-serving commissioners (Quello & Lee) in the observation period. Row 7 uses the most recent appointing president, rather than initial appointing president as $P_c$. Row 8 estimates random presidential intercepts to relax the pooling assumption of presidential party. Rows 9-10 fit the model to pre-1980 and post-1980 cases by adoption date. Row 11 dichotomizes votes (majority vs. non-majority) for the post-1980 period. These results strongly confirm the finding that commissioner partisan affiliation is the primary determinant of voting behavior.

The trend appears to be more recent, and one might wonder whether it stems from sampling variability: given that (a) cross-party commissioners typically are appointed later in the presidential administration, (b) the term length was reduced in 1986, and (c) several commissioners resign with a new administration of the opposite party, we might expect cross-party extremism by sampling variability alone. Using a negative-binomial model for the number of decisions in which commissioners participate marginally suggests cross-party commissioners sit on fewer cases throughout their tenure ($p$-value=0.05). But our estimates naturally take into account sampling variability due to the number of votes, and cross-party commissioners should in fact be less likely to resign when the administration changes hands to their party.

The left panel of Figure 8 suggests that cross-party commissioners are extreme only in the post-1980 period (i.e., after the Carter administration). One might hypothesize that the onset of divided government explains this rupture. But the evidence refutes this. Extreme commissioners are appointed during times of divided government and unified government (e.g., Furchtgott-Roth and Copps, respectively). Interacting $P_c$ with an indicator of whether the White House is controlled by a different party than the (a) Senate

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or the House and (b) the Senate alone yields no evidence for cross-party extremism as driven by divided
government. So did something else in the appointments process change in the 1980s? This has been
suggested by commentators in other contexts. For example, beginning in the late 1970s and early 1980s,
Senatorial deference to presidential nominees appeared to break down at the NLRB (Flynn, 2000), and a
new norm of “batching” two to three nominees in the same confirmation hearing, with a balance of pro-labor
and pro-management nominees, emerged.

To formally test this hypothesis, rows 8-9 in Figure 7 present results by fitting separate models to cases
adopted from 1965–80 and 1981–2006. The results strongly suggest that there was a sharp rupture in
appointments practice beginning in the 1980s: commissioner affiliation matters throughout, but the presi-
dential partisan affiliation coefficient becomes strongly negative in the post-1980 period, and statistically
insignificant from 0 in the pre-1980 period. The posterior probability that presidential and commissioner
coefficients differ across these two time periods is roughly 100%. Commissioner partisanship predicts vot-
ing behavior in both periods, but has a drastically higher effect after 1980; and cross-party commissioners
are more extreme only in the post-1980 period (cf. Krasnow et al., 1982, p. 41).

Further evidence bears out the hypothesis that a new Senatorial norm of batching developed for the
FCC just as for the NLRB. First, Furchtgott-Roth was himself proposed to the White House by then-Senate
Majority leader Trent Lott and the strong backing of House Commerce Chairman Bliley, a conservative
Republican, in a self-described “package” deal with the White House Personnel Office for three seats on
the FCC (along with Powell, backed by Senate Commerce Chairman McCain, and Kennard).32 Despite

Figure 8: Ideal Points by Appointing President, ordered by magnitude within Presidential Administration.

rumors of Furchtgott-Roth’s conservative approach, one trade reporter suggested executive deference to partisan-reserved slots: “[i]t would be highly unusual for White House not to follow recommendations on Republican nominations.” Rural Republicans and Democrats challenged the process of how such a package deal was negotiated, and Democrats charged that the White House had “listened to the wishes of two Senate Republicans and ignored the concerns of 15 Senate Democrats, including the entire Democratic Leadership.” But the package deal, which ultimately also included Tristani, was successful – all four were confirmed on Oct. 28-29. Second, Copps previously served as an aide to Hollings, the ranking Democrat on the Senate Commerce Committee, and was also part of a three-commissioner package with Abernathy and Martin. As one commentator noted: “[H]ad Michael Copps not been included it would have been a lot more difficult to get someone through the Senate Commerce Committee.” The batch sailed through, and was confirmed on May 25. In describing the Copps package, some hailed it as an “era of improved relations between lawmakers and telecommunications regulators.” And in the push for the subsequent Democratic seat, House Democratic leaders, contrary to the earlier procedural opposition to the Furchtgott-Roth package, even appealed to the apparent emerging norm: “In prior instances when minority party candidates were selected for FCC posts, deference has been given to the choices of Commerce Committee leaders from the same party on both sides of the Capitol.”

This qualitative evidence is also consistent with the fact that in the late 1970s the Senate conducted exhaustive studies into the appointments process, imposed more stringent financial disclosure and divestiture requirements, began to hold longer hearings, increasingly used strategic holds on nominations for political leverage, and increasingly scrutinized nominees (Graham and Kramer 1976; U.S. Senate Committee on Government Operations 1977; Bonafede 1987b; Mackenzie 1981; Deering, 1987, p. 82-83). Historically, this has been a shift linked with the rejections of Lefever for Assistant Secretary of State for Human Rights and Bork for the Supreme Court. Something appeared to change around the time Reagan assumed office. The apparent explanation of the shift (and the seemingly-paradoxical findings on extreme cross-party appointees) is that the Senate was flexing its advice and consent muscles, resulting in a new norm of confirming nominees in batches by dividing control between parties (Mackenzie, 1997, p. 29-34). And batching produces outliers. This is consistent with several other pieces of evidence (derived independently).

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33 E.g., Furchtgott-Roth apparently described universal service as “corporate welfare.”
34 Lott Recommends Powell and Furchtgott-Roth for FCC, COMM. DAILY, April 11, 1997.
First, the number of confirmations of multiple nominees increased significantly in the 1980s: measuring “batching” by the number of nominees confirmed two days apart, 24% of nominees were batched prior to 1980, compared to 48% after 1980 \( (p\text{-value}=0.02 \text{ from Fisher’s exact test}) \). Second, as can be seen in Figure 1, the holdover and vacancy rate increased dramatically. Prior to 1980 the average number of days a seat was vacant (excluding re-nominations) was 67, compared to 209 post-1980 (Kolmogorov-Smirnov \( p\text{-value}=0.00 \)) (see also [2001]). And 21% of terms resulted in holdovers prior to 1980, compared to 43% after 1980 \( (p\text{-value}=0.02 \text{ from Fisher’s exact test}) \). Lastly, while no commissioners had served as congressional staffers prior to 1980, 14% of commissioners after 1980 had \( (p\text{-value}=0.02 \text{ from Fisher’s exact test}) \). In short, considerable evidence points to a sharp rupture of the appointments process in the 1980s.

But why? Batching may have been an optimal strategy to reduce political transaction costs (Epstein and O’Halloran, 1999) in light of costly nomination fights of the 1970s. As Appendix C illustrates, the Senate and the Commerce committee became increasingly polarized along party lines during this period. By circumventing the need to aggregate preferences for every nomination, the division of control over a group of nominees can reduce the cost of nomination (both from the President, and minority and majority of the Senate), while increasing variance across nominees. And since the executive – typically more majoritarian in outlook (see, e.g., [1989] p. 279-80) – asserts more control for party-line appointees, batching would explain why cross-party appointees are even more extreme than party-line appointees.

In sum, the evidence strongly suggests that presidents, in facing a cross-party seat, are at the very least limited in their choices by partisan requirements, if not entirely deferential for cross-party appointments to senior congressional leaders post-1980.

5.4 Substantive Policy Effects

Even if presidents are constrained, does this matter substantively? After all, most presidents at least have the opportunity to appoint the median commissioner. At the outset, several aspects are worth noting. First, a large fraction of cases is not in fact decided by a full commission. Indeed, due in part to recusals and vacancies, roughly 63% of all cases are decided by fewer than 7 commissioners before 1983, and roughly 44% of all cases are decided by fewer than 5 commissioners after 1983. Second, appointments affect the probability of appointing the median commissioner in the future. Third, the majority does not always vote as a block; there is considerable heterogeneity in ideal points even amongst commissioners in the majority.

To investigate substantive policy effects, the left panel of Figure 9 plots the historical median commiss-

\[^{40}\] After 1980, cross-party appointees are also more than twice as likely to have prior staff experience, although the difference is not statistically significant (due to the small total number of commissioners with staff experience).

\[^{41}\] A similar logic favors delegation to the executive in the trade context ([2005]).
sioner’s ideal point (with posterior draws to account for uncertainty) over time. The units of analysis are natural commissions. There are more changes in the natural commission prior to 1982 due to differences in term lengths and size. One striking pattern is that the median ideal point is a relatively non-smooth step function over time, which is largely induced by the staggered nature of FCC terms and variation in natural commissions. In addition, there appears some degree of lag time until new presidents appoint the median. The text at the bottom of the graph depicts the appointing president of the commissioner who has the highest posterior probability of being the median commissioner for any given date. Several years after Clinton left office, his appointees remained in the median position, and not until two years into the Carter administration did Carter appoint the median. More severe lagged effects persist: due to a number of shifts around 1994 (e.g., between 1/31/1994 and 5/22/1994 only 3 active commissioners served on the FCC), Nixon’s influence lingered in the hands of Quello, who served 23 years on the FCC.

The right panel simulates counterfactual data if instead of making cross-party appointments, Democratic (Republican) presidents appointed party-line commissioners with the same ideology as the most liberal (conservative) party-line commissioner actually appointed by that president. Given the considerable difficulties outlined in Subsection 3.1, this is most likely an upper bound for the causal effect of partisan requirements. Nonetheless, it illustrates the potential substantive effects of partisan requirements. Most importantly, the range of median ideal points is much greater than in the left panel. Partisan requirements, by mandating the appointment of an ideologically-balanced commission moderate policy outcomes over time. The sharp swings in the right panel contrast with the incremental steps in the left panel. Without partisan re-
quirements, given rapid turnover, presidents could gain control relatively quickly, leading to sharper swings in regulatory policy. In contrast to the left panel, for example, George W. Bush almost immediately is able to appoint the median commissioner after 2001. (Carter’s appointees don’t serve as median commissioners because even the most liberal Carter appointees are moderates.) Partisan requirements, then, can be conceived to serve a temporal smoothing function. Yet even with these swings, the median is never as extreme as cross-party appointees.

6 Potential Threats to Validity

In this section I present a series of robustness and sensitivity analyses. The primary finding rejecting the hypothesis that cross-party appointees don sheep’s clothing remains strongly robust.

6.1 Ordering

This far, to capitalize on the heterogeneity of votes, I have assumed that votes on a given case are ordered in ideologically meaningful ways. A concurrence is meaningfully different from a straight vote for majority. Moreover, only one-directional disagreements with the majority are allowed. This is violated if one commissioner concurs on more liberal grounds from the majority and another dissents in a more conservative direction than the majority. The assumption is likely more reasonable than a dichotomous pooling assumption (e.g., that a concurrence provides the same information as a vote for the majority), but inferences may be biased if it is violated.

To investigate sensitivity to the ordering assumption, I conducted several sensitivity analyses. First, I dichotomized votes, and estimated a standard dichotomous ideal point model using each possible cutpoint (e.g., majority vs. non-majority, majority / concurrence vs. partial dissent / full dissent, no full dissent vs. full dissent). This trades off two sources of bias: bias due to false ordering and bias due to discarding cases, likely non-randomly. For example, using only cases where a full dissent is issued discards some 65% of the data. And using only cases with dissents leaves fewer than 350 cases in the dataset for post-1980. The dichotomous results, summarized in rows 3-5 of Figure 7 are largely similar, although the ideal point intervals are wider since we are discarding information. While commissioner partisan affiliation is still highly predictive, the posterior mass for presidential affiliation now includes the origin when pooling pre and post-1980 cases. This gives some reason to pause over the broad conclusion of cross-party extremism, but even then our conclusion would remain that holding constant commissioner partisan affiliation, there is

42 The existence of partisan requirements could of course also affect turnover.
no predictive power of presidential partisan affiliation over votes by FCC commissioners. On the other hand, dichotomization introduces several discrepancies with popular accounts, suggesting that the ordered model performs better: Cox, for example, is estimated to be the third most conservative commissioner, which is inconsistent with most accounts of Cox as a solid liberal. Moreover, row 11 in Figure 7 strongly confirms post-1980s results of cross-party extremism by dichotomizing majority vs. non-majority.

Second, reading a sample of cases with three or more types of votes suggests that the assumption is not violated. To begin, the ordering assumption only matters for roughly 19% of the sample. That said, for some cases it is difficult to assess whether the ordering assumption is violated as commissioners vote without issuing written opinions.

Lastly, some cases of course do not map onto our ideological dimension at all. The model is probabilistic, and explicitly allows for such cases. For example, in *Pacifica*, 5 of the 7 commissioners filed concurrences to the majority’s declaratory judgment against Pacifica radio station that George Carlin’s Filthy Words monologue was indecent. Lee and Washburn were the sole non-concurring commissioners in the majority’s judgment, which indicated that night-time broadcast of the monologue may not violate indecency regulations. Reid and Quello both concurred, but indicated that Filthy Words would be inappropriate no matter when broadcasted, while Robinson and Hooks urged a narrow reading of the majority opinion. These opinions do not accord with our ideal point estimates, and in fact the item discrimination parameter ($\Lambda_2$) is statistically indistinguishable from 0 – the 95% confidence interval is $(-1.90, 1.31)$.

### 6.2 Stationarity

So far, we have also assumed that ideal points are stationary – that commissioners do not change preferences over time. Because the terms are relatively short and staggered, this assumption seems reasonable, and is consistent with qualitative accounts (e.g., Cohen, 1986, p. 694). On the other hand, the data includes two commissioners who serve substantially longer periods than others: Quello and R. Lee served some 23 and 28 years, respectively (although Lee is only in the dataset for 16 years). The consequence may be that Quello or Lee “anchor” commissioners across time. If Quello has become more moderate over time, recent appointees to the FCC, such as Adelstein and Copps, may appear unduly liberal.

To investigate this possibility, I reestimate the model fitting relaxing stationarity by estimating separate ideal points for the three terms to which Lee and Quello were each reappointed during the observation period. A promising alternative is a temporal smoothing prior of Martin and Quinn (2002). Ideal point results are largely similar. Row 6 in Figure 7 plots parameter results, which are largely identical to Figure 6.

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43 *In re Pacifica Foundation Station WBAI*, 56 F.C.C.2d 94 (1975).
Lastly, even if the trend has been towards deregulation over the past few decades, that does not invali-
date ideal point measures. Many changes in the FCC are induced by exogenous technological and market changes, so commissioners serving in 1934, faced with new market circumstances, may certainly reach different substantive outcomes in 1996. This is a strength of the ideal point model – what fixes these dimen-
sions is not some substantive outcome (where the desirability, for example, of natural monopoly regulation may change as technological innovation induces competition), but relative voting records of commissioners.

6.3 Presidential Party

Two possible objections to the way in which we’ve measured the effects of presidential party exist.

We’ve assumed the effect of a Republican/Democratic president to be constant – \( \delta_3 \) does not vary across appointing presidents. This may not be the case if, for example, Clinton chooses more conservative nomi-
nees than Eisenhower, or if the institutionalization of the personnel process changed the role of presidential partisanship over time. To relax this assumption, I estimate a random intercept model, rewriting Equation 3 as:

\[
\begin{align*}
\phi_c & \sim N(\delta A_c + \gamma_p, 1) \\
\delta & \sim N_+(0, 25) \\
\gamma_p & \sim N(\mu_\gamma, \tau_\gamma^2)
\end{align*}
\]

where \( p = \{1, ..., P\} \) indexes each Presidency. I use diffuse priors for \( \mu_\gamma \) and \( \tau_\gamma^2 \). Row 8 in Figure 7 plots the marginal posterior of \( \delta \). The parameter on commissioner partisan affiliation remains robust. Appendix D plots random intercepts. Even relaxing the pooling assumption of presidential party, these results strongly suggest that the partisan affiliation of nominees matters, and weakly suggest that statutory requirements lead to more extreme cross-party appointees.

We’ve also used as presidential partisan affiliation the measure of the party of the initial appointing president. Since some commissioners are reappointed by presidents of differing parties, the measure may not be accurate. Several points bear mention. First, it would seem intuitively that presidents are more likely to reappoint more moderate commissioners, but our findings refute that. Second, this criticism is inapplicable for the vast majority of appointees. Few serve multiple terms. And of those that serve multiple terms, the party of the reappointer doesn’t necessarily differ from the party of the initial appointer. Quello, for example, is always appointed by a Republican. Bartley, for whom we only have some 6 months of data, is initially appointed by Truman, reappointed by Eisenhower, and then reappointed by Johnson; so he is correctly measured as a Democratic appointee under both coding criteria. The choice of initial appointing president matters only for two appointees: (1) Lee, initially an Eisenhower appointee, also serves as a
Johnson and Nixon appointee, and (2) Powell, a Clinton appointee who is reappointed and elevated as chair by George W. Bush. To assess sensitivity to this measurement, Row 7 of Figure 7 presents results from reestimating the model by changing the presidential affiliation for Lee and Powell. The results are virtually identical.

### 6.4 Remaining Concerns

Here I address several remaining methodological concerns. First, the paper assumes a single policy dimension. Overall, this seems reasonable given the purpose of learning about commissioners’ underlying regulatory philosophy. On the other hand, areas like content regulation may present a second (orthogonal) dimension not captured by the model (Hammond and Hill, 1993). As a statistical matter it is straightforward to incorporate a second dimension (see, e.g., Clinton et al., 2004), although computation, identification, and interpretation become more complicated. But most academic and popular accounts tend to discuss commissioners according to some conception of a single liberal / conservative dimension, and qualitative checks on cases and commissioners discussed in Subsection 5.1 bear this out.

Second, the analysis assumes that votes on cases are (conditionally) independent. Given the vast caseload of the FCC and the fact that multiple decisions are often made on a single underlying claim, one might argue that this assumption is violated, leading uncertainty intervals to be falsely narrow. Of course, this criticism applies to the overwhelming majority of applications of ideal point methods, and the assumption can be relaxed by modelling case dependence (Tuerlinckx and De Boeck, 2004).

Lastly, one might wonder about the interpretation of extreme cross-party commissioners when ideal points are identified only using non-unanimous cases. Can commissioners really be classified as more extreme post-1980, when the dissent rate has decreased from 17.1% in the pre-1980 period to 4.5% post-1980 (p-value from t-test≈0)? The intuition is that a unanimous case conveys little information about relative policy preferences. The dissent rate may vanish to 0 if all commissioners are gradually replaced with Marxists, but Marxists would still be identified as more liberal (as long as there is sufficient overlap with non-Marxists). Of course unanimous cases may still contain valuable information. The development of new measurement methods to address the concerns above may be fruitful areas for further research.

### 7 Conclusion

Writing the year the Interstate Commerce Commission was created, then-political science professor Woodrow Wilson concluded that “administration lies outside the proper sphere of politics” and that “[t]he object of
This study differs on both counts: administration is intricately interwoven in the sphere of politics, and the object of administrative study is not to remove executive methods from empirical experiment, but to execute methods that elucidate stable patterns in bureaucratic behavior.

Methodologically, this paper provides the first application of Bayesian ideal point models to voting records of independent commissioners. The methods are a straightforward generalization of techniques developed in psychometrics, educational testing, and political science. The application illustrates how the substantive knowledge of law – in pointing to fine-grained distinctions between types of votes – informs statistical analysis.

Substantively, some uncertainty remains about whether partisan requirements in fact increase the number of cross-party commissioners (due to low variation across agencies). But if they increase the number or change the type we have learned a considerable deal. First, contrary to the hopes of New Deal visionaries, partisanship explains a substantial amount of heterogeneity in FCC adjudications and rulemakings. Roughly three quarters of votes in non-unanimous cases can be explained by partisanship. Second, and more important to the question of partisan requirements, the partisan affiliation of commissioners has the largest and most robust explanatory power over votes compared to presidential affiliation. This refutes the idea that cross-party appointees don sheep’s clothing, and is consistent with partisanship requirements as a rational instrument for the enabling Congress (see, e.g., Moe [1989], Lewis [2004], Sunstein [2005]). Third, there is evidence that cross-party appointees post-1980 have been even more extreme than party-line appointees.

The evidence suggests a rupture in Senatorial norms, but is not quite as robust as the finding about commissioner affiliation. Either (a) post-1980 cross-party appointees are more extreme than party-line appointees, or (b) presidential affiliation has no independent effect at all. Whether or not this is desirable as a normative matter is, of course, another issue.

In sum, contrary to much anecdotal wisdom, there is solid empirical evidence that partisan requirements do in fact constrain. And contrary to Woodrow Wilson in 1887, the Congress, in implementing such requirements, appeared to know very well that partisanship was not a paper tiger.

Appendix

A Randomization Inference

To test whether commissioner partisan affiliation $A$ is independent of presidential partisanship $P$ when partisan requirements do not constrain an appointment, I selected 65 FCC unconstrained appointments (not commissioners), summarized in Table 3. These appointments are unconstrained either because of multiple vacancies or an independent serving on the commission.

<table>
<thead>
<tr>
<th>President’s Party</th>
<th>Democratic</th>
<th>Republican</th>
</tr>
</thead>
<tbody>
<tr>
<td>Commissioner’s Party</td>
<td>Democratic</td>
<td>30</td>
</tr>
<tr>
<td></td>
<td>Independent</td>
<td>2</td>
</tr>
<tr>
<td></td>
<td>Republican</td>
<td>5</td>
</tr>
</tbody>
</table>

Table 3: Summary of 65 unconstrained FCC appointments

Using the same notation introduced in Subsection 4 and letting $A_i = 0$ when $i$ is independent, the test statistic is a simple difference-in-means of partisanship:

$$
\tau_{obs} = \frac{\sum_i^C (A_i | P_i^{obs} = 1)}{\sum_i^C P_i^{obs}} - \frac{\sum_i^C (A_i | P_i^{obs} = 0)}{\sum_i^C (1 - P_i^{obs})}
$$

In potential-outcomes notation, the null hypothesis is:

$$
A_i(P_i = 1) = A_i(P_i = -1) \forall i.
$$

If $P$ were independently assigned for each commissioner (and no independents were appointed), we could use Fisher’s exact test for a $2 \times 2$ contingency table, but the unit of assignment is actually each Presidency, so the test would exhibit Type I error. To account for this, we can use randomization inference (Ho and Imai 2006; Donohue III and Ho 2007). For 12 administrations, 6 Democratic and 6 Republican, we randomly draw presidential partisanship $P^t$ to account for all historically-possible combinations of Presidencies, yielding $\binom{12}{6} = 924$ possible vectors of $P^t$. For each draw $t$, we can calculate the test statistic $\tau^t$ under the null. The exact p-value $P(\tau^{obs} \leq \tau^t) = \sum_{t}^I I(\tau^{obs} \leq \tau^t)$, where $I()$ is an indicator function, is 0.001. The last column of Table 1 calculates analogous p-values for other agencies.

B Case Selection

The aim of the study was to select, classify, and code all decisions taken by FCC commissioners in adjudications and rulemakings from 1965–2006. Westlaw provides coverage of 1 F.C.C. 2d (1965) through 104 F.C.C. 2d (1986), and the F.C.C. Rcd. from Nov. 1986 to the present. Both of these sources are the official record of all decisions by the FCC, and while the F.C.C. Rcd. includes considerably more information beyond votes by the commission (cases decided by each FCC bureau, daily digests, etc.) there is no reason to believe that coverage of commissioner votes changed across volumes. Automating data collection required fine-tuning Westlaw search strings to (a) account for variability in reporting practices over time, and (b) exclude reported decisions involving no commissioner votes (e.g., decisions solely by Administrative Law Judges). The search strings, specific to each time period, were:
1. 1965–1984: `fcc +1 Y –1900 +255 order rule! notice inquiry +10 adopted +3 Y +12 "by the commission" & da(Y)

2. 1985–1989: `order rule! notice inquiry +255 adopted +3 Y +12 "by the commission" & da(Y)

3. 1990: `fcc +10 Y –1900 +255 adopted +3 Y +20 order rule! inquiry notice +10 "by the commission" & da(Y)


5. 1996-2006: `fcc +1 Y –1900/2000 +10 adopted +3 Y +30 "by the commission" & da(Y)

where Y indicates the year. These search strings retrieve the large majority of all FCC adjudications and rulemakings, which enabled electronically coding case identification information. Nonetheless, these search strings miss a small fraction of cases. From a random sample of cases, roughly 5% are improperly excluded from the search results; compared to the F.C.C. Reporter, these exceptions include: (a) misspellings due to Westlaw (but not the FCC Reporter), see 85 F.C.C.2d (“commisison’); 85 F.C.C.2d 723 (“Adoted’); (b) slight variations in preamble, see 85 F.C.C.2d 713 (“Gentlemen:” in lieu of “By the Commission:’); 11 F.C.C. Rcd 14,171 (separate comment deadlines separating opinion from heading); 11 F.C.C. Rcd 14,100 (“F.C.C. No. 96’); and (c) failure by Westlaw to include salient case identification information (e.g., citation). This sample suggests that cases are highly likely to be missing completely at random, inducing only inefficiency in reported estimates. All case information was cross-validated by handcoding.

C Senate and Commerce Committee Polarization

Figure 10: Evidence of Senate and Commerce Committee Polarization. The left panel plots the difference in party median ideal points (DW-NOMINATE) across time for the full Senate, with bootstrapped 95% confidence intervals and a vertical line denoting 1980. The middle panel plots the same difference for members of the Senate Commerce Committee. The right panel plots party medians, and shading to denote range of ideal points, with a vertical line denoting 1980.

To investigate whether party polarization in the Senate and the Commerce Committee may be driving the post-80 cross-party extremism finding, I merged DW-NOMINATE estimates of Senate ideal points (Poole and Rosenthal, 1997) with congressional committee membership data (Nelson, 1993; Stewart III and Woon, 2005). Figure 10 plots results. The left panel plots the difference in the party median ideal points (the partisan gap), with bootstrapped 95% confidence intervals, for the full Senate over time. Although
increasing before 1965, the partisan gap is roughly constant for the first 15 years of FCC data (1965–1980), and increases significantly after 1980. The middle panel plots the same partisan gap within the Commerce Committee. After 1980, there is a sharp and statistically distinguishable increase in the gap. The right panel plots party median ideal points and ranges over time. Prior to 1980, there is some overlap in party ideal points, suggesting that bipartisan coalitions are easier to build. After 1980, ideal points of the parties separate completely. These results are consistent with batching as a means to overcome an increase in political transaction costs of confirmations.

D Random Intercept Results

![Random Presidential Intercepts](image)

Figure 11: Random Presidential Intercepts, sorted in magnitude. If presidential dominance is correct, we should see that Republican intercepts (in red) are positive and Democratic intercepts (in blue) are negative, which is not the case. Coefficients are statistically indistinguishable from 0, but there are distinguishable differences between presidents.

While each of the presidential intercepts intersects with the origin, Democratic intercepts tend to be positive, illustrating cross-party extremism. We can also distinguish presidential practices: in large part due to Furchtgott-Roth and Copps, there is a roughly 95% posterior probability that George W. Bush’s appointees, holding constant the partisan affiliation of commissioners, are more liberal than Clinton’s appointees.
References


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