

Media as Political Currency*

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Abstract

This paper studies how broadcast licenses become instruments of political coalition-building when executives retain discretion over media allocation. We examine this mechanism in Brazil’s 1987–88 National Constituent Assembly, linking all radio and TV licenses to every 1986 federal candidate. We develop a framework predicting that executives target ideologically proximate legislators, who align on institutional but not ideological votes, and gain electoral returns that scale with local media-market concentration. We document that winning a seat raised concession receipt by 13.6 percentage points, with larger effects among government-aligned legislators and null effects in every placebo election. Concession recipients show a large gap in government alignment on institutional versus ideologically divisive votes—the pattern the model predicts. The incumbency premium concentrates in non-competitive media markets where a new license dominates local information. The findings characterize broadcast concessions as discrete, market-structure-dependent productive political capital.

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1 Introduction

The strategic exchange of resources for legislative support is a central feature of democratic politics. Executives frequently rely on fiscal transfers and other forms of distributive spending to build and maintain congressional coalitions (Golden and Min, 2013; Stokes et al., 2013; Cruz and Keefer, 2015; Bonilla-Mejía and Morales, 2024; Brollo and Nannicini, 2012). Together, this literature shows that targeted benefits can shape legislative behavior, discipline coalition partners, and strengthen governing majorities. A parallel body of research highlights the political importance of the media, documenting how news exposure and media bias affect electoral behavior, political accountability, and long-run institutional development (Strömberg, 2015; Snyder and Strömberg, 2010; DellaVigna and Kaplan, 2007; Enikolopov et al., 2011; Grossman et al., 2022; Chong et al., 2017; Prat, 2018; DellaVigna and La Ferrara, 2015; Ferraz and Finan, 2008), and showing that executives who can capture media reporting gain substantial political leverage (Besley and Prat, 2006; Durante et al., 2019). Yet despite extensive evidence that media shapes electoral and policy outcomes, we know almost nothing about how politicians come to control media in the first place and whether executives strategically allocate this control to influence the political process. McMillan and Zoido (2004) provide a rare window into this allocation margin, documenting from leaked archives how Peru’s Montesinos valued control of a single television network at roughly 100 times the price of a congressional vote; yet clean quantitative evidence of legal, public-discretion allocation in a transitional democracy remains scarce. Building on theoretical work that models media capture as favor exchange between politicians and outlet owners (Szeidl and Szucs, 2021), this paper provides direct empirical evidence on the allocation side of that exchange.¹

Informational assets such as broadcast licenses constitute a distinct class of political

¹T. Boas and Hidalgo (2011) document, using an RD design, that incumbents are more likely to obtain community-radio licenses and that ownership raises vote shares—evidence on electoral behavior that complements our focus on the executive-legislative allocation mechanism.

resource from the fiscal transfers studied in the existing literature. Fiscal transfers are *consumptive*: the benefit is delivered, appropriated by the district, and then exhausted. Broadcast licenses are *productive* capital: they generate an informational stream that sustains the recipient’s electoral position as long as the holder remains politically active, at no recurring fiscal cost to the Executive. When a transitional executive faces a legislature with the power to write new constitutional rules, broadcast concessions therefore combine the targeting precision of individual-level patronage with the durability of a standing electoral asset. We show that concessions allocated through close-election bargaining in 1987–88 Brazil were retained selectively and are associated with a market-structure-dependent incumbency premium—the empirical signatures the productive-political-capital framework predicts.²

We study this mechanism in a canonical empirical setting—Brazil’s 1987–88 National Constituent Assembly—where executive discretion over media licensing was near-unilateral and the informational value of local radio was substantial. The mid-1980s in Brazil represented a period of profound institutional uncertainty, with an unpopular executive, a fragmented Congress, and the politically consequential task of drafting a new Constitution (Mainwaring, 1999). At that time, Brazilian regulation granted the president substantial discretion over the allocation and renewal of radio and television licenses (Lopes, 2018; Motter et al., 2019). Moreover, in many regions radio stations were the dominant source of local political information, giving their owners considerable influence over voter perceptions. This combination of discretion and informational power created a setting in which broadcast concessions functioned as targeted political transfers, exchanged for legislative support during a foundational moment of constitutional redesign.

To study this executive–legislative exchange, we assemble new data linking all radio

²The closest empirical antecedent is Bonilla-Mejía and Morales (2024), who use the same close-election RDD design to show that Colombian legislators trade votes for infrastructure projects. Section 3 develops the structural features that separate broadcast concessions from that pork-patronage framework.

and television concessions granted between 1982 and the 2000s to individual politicians. We combine these administrative records with detailed electoral data from the 1986 Federal Deputy election, digitized roll-call votes from the 1987–88 National Constituent Assembly (NCA), longitudinal election results from 1990 to 2010, and extensive biographical information on candidates, their families, and their prior media involvement. This dataset allows us to follow every candidate for the NCA—winners and losers—and track whether, when, and what type of broadcast outlets they acquired.

To guide our empirical analysis, we develop a simple theoretical framework in which the Executive allocates media concessions to maximize legislative support. In the model, owning a media outlet increases a legislator’s electoral valence, reducing the electoral cost of siding with the government. In equilibrium, the Executive targets legislators who are ideologically proximate, as they require lower compensating transfers. Our framework clearly predicts which legislators should be targeted with concessions, how recipients should align in legislative votes, and the electoral advantage they should experience.

Guided by this framework, our empirical strategy combines three complementary designs. For *allocation* we use a narrow-margin Regression Discontinuity Design (RDD) within party–state coalitions in the 1986 election: candidates who barely won became members of the NCA, while those who barely lost did not. For *legislative response* we estimate within-deputy partial correlations on the 451-deputy roll-call sample, with the within-deputy contrast between institutional and ideological topics differencing out deputy-level selection on overall pro-government disposition. For *heterogeneous persistence* we estimate heterogeneous ITT effects of the 1986 close-election win on subsequent electoral outcomes, splitting the sample by the pre-NCA *Fistel* broadcasting density per municipality (state-level), under the framework of Becker et al. (2013).

We organize our findings around three pieces of evidence, each confirming a prediction of the framework.

The first piece is allocation: barely elected deputies were 13.6 percentage points more likely to receive a radio or TV concession during 1987–88, with effects roughly 50 percent larger among government-aligned legislators (18–21 percentage points). Placebo designs on every federal election from 1982 to 1998 yield null effects in every cycle except 1986.

The second piece is legislative response. Recipients show a +33 to +43 pp within-deputy gap in government alignment on institutional versus ideological constitutional items across pooled specifications, rising to +45.4 pp in the within-deputy panel; because the gap is identified from the within-deputy contrast between each deputy’s two index values, deputy-level selection on overall pro-government disposition is differenced out by construction (Cinelli–Hazlett $RV_{q=1} = 0.246$ bounds confounders at any level).

The third piece is heterogeneous persistence. The post-NCA incumbency premium concentrates in non-competitive media markets where a single new license dominates the local information environment: the non-competitive-minus-competitive difference on being elected in 1994 is +13.9 pp at baseline, rising to +33.4 pp under donut specifications. This market-structure gradient is the strongest mechanism test the setting allows: a consumptive transfer—fiscal pork or patronage—generates no such conditional persistence, while a productive informational asset generates lasting electoral returns precisely where it dominates the local media landscape and decays where it does not.

The three pieces characterize broadcast concessions as a politically targeted, individually held, market-structure-dependent durable resource.

This paper contributes to three strands of the literature. First, we advance research on how media structure shapes political competition (e.g., Strömberg, 2004; Gentzkow et al., 2011; Gavazza et al., 2018) and how politician-owned media produces durable electoral advantages (Durante et al., 2019; Barone et al., 2015; Cavgias et al., 2024a). T. Boas and Hidalgo (2011) provide the closest predecessor, showing with an RD design that incumbents are more likely to obtain community-radio licenses and that ownership raises their vote

share. A related demand-side literature shows that media disciplines politicians: press coverage raises voter information (Snyder and Strömberg, 2010) and audit publicity deters corruption (Ferraz and Finan, 2008). We study the supply side of the same relationship: politicians captured broadcast licenses before that discipline could operate, shaping the information environment to their advantage.³ We extend this to a setting where licenses are allocated directly by the national Executive as a legislative-bargaining tool, provide causal evidence on the allocation margin, and document that the incumbency premium is heterogeneous in the dimensions the productive-political-capital framework predicts.

Second, the paper speaks to the literature on patronage and distributive politics (e.g., Golden and Min, 2013; Stokes et al., 2013; Cruz and Keefer, 2015; Bonilla-Mejía and Morales, 2024). We show that informational assets constitute a structurally distinct class of political currency: unlike fiscal transfers, which are exhausted at delivery, a broadcast license generates a durable stream of electoral valence at no recurring cost to the Executive. This productive-vs-consumptive distinction has testable implications for how the recipient’s political career evolves—implications we bring to the data in Section 5. The theoretical structure follows the favor-exchange environment of Szeidl and Szucs (2021) with the government as allocator and legislators as recipients.

Third, the heterogeneous-persistence result is the strongest mechanism test the paper delivers and the finding that most directly separates the productive-capital interpretation from pure patronage. A consumptive transfer generates no conditional persistence by market structure; a productive informational asset should generate lasting returns precisely where it dominates the local media landscape and attenuate where it does not. The post-NCA incumbency premium does exactly this: it concentrates in non-competitive markets and decays in saturated ones. This market-structure conditionality connects to the media-economics literature on why media influence is largest where outlets are scarce

³We do not study whether recipients adjusted editorial content after receiving a concession; evidence on the content channel in related settings is provided by Gentzkow and Shapiro (2010).

(Strömberg, 2004; Gavazza et al., 2018), and distinguishes the productive-capital mechanism from both consumptive transfers and unconditional valence advantages.

The remainder of this study is organized as follows. Section 2 explains why Brazil’s transition to democracy—and the 1986 Constituent Assembly in particular—created a rare opportunity to observe how executives deploy control over broadcasting licenses. Section 3 formalizes the mechanism motivating our analysis, presenting a simple framework in which media concessions raise legislators’ electoral valence and lower the ideological cost of supporting the Executive. Section 4 describes the data sources and variable construction, detailing how we link electoral records, media concession data, and legislative voting behavior. Section 5 tests the model’s predictions through three pieces of evidence—allocation (the close-election RDD on concession receipt), legislative response (within-deputy issue-heterogeneity contrast on roll-call votes), and heterogeneous persistence (the market-structure heterogeneity in the incumbency premium). Finally, Section 6 concludes by discussing the broader implications of our findings for democratic accountability, institutional design, and the political economy of media control.

2 Background

2.1 Broadcasting concessions and the regulatory framework

Broadcasting concessions. A broadcasting concession (*outorga*) in Brazil is a government grant of the exclusive right to operate a radio or television station on a specified frequency in a specified municipality, typically for ten years (radio) or fifteen years (television), renewable. Because the electromagnetic spectrum is a finite public resource, a concession is an economically valuable asset: it carries a barrier to entry (competitors are prohibited from broadcasting on the same frequency in the same market), an advertising revenue stream, and, in the local media markets that dominate Brazilian radio in particu-

lar, a near-monopolistic position on political information. For politicians, the concession offers a second, distinct dividend—the ability to shape the informational environment in which their own electoral performance is judged—which is a margin that standard commercial valuation of the asset does not capture. The empirical strategy in this paper exploits precisely this combination of economic and political value.

The allocation mechanism distinguishes the Brazilian setting from the OECD norm. In most large democracies, broadcasting licenses are allocated by independent regulatory bodies insulated from executive discretion (Hallin and Mancini, 2004). Brazil is unusual in that the president had direct discretionary authority over AM and television concessions, with only a thin technical filter from the Ministry of Communications and congressional ratification serving largely as a rubber stamp. This concentration of allocation authority in the presidency is what makes concessions a natural instrument of executive-legislative bargaining.⁴

Brazil's system of media regulation historically granted the federal executive extensive discretion over the allocation and renewal of radio and television licenses. The main legal framework from the 1960s⁵ allowed the president to designate the holder of AM radio and national TV concessions from among technically qualified applicants, while the Ministry of Communications exercised similar discretion over FM broadcasters (Motter et al., 2019). During the military regime, Decree-Law 236/1967 expanded these powers by allowing the state to fine, suspend, or revoke licenses without judicial oversight, a threat frequently used to discipline media outlets even though few revocations became formal actions (Lopes, 2018). These institutions created a regulatory environment in which broad-

⁴Qualitative evidence from Peru under Fujimori suggests that political capture of broadcast media can be economically substantial even under formally competitive arrangements: McMillan and Zoido (2004) document that Montesinos valued the editorial line of a single television network at roughly 100 times the price of a congressional vote—the authoritarian-exchange analogue of the legal-democratic mechanism studied here.

⁵Specifically the 1962 Código Brasileiro de Telecomunicações and the 1963 Regulamento dos Serviços de Radiodifusão. A contemporary example of political discretion in the allocation of radio concessions is reproduced in Figure B1.

cast concessions were highly valuable and heavily dependent on presidential favor.

2.2 The 1986 National Constituent Assembly

By the early 1980s, Brazil was undergoing a process of redemocratization, with the military gradually transferring political authority back to civilian leaders after two decades in power. President José Sarney, who unexpectedly assumed office after Tancredo Neves's death, inherited a fragile economic environment marked by high inflation and the eventual failure of successive stabilization plans. Politically, Sarney lacked a cohesive partisan base and governed a Congress characterized by extreme fragmentation and weak party discipline. The largest governing party, the PMDB (Brazilian Democratic Movement Party), was a broad and internally heterogeneous centrist coalition that had served as the main legal opposition under the military regime. The PFL (Liberal Front Party), formally a center-right party and an important coalition partner, was composed largely of politicians who had supported the military regime and then repositioned themselves during the transition. Both parties exhibited substantial rates of legislative defection and weak internal cohesion (Mainwaring, 1999), complicating the Executive's efforts to build stable majorities.

The nationwide elections held in November 1986 selected the members of the National Constituent Assembly (NCA), composed of 487 deputies and 72 senators, and were strongly influenced by the temporary popularity generated by the *Plano Cruzado*, an ambitious price-freeze stabilization program. Shortly after the election, however, inflation resurged and the plan quickly collapsed, leading to a sharp and well-documented decline in the government's popularity.⁶ Although the PMDB–PFL coalition secured a large number of seats, it remained internally divided (Motter et al., 2019). As a result, the Executive needed to assemble majorities vote by vote on several contentious constitutional issues:

⁶Appendix A, and particularly Figure B2, reproduces newspaper headlines from late 1986 to 1988 documenting the collapse of the *Cruzado* plan and the rapid decline in presidential approval.

some were high-stakes institutional questions on which the government held a unified position—most notably the length of the presidential term and the choice of a presidential over a parliamentary system of government—while others split the governing coalition along pre-existing ideological lines, including property rights and agrarian reform.

In this environment, media concessions functioned as a particularly effective political currency. Broadcast outlets were economically attractive and electorally powerful given the central role of radio and television as sources of political information in Brazil, especially in areas with limited newspaper circulation (T. C. Boas, 2005; Porto, 2003; Lopes, 2018). Local radio stations, in particular, were often the primary venue for political news in smaller municipalities (Lopes, 2018). Historical accounts describe an uptick in the distribution of concessions during periods of intense political bargaining. For example, Motter et al. (2019) documents that during Sarney’s presidency, approximately ninety legislators received radio and TV concessions over the course of the NCA. Although these accounts are descriptive and not intended to establish causal patterns, they underscore how valuable media outlets were perceived to be as bargaining instruments in the very period analyzed in this paper.

The phenomenon is not confined to a small number of outliers. Table B1 (Appendix A) lists twelve representative politician-broadcasters whose links to radio or television are documented in archival and secondary sources, including the congressional investigations reconstructed by Motter et al. (2019) and the broadcasting ownership records compiled by Lopes (2018). The cases span eleven states and both large governing parties; ownership is frequently registered through spouses, children, or close business associates, a feature that our matching procedure (Section 4) is designed to capture.

The structure of Brazil’s media market further magnified the political returns to owning broadcast outlets. Television expanded rapidly during the military regime as part of a broader modernization strategy but evolved into a highly concentrated private market

dominated by a small number of conglomerates (Lima, 1988; Motter et al., 2019). A sizable literature documents how media owners have historically used their outlets to promote political allies and reduce the visibility of rivals (Costa and Brener, 1997; Santos and Capparelli, 2005; Stevanim and Santos, 2011). Complementing this qualitative evidence, recent quantitative work shows that owning a radio station increases a politician’s vote share (T. Boas and Hidalgo, 2011; Lopes, 2018) and that voters respond to corruption audits only when local media are not controlled by political elites (Gehrke, 2019). Together, these institutional and informational features created a political landscape in which the executive had strong incentives to allocate media concessions strategically, legislators had strong incentives to obtain them, and media ownership plausibly shifted electoral prospects.

These characteristics make the Brazilian NCA a particularly compelling environment for studying how executives allocate media concessions, how these concessions shape legislative behavior, and how they translate into long-term political advantage.

3 Conceptual Framework

This section develops a one-shot allocation game in which the Executive uses broadcast concessions as discrete, durable, market-structure-dependent political capital to secure legislative support during a constitutional bargaining episode. Three features distinguish the framework from canonical pork-patronage models (Lindbeck and Weibull, 1987; Dixit and Londregan, 1996; Bonilla-Mejía and Morales, 2024): concessions are binary and identity-embodied rather than divisible; legislators face issue-differentiated deviation costs so that the same license can flip an institutional vote but not a distributive one; and the asset’s electoral value depends on local media-market structure. The framework derives three empirical predictions taken to the data in Section 5, drawing on the media-capture environment of Besley and Prat (2006), the government-control-of-media framework of Gehlbach and Sonin (2014), and the political-exchange structure of

Groseclose and Snyder Jr. (1996).

Environment. There are N legislators $\ell \in \{1, \dots, N\}$, each representing a single-member district with median voter ideal point $p_\ell^* \in [0, 1]$ drawn from a continuous distribution F . The Executive E has $K < N$ broadcast licenses available for allocation during the legislative term and announces a (possibly stochastic) allocation rule before votes are cast. We denote the allocation by $b_\ell \in \{0, 1\}$, with $b_\ell = 1$ if legislator ℓ receives a license and 0 otherwise, subject to $\sum_\ell b_\ell \leq K$. The constraint $K < N$ reflects spectrum scarcity: radio frequencies are physically finite, and the Ministério das Comunicações processed concessions through a regulated pipeline (the *Fistel* registry) whose throughput per period was institutionally bounded (Lopes, 2018)—placing the 1987–88 NCA setting clearly in the under-supply regime ($|A^*| > K$), where the Executive rations within the targetable set.⁷

The legislature votes on two classes of issues. *Institutional* issues—indexed I —concern the structure of executive power (e.g., system of government, length of presidential term). The Executive has a unified preferred position $g^I \in \{0, 1\}$, and voters’ preferences on these issues exhibit low ideological salience: knowledge of the issue is thin and partisan priors are weak. *Ideological* issues—indexed D —concern redistribution and the regulation of economic life (e.g., agrarian reform, labor rights, property). The government may or may not have a unified position; voters hold strong, ideologically anchored priors. Each legislator chooses a vote $v_\ell^I, v_\ell^D \in \{0, 1\}$.

Electoral game. Reelection is contested between the incumbent ℓ and a challenger who positions at the district median on both issues. Each voter j in district ℓ draws an idiosyncratic preference shock $\varepsilon_j \sim U[-\sigma, \sigma]$, capturing unobserved heterogeneity in political

⁷In the 1987–88 NCA period, roughly 80 new concessions were granted to legislators in our RDD sample, against a governing coalition of approximately 366 PMDB and PFL deputies (historical accounts document approximately 90 across the full NCA membership including senators). An alternative microfoundation treats K as endogenous: the Executive stops granting concessions when the marginal political cost of an additional license equals the marginal legislative benefit. We view spectrum scarcity as the more binding constraint in this setting, but a political-cost model delivers the same qualitative predictions.

tastes. Voter j supports the incumbent if and only if

$$-\omega^I(v_\ell^I - \hat{v}_j^I)^2 - \omega^D(v_\ell^D - \hat{v}_j^D)^2 + \Gamma b_\ell + \varepsilon_j \geq -\omega^I(\hat{v}_j^I - p_\ell^*)^2 - \omega^D(\hat{v}_j^D - p_\ell^*)^2, \quad (1)$$

where $(\hat{v}_j^I, \hat{v}_j^D)$ are the voter's ideal positions, $\omega^I < \omega^D$ are issue-specific salience weights—voters discount the institutional dimension—and $\Gamma \equiv \Gamma(m) > 0$ is the valence delivered by the broadcast license through programming, name recognition, and editorial reach (m denotes local market density and is suppressed in the static game; Prediction 3 makes this argument explicit). Aggregating over voters in district ℓ , the incumbent's *expected* vote share—the fraction of voters for whom the support condition holds—takes the form

$$s_\ell(v_\ell^I, v_\ell^D, b_\ell) = \frac{1}{2} - \omega^I \Delta_\ell^I(v_\ell^I) - \omega^D \Delta_\ell^D(v_\ell^D) + \Gamma b_\ell, \quad (2)$$

where $\Delta_\ell^j(v_\ell^j) \geq 0$ is the loss in median-voter utility from a vote against district preference on issue j , normalized so that $\Delta_\ell^j(g^j) = 0$ when the Executive's position coincides with the district's.

Legislator's problem. Legislators are office-motivated and choose (v_ℓ^I, v_ℓ^D) to maximize their expected vote share s_ℓ , which is equivalent to maximizing their reelection probability since $\Pr(s_\ell \geq \frac{1}{2})$ is increasing in s_ℓ . The cost of voting with the Executive on issue j when this conflicts with the district median is the lost expected vote share

$$c_\ell^j = \omega^j \cdot \Delta_\ell^j, \quad j \in \{I, D\}, \quad (3)$$

which inherits the salience asymmetry $\omega^I < \omega^D$. We make the distributional assumption that Δ_ℓ^I and Δ_ℓ^D are drawn from continuous distributions with bounded supports $\Delta_\ell^I \in$

$[0, \bar{\Delta}^I]$ and $\Delta_\ell^D \in [\underline{\Delta}^D, \bar{\Delta}^D]$ with $\underline{\Delta}^D > 0$, and we restrict the parameter space so that

$$\omega^I \bar{\Delta}^I \leq \Gamma < \omega^D \underline{\Delta}^D. \quad (4)$$

Under (4), $c_\ell^I \leq \Gamma < c_\ell^D$ holds with probability one: a license is valuable enough to flip an institutional vote but not a distributive one. The restriction $\omega^I < \omega^D$ reflects a documented feature of Brazilian electoral politics: distributive issues (agrarian reform, labor rights, property) crystallized sharp ideological cleavages with strongly positioned constituent bases, while institutional questions (system of government, presidential term length) carried low mass salience and were contested primarily among political elites (Pilatti, 2008; Mainwaring, 1999).⁸

Given an allocation b_ℓ , the change in expected vote share from aligning with the Executive on issue j is $\Gamma b_\ell - c_\ell^j$, which is positive iff $\Gamma b_\ell \geq c_\ell^j$. The legislator's best response is therefore

$$v_\ell^j = g^j \iff \Gamma b_\ell \geq c_\ell^j. \quad (5)$$

By (4), $b_\ell = 1$ flips the institutional vote but not the ideological vote, while $b_\ell = 0$ leaves the legislator at the district median on both. The asymmetry in (4) is the structural counterpart of the empirical issue-heterogeneity result documented in Section 5.

⁸The asymmetry is about *voter* salience in retrospective evaluation, not the Executive's priority ordering—institutional votes were cheap to flip precisely because constituents discounted them, even as the Executive prized them most. Cavgias et al. (2024b) show the first post-Assembly presidential election was dominated by economic concerns rather than the institutional choices the Assembly had just made. Formally, the restriction survives the limit $\omega^I \rightarrow \omega^D$ provided the supports are separated ($\bar{\Delta}^I < \underline{\Delta}^D$), so it requires only that distributive issues carry *some* additional electoral cost at every point in their support.

Executive’s allocation problem. The Executive seeks to maximize the number of institutional votes aligned with g^I .⁹ She solves

$$\max_{b \in \{0,1\}^N} \sum_{\ell=1}^N \mathbb{1}[\Gamma b_{\ell} \geq c_{\ell}^I] \quad \text{s.t.} \quad \sum_{\ell=1}^N b_{\ell} \leq K. \quad (6)$$

The solution to (6) is straightforward: define the *targetable set*

$$\mathcal{A}^* = \{\ell : c_{\ell}^I \leq \Gamma\}, \quad (7)$$

which contains every legislator whose institutional vote can be flipped at cost Γ . If $|\mathcal{A}^*| \leq K$ the Executive allocates a license to every $\ell \in \mathcal{A}^*$. If $|\mathcal{A}^*| > K$, any subset of K legislators drawn from \mathcal{A}^* is an equilibrium allocation; the equilibrium vote count is unique.¹⁰

Equilibrium. In the Subgame-Perfect Nash Equilibrium, the Executive draws from \mathcal{A}^* and recipients vote $v_{\ell}^I = g^I$; non-recipients vote at the district median p_{ℓ}^* .

Commitment is a maintained assumption: the Executive announces the allocation schedule before votes are cast, grounded in the public registration of concessions in the *Diário Oficial da União* prior to the relevant constitutional votes.¹¹

Three testable implications. The equilibrium yields three predictions taken to the data in Section 5. The first two follow directly from the static allocation game. For the third, we let the valence gain depend on local media-market conditions by writing $\Gamma(m)$, where m denotes the number of pre-existing broadcasting stations in the legislator’s electoral

⁹The ideological dimension is excluded from the objective because the governing coalition held internally divided positions on redistributive issues; the Executive’s net gain from securing any one faction’s ideological vote was cancelled by the cost of alienating the opposing faction. All qualitative results hold under a weak Executive preference on ideological issues provided $c^D > c^I$.

¹⁰The within-coalition gradient documented in Table 2 is consistent with any priority rule that rewards ideological proximity; we explore this heterogeneity in Section 5.

¹¹A retrospective reading—rewarding deputies who already voted loyally—is observationally equivalent given roll-call data alone. We do not model the reputational equilibrium explicitly; the *Diário Oficial* record serves as the institutional microfoundation.

market. Following Strömberg (2004), a broadcaster’s marginal contribution to political coverage decreases with the number of competing outlets: each station reaches a smaller audience share when information provision is more dispersed. We therefore impose the monotonicity condition

$$\Gamma'(m) < 0, \tag{8}$$

so that a new license dominates the local information environment in a thin market and contributes marginally in a saturated one. Under (8), every comparative static of the model with respect to Γ translates directly into a comparative static with respect to market thinness (lower m).

1) Targeting. The set of recipients \mathcal{A}^* contains legislators whose institutional deviation cost c_ℓ^I is below Γ . Equivalently, recipients are concentrated among legislators ideologically proximate to the Executive’s institutional position g^I . Legislators distant from g^I on the institutional dimension—typically deep opposition—fall outside \mathcal{A}^* and receive no licenses. **2) Issue-differentiated voting.** Equation (4) delivers a sharp prediction on roll-call behavior. Recipients vote with the Executive on institutional issues but not on ideological issues; non-recipients vote at the district median on both. The recipient–non-recipient gap therefore satisfies

$$\underbrace{\mathbb{E}[v_\ell^I | b_\ell = 1] - \mathbb{E}[v_\ell^I | b_\ell = 0]}_{> 0 \text{ (large)}} - \underbrace{\mathbb{E}[v_\ell^D | b_\ell = 1] - \mathbb{E}[v_\ell^D | b_\ell = 0]}_{\approx 0} > 0.$$

This within-deputy contrast—the institutional-versus-ideological Loyalty Index gap—is the model’s signature prediction and the empirical object documented in Table 3. Throughout Section 5, v_ℓ^I is operationalized as the *Institutional Index* (average government-alignment across the two institutional votes) and v_ℓ^D as the *Ideological Index* (average government-alignment across the five redistributive and regulatory votes).

3) Reelection enhancement. License holders enjoy an expected vote-share advantage of

$\Gamma(m) > 0$ relative to non-holders, translating into a higher reelection probability. Since $\Gamma'(m) < 0$, this advantage is larger in thinner markets: a new station dominates the local information environment when competition is sparse, and contributes marginally in an already saturated market. The empirical analysis tests this comparative static by interacting the close-election RDD with a measure of pre-NCA media-market competitiveness.

The framework departs from canonical pork models in two structural ways: the allocation is discrete rather than divisible, and the deviation-cost technology is issue-differentiated ($c^I < c^D$)—the feature that generates the Institutional-minus-Ideological differential documented in Section 5.

4 Data

Data sources. The empirical analysis draws on three datasets. First, we collect the universe of radio and television concessions granted by the Brazilian Ministry of Communications between 1979 and 2003 from the Ministry’s administrative records, obtained through a *Lei de Acesso à Informação* (LAI) request. These files list the firm (identified by its CNPJ, Brazil’s corporate tax identifier), the municipality and state of service, the service type (AM, FM, shortwave radio, or television), and the date of the grant. Second, we obtain the firm-level *Cadastro Nacional de Empresas* (CNE), a registry of all Brazilian firms that records the CNPJ, partner names and individual tax identifiers, and dates of incorporation and partner entry. Third, we use the electoral records of the *Tribunal Superior Eleitoral* (TSE) for the 1986 federal deputy elections that constituted the NCA, which cover candidate names, states, parties, vote counts, and ballot outcomes.

Matching concessions to politicians. We link firms to politicians in two stages. First, we perform an exact full-name match between CNE partners and 1986 NCA candidates

within state, extending the search to surname-level matches while excluding the six most common Brazilian surnames (Silva, Santos, Oliveira, Souza, Pereira, Lima) to avoid spurious links. This blind procedure yields a conservative baseline. Second, we use biographical material from FGV-CPDOC and the Chamber of Deputies' memory page to identify cases in which concessions were held by spouses, children, or close business associates of politicians. The biographical validation is the step at which the recurring use of *laranjas*—front men and straw owners—in Brazilian media politics becomes measurable. Qualitative accounts document such arrangements in multiple states (most prominently Bahia, Maranhão, Alagoas, Paraná, and Goiás), and the family-level matching is designed to capture them. Nonetheless, some cases of indirect or concealed ownership certainly remain undetected, biasing our ownership indicator toward zero. We denote by $\text{Concession}_i \in \{0, 1\}$ the indicator that deputy i received a radio or television concession during 1987–88; Radio_i and TV_i denote the type-specific variants.

Proven historical records and the two-coding strategy. We complement the automated CNE-TSE match with a second, independent source: the concession histories reconstructed from original Ministry of Communications paper files by Motter et al. (2019) for a congressional investigation into broadcast policy during the Sarney government. These proven historical records cover the full membership of the Constituent Assembly—documenting both recipients and pre-existing owners—and capture indirect ownership through spouses, relatives, and proxy shareholders (*laranjas*) that a tax-registry match cannot recover. We use two treatment codings throughout: *automated match only* (direct formal ownership, strict lower bound) and *archival-augmented* (direct plus documented indirect ownership, our preferred measure). Because concealed and indirect ownership channels are invisible to tax-registry matching, the automated match understates the true extent of political media ownership during the NCA period and should be read as a conservative lower bound; the archival records raise that lower bound by documenting

the well-known *laranja* arrangements. As these historical records were compiled from the parliamentary context, they primarily enrich the treated-side coverage rather than the comparison group—a feature we acknowledge transparently by reporting all headline results under both codings in Appendix Table A3.

Two-coding sensitivity: headline results. Appendix Table A3 reports all three headline results under both codings. The close-election RDD on concession receipt ranges from +6.0 pp (automated match, a conservative lower bound) to +13.6 pp (archival-augmented, our preferred estimate); the lower bound confirms that political assignment is detectable from direct formal ownership records alone. The pooled Loyalty Index is sensitive to coding choice, but the within-deputy Institutional-minus-Ideological differential—identified from each deputy’s own voting record—is stable and large under both codings (+0.380 vs. +0.454, each significant at the one-percent level).

Voting during the National Constituent Assembly. The 1988 NCA was a landmark event in Brazil’s transition to democracy, and several contentious topics—agrarian reform, labor rights, property ownership, and the structure of government—revealed sharp divisions between progressive and conservative blocs. The Sarney administration engaged actively in shaping legislative outcomes, which is the mechanism our empirical strategy targets. We construct roll-call data for seven of the most contested topics of the NCA—Property Rights, Working Hours, Right to Strike, Presidential System of Government, Agrarian Reform, Employment Stability, and the Five-Year Presidential Term—from the Chamber of Deputies’ published proceedings.¹² The government’s preferred position on each sub-vote is identified from Mario Covas’s floor speeches as government leader in the Assembly. For each deputy and each topic, we compute the share of sub-votes cast in the government’s direction, treating abstentions as opposed to the government’s position

¹²Available at: <https://www2.camara.leg.br/atividade-legislativa/plenario/discursos/escrevendohistoria/25-anos-da-constituicao-de-1988/momentos-marcantes-da-constituente>.

when the government supported the proposal and aligned when the government opposed it. The deputy–topic alignment score $a_{ij} \in [0, 1]$ ranges from 0 (complete opposition) to 1 (full alignment). We aggregate these into three deputy-level indices: the *Loyalty Index* L_i (mean of a_{ij} across all seven topics), the *Institutional Index* I_i (mean across the two institutional topics), and the *Ideological Index* D_i (mean across the five ideological topics). These indices are the dependent variables in the roll-call regressions of Section 5.¹³

Electoral outcomes. To measure the post-NCA political persistence of concession recipients, we use TSE records of every federal, state, and municipal election from 1990 onward. The data cover candidate names, offices contested, parties, vote counts, and ballot outcomes. We link NCA candidates to subsequent elections using name-state-party matching analogous to the concession-matching procedure, and we use the resulting deputy-level panel to construct the incumbency-persistence outcomes of Section 5.

5 Empirical Analysis

5.1 Allocation: close-election RDD on concession receipt

Table A9 in Appendix A reports summary statistics for the main estimation samples.

Descriptive Evidence. We begin by documenting broad descriptive patterns in the evolution of media concessions during the democratic transition. Figure 1 plots the cumulative number of radio and television licenses from 1979 to 1990 using administrative records from the Ministry of Communications. The figure reveals two patterns. First, television concessions expanded rapidly in the early 1980s—particularly between 1979

¹³The Chamber’s published proceedings are distributed as scanned PDFs. We extract the roll-call matrix via OCR, validate each extracted vote manually against the original PDF, and classify each sub-vote by topic. Appendix C describes the extraction pipeline, the validation procedure, and the topic-level sub-vote assignments in detail.

and 1982—but their growth slowed substantially thereafter, including during the National Constituent Assembly (NCA). Second, and in sharp contrast, radio concessions exhibit a pronounced surge precisely in 1987–88, coinciding with the NCA period. This divergence reflects that radio stations were significantly cheaper to install and operate and had a more localized reach, making them a more flexible political resource than TV stations, whose expansion had largely stabilized before the NCA.

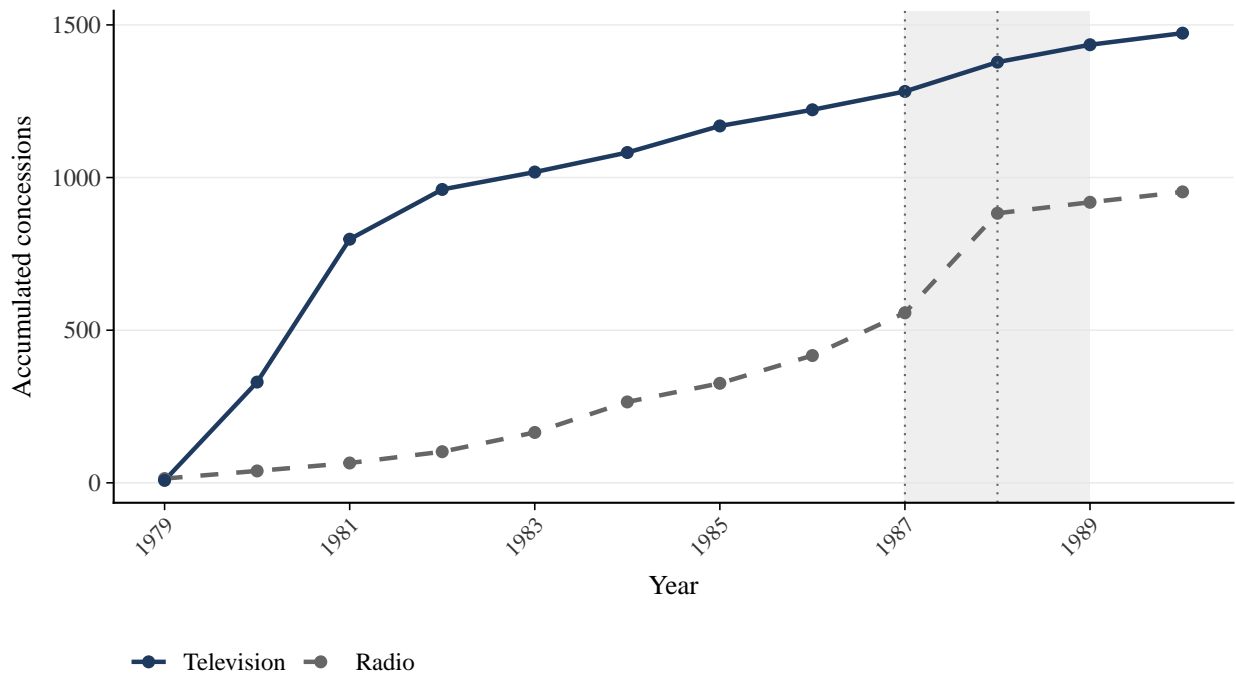


Figure 1: Accumulated TV and Radio Concessions, 1979–1990

Notes: Cumulative radio and television concessions from LAI administrative records of the Ministry of Communications. Vertical dashed lines mark the 1987–88 Constituent Assembly.

Figure 2 compares the evolution of media concessions for candidates who won or lost the 1986 election. While both groups experienced a gradual increase over the decade, a clear divergence emerges in 1987–88, with elected deputies receiving substantially more concessions during the NCA bargaining window—consistent with targeted allocation in exchange for legislative support. The analysis that follows formally tests this pattern and its downstream consequences for voting behavior and electoral persistence.

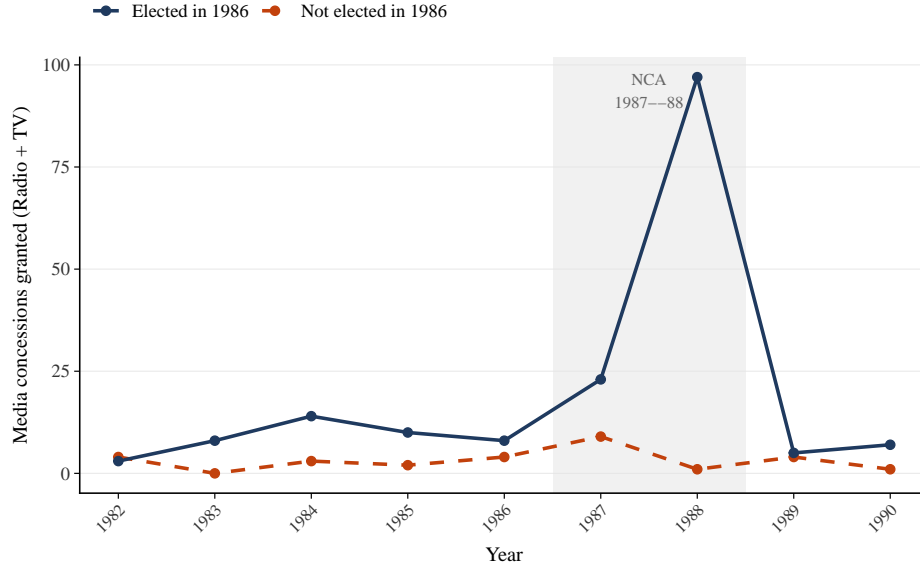


Figure 2: Media Concessions Granted Over Time, by 1986 Electoral Outcome

Notes: Annual count of radio and television concessions received by candidates who were elected (*Elected in 1986*) versus those who ran but were not elected (*Not elected in 1986*) in the NCA election. The shaded band marks the National Constituent Assembly bargaining window (1987–88). Data from LAI requests to the Ministry of Communications, matched to the 1986 federal candidate registry.

Effect of Electoral Victory on Media Concessions. To identify the effect of being elected as a National Constituent Assembly (NCA) deputy, we leverage quasi-random variation from close electoral races within Brazil’s open-list proportional representation system. Our running variable follows Novaes (2024), comparing candidates within the same party-coalition and state according to their distance to the effective electoral threshold.¹⁴ This design allows us to compare candidates who narrowly secured a seat with those who narrowly missed, under the assumption that these individuals are otherwise similar with respect to relevant political and personal characteristics.

The design is well-powered and the sample structure is as expected under open-list proportional representation. The RD population consists of 2,614 candidates for whom

¹⁴Let V_{ics} denote the vote total of candidate i running on coalition c in state s , and let V_{cs}^* denote the vote total of the last elected candidate from coalition c in state s (the coalition’s effective seat threshold). The running variable is $x_{ics} = (V_{ics} - V_{cs}^*)/V_{cs}^*$, so that $x_{ics} > 0$ for elected candidates and $x_{ics} < 0$ for defeated ones; the cutoff is $x = 0$.

the running variable is well-defined. Appendix Figure A4 displays the distribution of the running variable, showing a concentration of mass near the cutoff and a rightward skew reflecting the larger number of candidates who narrowly lost relative to those who narrowly won. At the MSE-optimal bandwidth of approximately 1.4, roughly 76% of the RD population falls within the estimation window, with 1,622 observations to the left and 359 to the right. The asymmetry reflects the structure of open-list proportional representation: each coalition produces many marginal losers but few marginal winners.¹⁵

A clear discontinuity emerges at the electoral threshold. Figure 3 plots binned averages of the probability of receiving a media concession against the vote margin relative to the coalition threshold, with local-linear fits on each side of the cutoff. Candidates who barely won were substantially more likely to receive a concession than those who barely lost, confirming that concession allocation was systematically related to electoral success.

Panel A of Table 1 formalizes this result. Narrowly elected candidates were 13.6 percentage points more likely to obtain a media concession (archival-augmented coding; automated match alone yields 6.0 pp), with bias-corrected p -values below one percent for composite and radio outcomes. Donut specifications that exclude observations within 0.10 and 0.25 of the cutoff yield estimates of 14.3 and 10.5 pp, flat or rising as the donut widens and inconsistent with strategic manipulation. The marginal-pair RDD yields +5.5 pp (Appendix Table D7), well within the main confidence interval.

Disaggregating the results by concession type reveals a pattern consistent with the institutional context described earlier. The RD effects are driven overwhelmingly by radio concessions, which expanded sharply during 1987–1988. Television concessions, which had mostly been allocated earlier in the decade, show smaller and less precise effects. This

¹⁵With the symmetric MSE-optimal bandwidth, 359 of 487 right-side observations (74%) fall within $h^* = 1.44$. To verify the algorithm is not destabilized by this thin right side, we re-estimate using `msetwo`, which selects separate optimal bandwidths for each side. `msetwo` sets $h_l = 1.55$ and $h_r = 1.63$, widening the right-side window and adding nine additional right-side observations; the point estimate is unchanged at +13.5 pp ($p < 0.001$).

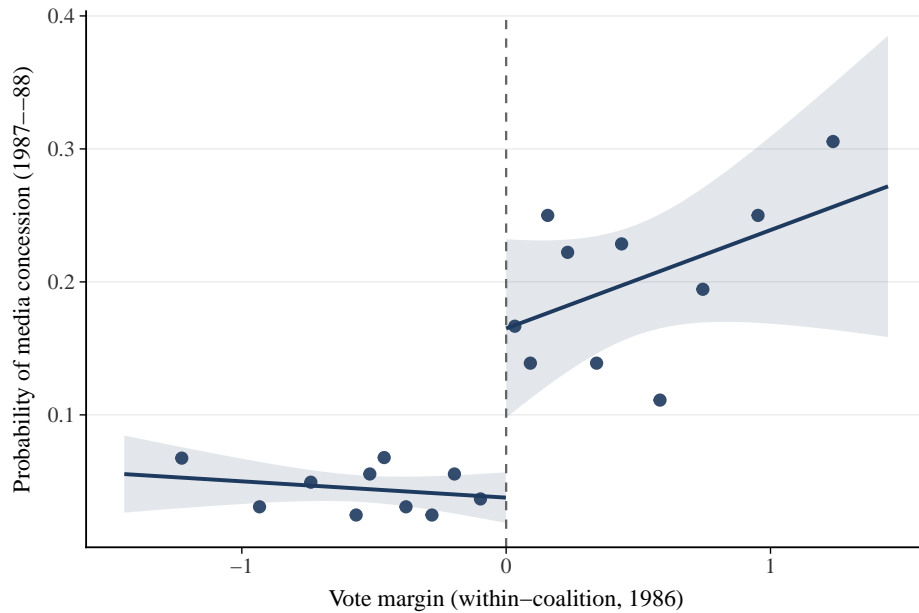


Figure 3: RD plot: probability of receiving a media concession by 1986 vote margin

Notes: Binned averages with local-linear fits on each side of the cutoff. The window shown extends to $1.4\times$ the MSE-optimal bandwidth used for inference (Panel A of Table 1); the wider visualization window is a presentation choice that lets the reader see the local trend on each side without changing the inferential bandwidth, following the visualization conventions of Calonico et al. (2014).

pattern is consistent with historical accounts (Motter et al., 2019) documenting how radio licenses—being cheaper, more numerous, and more locally influential—became the primary instrument of political exchange during the constitution-writing period.

Balance tests for predetermined covariates are in Appendix A.¹⁶

Appendix Table A2 shows that, with the exception of PFL membership (+14.4 pp, $p = 0.002$), observable characteristics evolve smoothly at the cutoff. The PFL discontinuity reflects a mechanical feature of open-list proportional representation discussed in the covariate-smoothness paragraph below; the covariate-adjusted RDD addresses it directly. As an additional robustness check, Appendix Table D1 confirms that the main RD

¹⁶Covariates are restricted to 1986 and pre-1986 records: post-1986 TSE filings are observable only for candidates who ran again after 1986, a group directly shaped by the electoral outcome, which would introduce the imbalance the test is designed to detect. Ideological positions use party-level scores from Power and Zucco Jr (2009) assigned by 1986 party affiliation.

estimates are virtually identical when restricting the sample to candidates who had not previously held elected office.

The estimates are stable across bandwidth choices and alternative running-variable constructions (Appendix Figure A6 and Table D9).

Manipulation of the running variable would require coordinated reordering of candidates on open lists in response to anticipated vote totals—implausible in the 1986 context. The local-polynomial density test of Cattaneo et al. (2020) finds no evidence of strategic sorting in the PMDB-only sample ($T = 1.21$, $p = 0.227$); apparent rejections for the governing-coalition and right-opposition subsamples reflect a structural feature of open-list PR in which coalitions mechanically produce more marginal losers than winners, not strategic manipulation—donut specifications restore the pooled p -value to 0.89 (Appendix Table A1 and Figure A5).¹⁷

Covariate smoothness at the 1986 cutoff. Appendix Table A2 reports the RDD covariate-smoothness test. Six of seven covariates pass cleanly. PFL membership shows a residual discontinuity (+14.4 pp, $p = 0.002$), reflecting a mechanical feature of open-list proportional representation in which PFL accumulated proportionally more near-margin losers than winners relative to PMDB.¹⁸ The covariate-adjusted RDD (Table 1) enters PFL and prior ownership directly and leaves the substantive conclusions unchanged.

Covariate-adjusted RDD. Panel B of Table 1 reports the covariate-adjusted result. Point estimates fall by roughly one fifth across all three outcomes (any media: 13.6 \rightarrow 10.3 pp; radio: 13.1 \rightarrow 10.8 pp; TV: 4.8 \rightarrow 4.1 pp)—the magnitude implied by the observed balance

¹⁷Appendix Table D6 confirms the marginal-pair sample passes cleanly ($T = 0.02$, $p = 0.98$ within PMDB). Appendix Table D2 reports full donut-robustness by party group.

¹⁸Appendix Figure A7 plots the loser-to-winner ratio within the bandwidth across all bandwidth choices for PMDB and PFL separately. The PFL ratio (≈ 2.93 at the MSE-optimal bandwidth) is 1.8 \times higher than PMDB's (≈ 1.64) and stable across bandwidth choices; PMDB's ratio is also stable and close to the pooled sample average. The imbalance is therefore a structural feature of list allocation under open-list proportional representation, not a sign of manipulation.

gaps—and all three remain significant at the one-percent level for radio and any-media. Pre-treatment controls do not affect the core discontinuity.

Table 1: RDD Estimates with and without Pre-Treatment Covariate Adjustment

	TV or Radio	Radio	TV
<i>Panel A: Baseline (no covariates)</i>			
RD estimator	0.136*** (0.036)	0.131*** (0.035)	0.048** (0.023)
Robust 95% CI	[0.059, 0.216]	[0.058, 0.211]	[0.001, 0.100]
Bandwidth (MSE)	1.44	1.23	1.65
Eff. obs. (left / right)	1622 / 359	1541 / 343	1695 / 368
% of RD population in BW	75.8%	72.1%	78.9%
<i>Panel B: Calonico-Cattaneo-Farrell-Titiunik (2019) covariate-adjusted</i>			
RD estimator	0.103*** (0.032)	0.108*** (0.032)	0.041** (0.021)
Robust 95% CI	[0.033, 0.170]	[0.041, 0.181]	[0.000, 0.090]
Bandwidth (MSE)	1.41	1.20	1.30
Eff. obs. (left / right)	1614 / 355	1537 / 340	1562 / 348
% of RD population in BW	75.3%	71.8%	73.1%
RD population	2,614		

Notes: Panel A reports the baseline linear local-polynomial RDD without covariate adjustment. Panel B adds pre-determined covariates additively using the estimator of Calonico et al. (2019). The covariate set is 1982 office-holding, 1982 electoral participation, pre-1986 ownership of any media concession and separately of radio and TV concessions, and 1986 PMDB and PFL dummies. All specifications use a triangular kernel with the MSE-optimal bandwidth computed separately for each outcome and specification, and nearest-neighbor variance estimation. Eff. obs. (left / right) are the effective observations inside the bandwidth on each side of the cutoff; % of RD population in BW is the fraction of the 2,614-candidate RD population that lies inside the estimation window. Conventional standard errors in parentheses; robust 95% confidence intervals bias-corrected following Calonico et al. (2014). * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

A complementary placebo test asks whether the discontinuity is specific to the NCA bargaining window or whether it instead reflects a generic incumbency channel through which 1986 winners receive concessions in any period. Appendix Figure A2 estimates the same close-election RDD on concessions granted in four distinct time blocks: 1983–1986 (pre-NCA), 1987–1988 (NCA), 1989–1990 (post-NCA), and 1991–1994 (well after the NCA). The discontinuity is sharply concentrated in 1987–1988: the RDD coefficient

of 13.6 pp is roughly three times larger than the pre-NCA and post-NCA estimates (both around 4.5 pp and individually insignificant at conventional levels), and the 1991–1994 estimate is precisely zero. Appendix Table A10 and Appendix Figure A8 extend the same logic to the close-election dimension by re-running the design for every federal Chamber-of-Deputies election from 1982 to 1998, defining the outcome as concession receipt during the subsequent legislative term. The discontinuity reproduces only in the 1986 row; the 1982, 1990, 1994, and 1998 elections deliver null or slightly negative estimates. The 1987–88 pattern is specific to the NCA context, not a generic feature of Brazilian close elections in this period.

A central prediction of our theoretical model presented in Section 3 is that the Executive should strategically target legislators for whom support is “cheaper” in ideological terms. That is, ideologically aligned deputies should be more responsive to media concessions, and thus more likely to receive them, whereas distant ideological opponents should require transfers so large that offering them concessions becomes suboptimal. To examine this theoretical prediction, we estimate heterogeneous RD effects by political alignment. Table 2 reports separate estimates for government supporters, the left opposition, and the right opposition. The results closely follow the model’s predictions. Government-aligned legislators exhibit large and precisely estimated discontinuities, with narrowly elected supporters of the PMDB-led governing coalition being 18–21 percentage points more likely to receive a concession than narrowly defeated coalition candidates. In contrast, left-opposition candidates display no meaningful discontinuity, with point estimates close to zero across all concession types. This is exactly what the model predicts: when ideological distance is large, the compensating transfer required to secure support becomes prohibitively high, making these legislators unattractive targets.¹⁹

For the right opposition, ideologically closer to the Executive than the left but still

¹⁹Appendix Table D4 shows that these heterogeneous effects do not change significantly when restricting the sample to candidates without prior electoral experience, reinforcing that the alignment patterns are not driven by previous incumbency.

outside the governing bloc, the estimated effects are positive but smaller and statistically weaker. This intermediate pattern is consistent with the model’s comparative statics: as ideological distance increases, the optimal transfer declines smoothly and eventually becomes zero.²⁰

Taken together, the RD estimates confirm that concession allocation responded systematically to electoral outcomes and ideological proximity—precisely the pattern the model predicts for a productive political asset targeted at pivotal legislators.

Table 2: RDD Estimates by Political Alignment (Linear Specification)

	TV or Radio	Radio	TV
Panel A: Government-Aligned Legislators			
RD Estimator	0.206** (0.051)	0.182** (0.047)	0.090* (0.034)
Eff. number obs.	1015	1033	1045
CCT-optimal BW	0.979	1.010	1.038
Panel B: Left Opposition			
RD Estimator	-0.040 (0.033)	-0.036 (0.028)	-0.007 (0.015)
Eff. number obs.	334	251	331
CCT-optimal BW	0.564	0.481	0.543
Panel C: Right Opposition			
RD Estimator	0.098 (0.062)	0.099 (0.065)	0.038 (0.049)
Eff. number obs.	483	553	599
CCT-optimal BW	0.870	1.094	1.124

Notes: All estimates from `rdrobust` using a linear specification ($p = 1$). Standard errors in parentheses. Asterisks denote significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Effective number of observations and CCT-optimal bandwidths reported.

²⁰A formal test of this monotone gradient would require party-level RD estimates across a sufficient number of parties. The Power and Zucco Jr scores used for balance tests are party-level, and only PMDB and PFL have enough marginal winners ($N \geq 100$) to support reliable RD inference. Estimating separate RD effects for the four additional parties with at least 15 marginal winners yields directionally consistent patterns—PT (Power score 1.06) shows a negative point estimate and PFL (Power score 7.47) the largest positive effect—but the coefficients are too imprecise to constitute a formal monotonicity test (four of six estimates are statistically indistinguishable from zero). We therefore rely on the three-group alignment split as the empirically tractable operationalization of the model’s comparative statics.

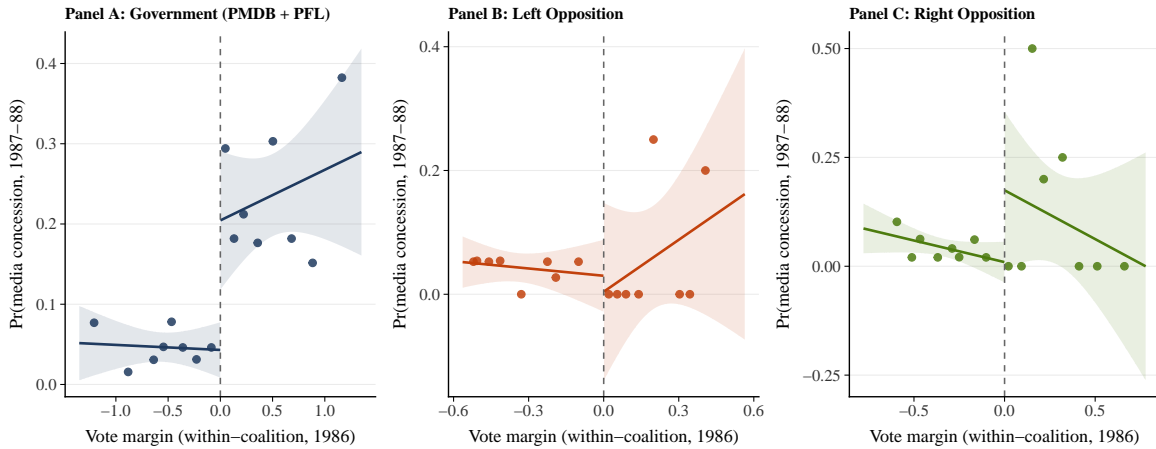


Figure 4: RD plots: probability of receiving a media concession by 1986 vote margin, by political alignment

Notes: Each panel shows binned averages of the probability of receiving a media concession in 1987–88 against the within-coalition vote margin, with local-linear fits on each side of the cutoff. Panel A restricts to government-aligned legislators (PMDB and PFL, Sarney’s governing coalition); Panel B to left-opposition parties (PDT, PT, PSB, PCB, PCDOB); Panel C to the right bloc (non-PMDB/PFL non-left parties). RD estimates and bandwidths match Table 2.

5.2 Legislative Response: within-winner roll-call associations

Having established the allocation margin, we now ask whether concession recipients voted differently in the National Constituent Assembly. Because only elected candidates cast NCA votes, the RD design cannot be applied to roll-call outcomes; the evidence below is therefore best read as *within-winner conditional associations*, complementary to the causal allocation result.

With that qualification in mind, the NCA setting provides a natural two-way identification of issue-differentiated behavior. The Sarney administration held a unified, publicly stated position on two institutional questions—the presidential system of government and the five-year presidential term—while legislative blocs split along pre-existing ideological lines on five topics: agrarian reform, right to strike, job stability, work hours, and property rights (Pilatti, 2008). This two-class structure maps directly onto the Institutional and Ideological indices estimated below.

We rely on the roll-call dataset described in Section 4 and Appendix C. For each of the 451 deputies and each topic, we measure the proportion of votes aligned with the government’s publicly stated preferences.²¹ We estimate within-winner partial correlations via

$$Y_i = \alpha + \beta \cdot \text{Concession}_i + \delta_{sp(i)} + \gamma x_i + \varepsilon_i, \quad (9)$$

where $Y_i \in \{L_i, I_i, D_i\}$, $\delta_{sp(i)}$ denotes state-by-party fixed effects, and x_i includes the 1986 vote margin and pre-1987 concession holdings. These controls absorb persistent ideological, regional, and pre-existing-ownership heterogeneity; residual selection concerns are addressed in the sensitivity analysis below. We interpret the patterns that follow as *consistent with* the quid-pro-quo mechanism.

Table 3 reports the results. Panel A shows that the pooled Loyalty Index is large and stable: concession recipients voted with the Executive on 12.8 to 15.0 percentage points more of the constitutional agenda than non-recipients (all $p < 0.001$). Panel B shows the key contrast. For ideological items, partial correlations are small and statistically indistinguishable from zero once state-by-party fixed effects are introduced (-3.7 pp in Column 2, $p = 0.26$).²² For institutional items, large positive associations emerge: the Institutional Index coefficient is $+0.296$ ($p < 0.001$) in Column 2, with deputies who received a concession being 24.3 pp more likely to vote for the presidential system and 34.8 pp more likely to support the five-year term. The implied Institutional-minus-Ideological gap ranges from $+33.3$ pp to $+43.0$ pp, statistically significant at the one-percent level in every col-

²¹Of the 487 candidates who won seats in the 1986 NCA election, 475 are linked to the roll-call records; the remaining 12 could not be matched to the voting archive (deaths during the Assembly, mid-term substitutions, or unresolved name-disambiguation). Of these 475, a further 24 are dropped as singletons when state-by-party fixed effects are absorbed, leaving the estimation sample of 451.

²²Job Stability and Right to Strike are apparent exceptions at the topic level: their individual coefficients are $+0.220$ ($p = 0.010$) and -0.115 ($p = 0.018$) in Column 2 before multiplicity adjustment. Job Stability’s positive sign reflects that the government supported moderate provisions and opposed only the most rigid dismissal constraints (see Appendix C); concession recipients were thus marginally more likely to support the government’s centrist position. Right to Strike’s negative sign reflects that the government opposed the most expansive strike guarantees. Neither coefficient survives Romano–Wolf correction (RW $p = 0.147$ for both), illustrating why individual topic coefficients without multiplicity adjustment can be misleading in multi-outcome analyses.

Table 3: Pooled Within-Winner Concession Effects: Loyalty Index and Ideological/Institutional Components

	(1) OLS	(2) + State-party FE	(3) + Pre-1987 media controls	RW adj. <i>p</i> (FWER, Col. 3)
<i>Panel A: Loyalty Index (mean of 7 topic supports per deputy)</i>				
Concession (1987–88)	0.150*** (0.024)	0.130*** (0.027)	0.128*** (0.029)	–
Observations	475	451	451	
<i>Panel B: Heterogeneous Effects on Divisive and Non-Divisive Issues</i>				
Non-Divisive Index (institutional votes)	0.363*** (0.039)	0.296*** (0.048)	0.305*** (0.053)	–
Presidential System	0.296*** (0.049)	0.243*** (0.047)	0.248*** (0.052)	0.002***
Five-Year Term	0.431*** (0.048)	0.348*** (0.061)	0.362*** (0.067)	<0.001***
Divisive Index (ideological votes)	-0.062** (0.024)	-0.037 (0.033)	-0.038 (0.034)	–
Agrarian Reform	-0.106*** (0.029)	-0.075** (0.031)	-0.093*** (0.028)	0.032**
Right to Strike	-0.080* (0.041)	-0.115** (0.043)	-0.115** (0.046)	0.147
Job Stability	0.319*** (0.040)	0.220** (0.085)	0.197** (0.083)	0.147
Work Hours	-0.028 (0.035)	-0.043 (0.040)	-0.042 (0.043)	0.632
Property Rights	-0.100* (0.054)	-0.031 (0.065)	-0.032 (0.066)	0.732
Observations (non-divisive / divisive)	475 / 431	451 / 406	451 / 406	

Notes: Deputy-level regressions ($n = 475$ in Column 1; $n = 451$ in Columns 2–3 as singleton state-party cells are dropped when fixed effects are absorbed). Panel A regresses the Loyalty Index—the mean support rate across the seven constitutional topics—on an indicator for having received a TV or radio concession in 1987–88. Panel B regresses the Institutional Index (mean of presidential system and five-year-term supports) and the Ideological Index (mean of agrarian reform, right to strike, job stability, work hours, and property rights) on the same indicator; indented rows show the coefficient from the same specification run separately on each constituent topic. Column 1 is OLS with no fixed effects; Column 2 adds state-by-party fixed effects; Column 3 further controls for the 1986 running variable and for the number of TV and radio concessions held in 1982–1986. The “RW adj. *p*” column reports Romano–Wolf (2005) step-down family-wise error rate (FWER) adjusted *p*-values for the seven per-topic coefficients from Column 3, obtained via wild-cluster bootstrap ($B = 4,999$ Rademacher draws, clusters = state-by-party cells). RW adjustment is applied jointly to all seven topics; index rows (Panel A, Institutional and Ideological indices) are not subject to this correction and show “–”. Standard errors clustered at the state-by-party level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$ for both conventional and RW-adjusted columns. The Institutional-minus-Ideological gap reported in the text (+33 to +43 pp across columns) is estimated directly as the coefficient on Concession $\times \mathbb{1}\{\text{Institutional}\}$ in a stacked deputy–index regression, not as the arithmetic difference between the two index rows; the +45.4 pp figure in the text corresponds to the same interaction estimated with deputy fixed effects absorbed (Appendix Table A11).

umn. Because the interaction term is identified from the within-deputy contrast between each deputy's two index values, it is unaffected by deputy-level selection into concessions. Romano–Wolf FWER adjustment confirms: only Presidential System (RW $p = 0.002$) and Five-Year Term (RW $p < 0.001$) survive at the one-percent level; all ideological items fail to clear either threshold (RW $p \geq 0.147$). The two-coding sensitivity in Appendix Table A3 shows the within-deputy differential is stable under both codings (+0.380 vs. +0.454, each significant at one percent).

Sensitivity to selection on unobservables. The within-deputy contrast differences out all deputy-level selection by construction; robustness to issue-by-deputy-level confounders is documented in Appendix Table A11 (Cinelli–Hazlett $RV_{q=1} = 0.246$).

Table 4 decomposes the within-winner associations by political alignment. Government-coalition deputies (PMDB+PFL, $N = 366$, 93 concession recipients) display the largest and most precisely estimated differential: the Institutional minus Ideological gap is +40.9 percentage points ($p < 0.001$). Left-opposition deputies ($N = 47$) received only two concessions under the broad ownership coding—both via indirect family holdings—too few to support regression analysis; their near-absence from the concession pool confirms that the Executive did not direct this instrument toward ideological opponents. Right-opposition deputies ($N = 62$, 12 recipients) show a positive but statistically imprecise differential of +15.6 pp ($p = 0.307$). The productive-capital channel thus operated most sharply within the government coalition, where the alignment between concession holders and institutional vote targets was tightest.²³

Taken together, the results are consistent with the prediction developed in Section 3 that media concessions operated as targeted transfers associated with differential align-

²³Appendix Table A7 re-estimates the four Loyalty-Index-family coefficients separately for PMDB and PFL. The Institutional-minus-Ideological differential is positive and significant in both parties (+37.8 pp for PMDB, $p = 0.001$; +16.4 pp for PFL, $p = 0.010$), with the PMDB estimate roughly twice as large, consistent with PMDB legislators spanning a wider ideological range on the institutional items.

Table 4: Within-Winner Concession Effects by Political Alignment Group

	Government (PMDB+PFL)	Right Opposition
<i>Panel A: Loyalty Index</i>		
Concession	0.103*** (0.030)	-0.024 (0.087)
N deputies	366	62
N concession recipients	93	12
<i>Panel B: Non-Divisive and Divisive Indices</i>		
Non-Divisive Index	0.238*** (0.058)	0.043 (0.215)
Divisive Index	-0.041 (0.028)	0.021 (0.080)
Non-Div – Divisive	0.409*** (0.067)	0.156 (0.150)
State-party FE		Yes
Pre-1987 media controls		Yes

Notes: Specification matches Column 3 of Table 3: OLS with state-by-party fixed effects, controlling for the 1986 running variable and pre-1987 media holdings. Groups: Government = PMDB+PFL (Sarney’s governing coalition); Right Opposition = other non-left parties. Left-opposition deputies ($N = 47$) received 2 concessions under the broad ownership coding (indirect family holdings)—too few for regression—and are excluded. Standard errors clustered at the state-by-party level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

ment among pivotal legislators.²⁴ We find no systematic relationship between concessions and votes on issues where ideological positions were steep and costly to shift. Instead, the associations concentrate on decisions that were especially valuable to the Executive and for which deputies closer to the government were more likely to align at lower ideological cost.

²⁴A natural question is whether one could test the productive-vs-consumptive distinction by running the same close-election RDD on a fiscal-pork analog. The relevant institutional instrument did not yet exist: individual budget amendments (*emendas parlamentares individuais*) were created by the 1988 Constitution itself and first executed in the 1989 federal budget (Limongi and Figueiredo, 1998; Pereira and Mueller, 2002; Pereira and Mueller, 2004). Before 1989, deputy-level fiscal benefits flowed through collective bench amendments and informal channels that cannot be linked to individual legislators in the way a discretionary broadcast license can—making broadcast concessions one of the few legibly individual-level resources the Executive controlled during the NCA bargaining episode.

Having shown that concession receipt predicts differential alignment on institutional—but not ideological—items, we examine whether the asset translated into durable electoral advantage after the NCA.

5.3 Electoral Advantage

The close-election RDD bundles incumbency and concession receipt into a single treatment and cannot separate the two channels; concession receipt among the elected is endogenous to electoral strength. The exercises below are therefore descriptive benchmarks.

Appendix Figure A3 splits marginal 1986 winners into concession recipients ($N = 80$) and non-recipients ($N = 407$), using barely-lost candidates as the shared counterfactual. Because concession receipt among winners is endogenous, the gap between the two series is illustrative, not a causal estimate.

The no-concession series peaks at +35.0 pp elected in 1990 and decays to +4.3 pp by 2002, tracing the finite-horizon incumbency premium. The concession-recipient series is larger in every year—+38.9 pp in 1990, +22.6 pp in 1998, +20.5 pp in 2002—roughly four to five times the no-concession estimate in the later years, and decays more slowly, consistent with the broadcasting asset extending electoral durability beyond the mandate.²⁵

5.4 Heterogeneous Persistence

Productive political capital: heterogeneous persistence. We now turn to the third piece of evidence: the pattern of post-NCA persistence that the productive-political-capital framework predicts. The framework distinguishes broadcast concessions from fiscal transfers along two dimensions. A license is *discrete and identity-embodied*—the asset stays with its holder rather than being consumed—and its electoral value is *market-*

²⁵The concession-recipient series uses only 43 right-side observations ($SE \approx 15$ pp); concession receipt is also endogenous to unobserved political strength, so these estimates should not be read causally.

structure-dependent—a single new station dominates information in a thin market and contributes little in a saturated one. The market-structure dimension predicts that the post-NCA incumbency premium should concentrate in non-competitive media markets (the market-competitiveness heterogeneity estimated below).

The exercise identifies the heterogeneity of the close-election incumbency premium by pre-NCA media-market competitiveness.

State-level heterogeneity by media market competitiveness. This heterogeneity cut asks whether the post-NCA incumbency premium varies with the local information environment in the way the productive-political-capital model predicts. A station in a market with many competitors adds little to the politician’s marginal visibility, while one in a thin market may dominate the local information environment. We split the sample by pre-NCA media-market competitiveness—measured as the number of *Fistel* broadcasting registrations per municipality in 1985, divided at the state-level median—and re-estimate the RDD separately for candidates from competitive (above-median, $N = 1,978$) and non-competitive (below-median, $N = 636$) states.²⁶ The non-competitive-minus-competitive difference is estimated from a pooled local-linear regression on the union bandwidth as the coefficient on $\mathbb{1}\{\text{running} \geq 0\} \times \mathbb{1}\{\text{non-competitive}\}$, oriented so that positive values

²⁶We measure competitiveness at the state level rather than the municipality level for two reasons. First, federal deputies compete in open-list proportional elections held statewide; no single municipality constitutes a deputy’s electoral market. Second, municipality-level vote shares from the TSE are available only from 1998 onward, precluding the use of pre-NCA vote distributions to assign candidates to a sub-state market. The state-level density—number of *Fistel* registrations per municipality, averaged within a state—is therefore the finest geographic unit at which a competitiveness index can be assigned consistently to all candidates in the RDD sample. Within-state heterogeneity in market structure would tend to attenuate the measured non-competitive–competitive gap; our estimates should be read as a lower bound on the true heterogeneity. The continuous-interactor specification in Appendix Table A4, which uses the full state-level *Fistel* density rather than a binary split, is directionally consistent. Regarding registry composition: the *Fistel* database through 1985 contains 2,574 registrations—363 radio (217 FM, 146 AM) and 2,211 television, of which 2,089 (81%) are retransmission stations (RTV) that passively relay a generator’s signal without independent editorial operation, and 35 are generators (TV/GTVD). Our competitiveness index therefore captures the saturation of the local broadcasting environment—including relay infrastructure—rather than the count of independent editorial voices. This is the appropriate measure for the mechanism: a market saturated with relay stations already delivers political information to viewers, reducing the marginal informational reach of any new concession.

indicate a larger incumbency effect in non-competitive markets. The sample is cut at 2002 because the share of NCA deputies retaining their broadcasting concession drops sharply after that date, so post-2002 coefficients conflate persistence of the political effect with attrition of the underlying ownership base.

Figure 5 reports the results. In 1990, both market types produce large and comparable incumbency effects on both outcomes, with the competitive states marginally higher (+30.2 pp elected, +45.2 pp ran, both $p < 0.001$) than the non-competitive ones (+23.0 pp, +35.8 pp, both significant). The two trajectories diverge starting in 1994. In non-competitive markets the 1986 incumbency premium *persists* for a second cycle: marginal 1986 winners are +20.5 pp more likely to be elected as federal deputies in 1994 ($p = 0.009$) and +26.4 pp more likely to run ($p = 0.004$). In competitive markets the same winners have no detectable advantage in 1994 on either margin (+6.2 pp elected, $p = 0.10$; +6.1 pp ran, $p = 0.20$). The non-competitive-minus-competitive differential makes this contrast precise: +13.9 pp for elected ($p = 0.099$) and +17.2 pp for ran ($p = 0.078$), both marginally significant on either side of the ten-percent level. By 1998 the non-competitive margin remains higher but with wider confidence intervals, and by 2002 both market types have faded to small positive effects that cannot be distinguished from one another.

Appendix Tables A5 and A6 confirm robustness to donut specifications, covariate adjustment, and bias-corrected inference; power and MDE calculations are in the latter.²⁷

The pattern confirms the framework’s prediction: a station in a thin market dominates the local information environment and translates into a second-cycle incumbency

²⁷Aggregating post-mandate outcomes into a single binary—“ever elected in at least one of 1994, 1998, or 2002”—concentrates all identifying variation into one test and eliminates multiple-testing dilution. The pooled RDD yields a non-competitive effect of +27.1 pp (SE = 8.6, $p = 0.005$) and a competitive effect of +12.9 pp (SE = 4.4, $p = 0.011$), for a difference of +13.9 pp (SE = 10.3, $p = 0.18$). The point estimate is economically large—one third of the non-competitive control mean—but imprecisely estimated because the non-competitive subsample contains only 636 candidates compared to 1,978 in the competitive group, and the aggregate bandwidth places only ~350 non-competitive candidates within the estimation window. The power constraint is therefore structural: the 1:3 sample split between market types limits the precision of any interaction regardless of how outcomes are aggregated.

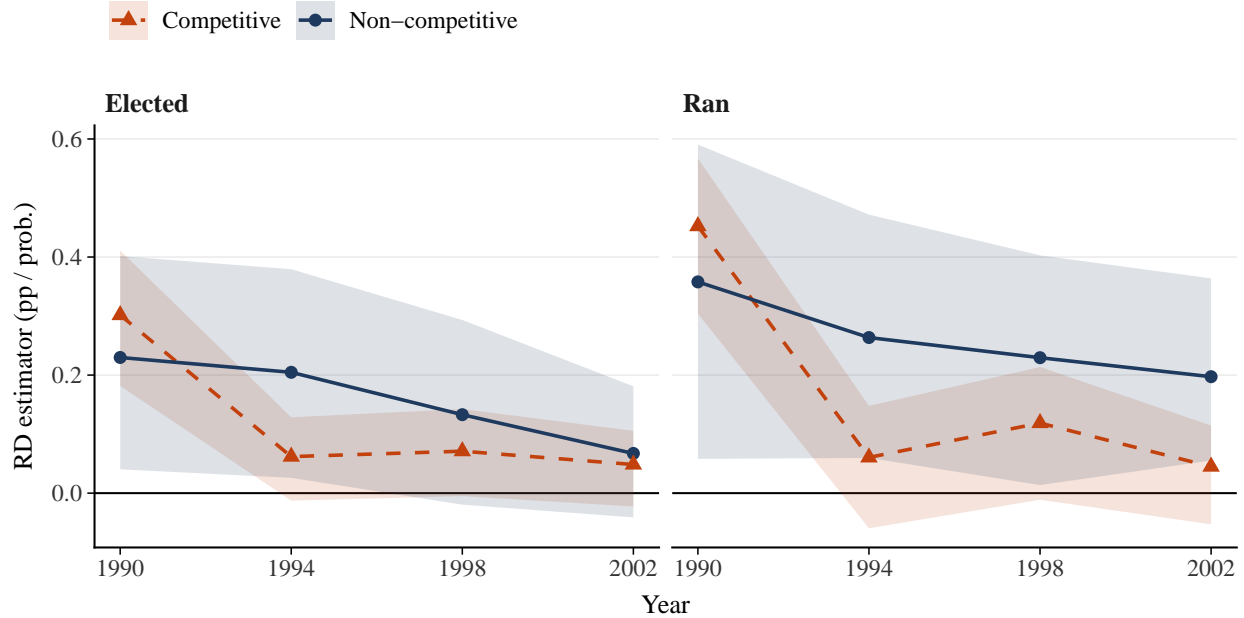


Figure 5: RDD Incumbency Effects by Media Market Competitiveness

Notes: Local-linear RDD estimates of the effect of barely winning the 1986 NCA election on probability of being elected (left panel) or running (right panel) as a federal deputy in each subsequent election year, for candidates from states with below-median (non-competitive, $N = 636$) and above-median (competitive, $N = 1,978$) pre-NCA Fistel broadcasting density per municipality in 1985. Triangular kernel with MSE-optimal bandwidth selected within each subsample. Bands are 95% confidence intervals. Standard errors clustered at the state-party level.

premium; in a saturated market, the marginal informational contribution is small and the advantage dissipates within one election.

Synthesis: three pieces and the productive-political-capital reading. The empirical evidence assembles three pieces of the productive-political-capital framework. *Allocation:* the close-election RDD identifies a +13.6 pp effect of barely winning the 1986 NCA seat on receipt of a 1987–88 broadcasting concession, with placebo cycles null. *Legislative response:* within-deputy contrast isolates a +33 to +43 pp gap across pooled specifications, rising to +45.4 pp in the within-deputy panel, concentrating on institutional items and absent on ideological ones. *Heterogeneous persistence:* the post-NCA incumbency premium concentrates in non-competitive media markets and is absent in competitive ones (non-

competitive-minus-competitive of +13.9 pp on elected-1994; robustness in Appendix Tables A4 and A6). The three pieces confirm the framework’s core predictions.

6 Conclusion

This paper studies how control over broadcast media is allocated, retained, and translated into legislative behavior during a foundational moment of institutional redesign. Using a new dataset linking nearly universal information on radio and television concessions to individual politicians, combined with close-election RDD designs and a simple conceptual framework, we present three pieces of evidence that together characterize broadcast concessions as productive political capital—an asset that is discrete, identity-embodied, and market-structure-dependent.

On *allocation*, barely elected NCA members were 13.6 percentage points more likely to receive a concession in 1987–88, with larger effects among government-aligned legislators and null effects in every placebo election. On *legislative response*, concession recipients show large within-deputy shifts in government alignment concentrated on institutional but not ideological votes—a contrast identified from each deputy’s own voting record, which differences out deputy-level selection by construction. On *heterogeneous persistence*, electoral returns to the 1986 margin survive a full cycle only where local media markets are thin; in competitive markets the advantage dissipates after the first election. This market-structure conditionality is the sharpest mechanism test the paper delivers: a consumptive transfer generates no such conditional persistence, while a productive informational asset generates lasting returns precisely where it dominates the information landscape. Identification limits are detailed in Appendix A.

These findings contribute to several strands of research. First, they expand the media-politics literature by documenting not only how media influences voters, but how control

over media is allocated in the first place, and how this allocation interacts with local market structure to shape post-allocation political competition. Existing work emphasizes biased coverage or media entry; we document how a politically allocated broadcast license operates as a discrete, identity-embodied asset whose post-allocation electoral value is conditioned by the structure of the market into which it is placed. Second, and most importantly for mechanism, the market-structure heterogeneity in the post-NCA incumbency premium is the strongest finding separating the productive-capital story from pure patronage. A consumptive transfer—fiscal pork, cash payments, or generic patronage—generates no conditional persistence by market structure; only a productive informational asset should generate lasting returns where it dominates the local media landscape and decay where it does not. That the post-NCA incumbency premium follows precisely this pattern—persisting in non-competitive markets and decaying in competitive ones—connects our allocation margin to the media-economics finding that media influence is largest where outlets are scarce (Strömberg, 2004), and rules out the most natural alternative interpretation.

Scope conditions and external validity. The mechanism we document is not universal. Broadcast concessions function as political currency when three conditions hold jointly: (i) executive discretion over license allocation is high, (ii) local information markets are thin enough that a single new station can dominate political information for a non-trivial share of the electorate, and (iii) the political window is one of institutional flux, in which executive–legislative bargaining is dense and the stakes of individual votes are unusually high. Brazil’s 1986 NCA is an extreme point on all three dimensions: the president had near-unilateral allocation authority over AM and television concessions, radio penetration in the interior was near-universal with few competitors per municipality, and the Constituent Assembly drew up the new rules of the political game. We therefore expect the mechanism to be most visible in contexts that approximate this constellation—for

example, Peru in the early 1990s, where McMillan and Zoido (2004) document the authoritarian-exchange analogue; Italy under Berlusconi, where Durante et al. (2019) document the private-empire version; Argentina under Menem’s privatization wave; Hungary after the 2010 media-authority consolidation; or Turkey after the 2001 TMSF seizures. In consolidated democracies with independent regulators and dense information markets—the OECD “Liberal” and “Democratic Corporatist” configurations of Hallin and Mancini (2004)—we expect the mechanism to be attenuated but not absent: politically valuable media assets can still be allocated through regulatory discretion at the margin, even when formal auction mechanisms are in place. Mapping the boundaries of this mechanism empirically in these settings is an obvious next step for the literature.

A further consideration strengthens rather than weakens extrapolation. The allocation mechanism is uniquely Brazilian, as executive discretion over AM and television licenses at this scale is rare in the OECD, but the *outcome* is not: politicians who own broadcast outlets operate in most democracies, from Italy and Spain to Mexico and the Philippines. What distinguishes the Brazilian NCA is that the allocation episode was broad enough (approximately ninety legislators received concessions) to populate the treated group with a cross-section of governing-coalition politicians rather than a handful of particularly powerful ones. This breadth implies that concession recipients in our sample are less positively selected on political strength than media-owning politicians in countries where ownership is concentrated among the most prominent figures. If anything, our estimates of the allocation and legislative effects are attenuated relative to what a more selected sample would produce. Moreover, Brazil after 1989 operates as a standard multiparty electoral democracy with open-list proportional representation, competitive elections, and a free press; the institutional environment in which the concession’s electoral value is realized is not historically specific but common to a wide range of contemporary settings.

The results open several natural avenues for further research. A first is methodological: a clean causal estimate of the long-run electoral return on concession receipt would

require an identification design with parallel pre-trends across treated and never-treated cohorts that the single-cohort 1986 setting does not support. Settings with rolling allocation across cohorts—for example, license renewals over multiple legislative terms, or country-level allocations with staggered timing—would allow direct identification of the long-run electoral effect that our market-structure heterogeneity analysis can only signal indirectly. A second direction is to investigate the mechanisms through which concession recipients convert media ownership into political advantage: whether they face weaker electoral accountability, engage differently in distributive politics, or shape local information environments in ways that reinforce their political position. Direct measurement of legislative effort and constituent service among media-owning deputies is a natural next step. A third direction is comparative. The Brazilian NCA offers an unusually transparent setting for studying the political allocation of broadcasting rights, but discretionary media control has played an important role in many other democratizing episodes—in Latin America, Southern Europe, and Eastern Europe in particular. Documenting whether similar mechanisms operated in those settings would help identify the conditions under which broadcast concessions emerge as a preferred instrument of executive-legislative bargaining during institutional transitions.

Our results underscore the importance of regulatory frameworks that limit executive discretion over broadcast concessions and enforce conflict-of-interest rules for media-owning legislators. More broadly, the mechanism we document—the use of discretionary broadcast licensing as political currency during high-stakes institutional bargaining—is likely to operate in any context where executive discretion over spectrum allocation meets dense executive-legislative negotiation, a pattern whose scope conditions and boundaries future work can map empirically.

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A Additional Tables and Graphs

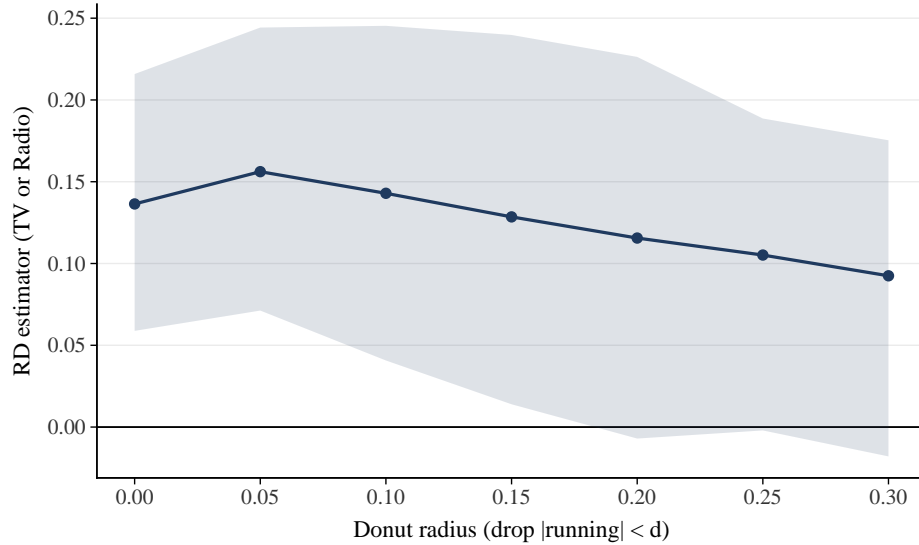


Figure A1: Donut-sensitivity path of the 1986 RD estimator

Notes: Each point plots the local-linear RD estimator of the effect of narrowly winning a 1986 NCA seat on the probability of receiving a radio or TV concession in 1987–1988, estimated after dropping all observations with $|running| < d$ for d on the horizontal axis. Bandwidth is MSE-optimal at each d ; kernel is triangular; confidence bands are 95% bias-corrected following Calonico et al. (2014). The estimator is flat or rising as the donut widens to 0.10 and only modestly attenuates thereafter, inconsistent with strategic manipulation at the threshold.

Table A1: Local polynomial density test at the RD cutoff

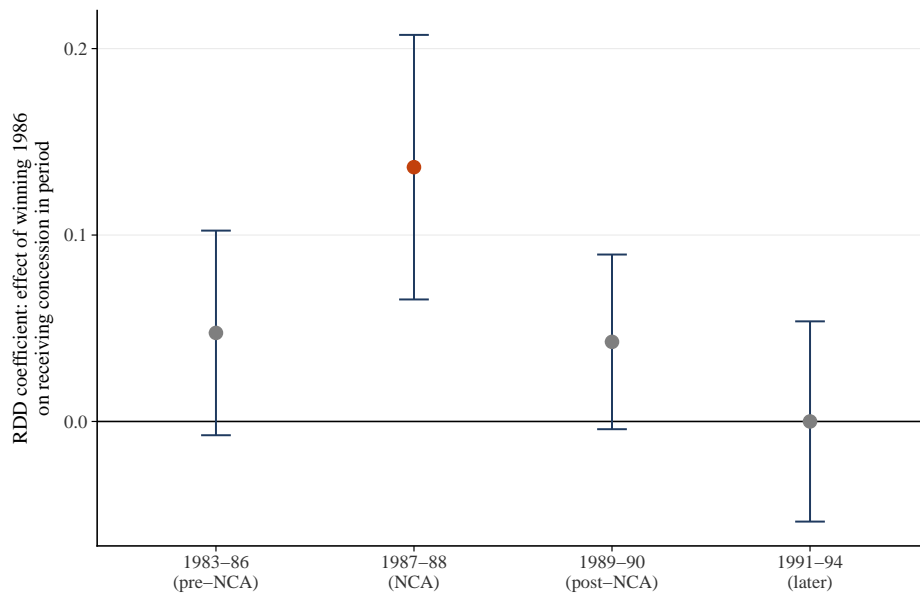
Sample	N	T_{CJM}	p -value
Government (PMDB + PFL)	1,142	3.002***	0.003
Left Opposition	566	0.312	0.755
Right Opposition	906	-4.177***	0.000

Notes: Local-polynomial density test of Cattaneo et al. (2020) at the seat-threshold cutoff, estimated separately by political alignment. Government: PMDB + PFL (Sarney’s governing coalition, $N = 1,142$); Left Opposition: PDT, PT, PSB, PCB, and PC do B ($N = 566$); Right Opposition: non-PMDB/PFL non-left parties ($N = 906$). The rejection for Government and Right Opposition reflects the same structural asymmetry documented for the pooled sample—proportional representation produces mechanically more marginal losers than winners per coalition list—and not strategic sorting. T_{CJM} is the jackknife T-statistic; p -values are two-sided. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

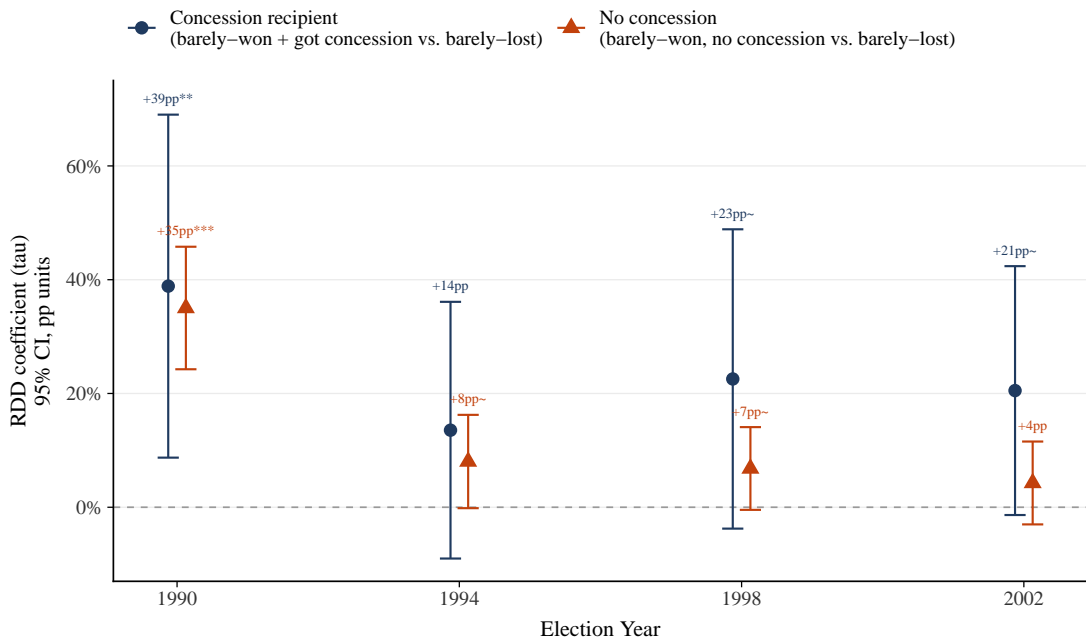
Table A2: Covariate Smoothness at the 1986 Electoral Cutoff

Variable	Estimate	SE	Robust 95% CI		<i>p</i> -value	Eff. <i>N</i>
			Lower	Upper		
Elected in 1982	0.136	0.067	-0.036	0.249	0.142	1454
Ran in 1982	0.093	0.068	-0.080	0.211	0.375	1495
Held media concession 1983–86	0.047	0.028	-0.015	0.111	0.133	1852
Held radio concession 1983–86	0.026	0.024	-0.030	0.078	0.390	1775
Held TV concession 1983–86	0.020	0.017	-0.015	0.058	0.256	1976
PMDB in 1986	0.146	0.066	-0.023	0.261	0.101	1479
PFL in 1986	0.144***	0.043	0.056	0.246	0.002	1957

Notes: Each row reports the RDD estimate (conventional local-linear, $p = 1$) and robust 95% confidence interval from an `rdrobust` regression using each pre-treatment covariate as the outcome. Bandwidth: MSE-optimal, triangular kernel, nearest-neighbor variance estimator. Eff. *N* is the number of observations within the MSE-optimal bandwidth. *p*-values are robust (bias-corrected). * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

**Figure A2:** Temporal placebo: RDD estimates of winning the 1986 election on concession receipt, by period

Notes: Each point plots the RDD coefficient of barely winning the 1986 NCA election on the probability of receiving a media concession during the indicated period. All estimates use the same local-polynomial specification as Panel A of Table 1. The NCA period (1987–88, highlighted) is the only one with a statistically significant discontinuity ($p < 0.001$). Error bars are 95% bias-corrected confidence intervals.



Notes: Each point is an RDD estimate using the pooled MSE-optimal bandwidth. Right side split by concession receipt in 1987–88; left side (1986 losers) is the shared count

Figure A3: Suggestive evidence: RDD estimates split by concession receipt

Notes: Each point plots a local-linear RDD coefficient on the probability of being elected as a federal deputy in the indicated year, using the 1986 electoral margin as the running variable. The left side (barely-lost in 1986) is the shared counterfactual for both series. *Concession recipient*: right side restricted to the 80 winners who received a broadcasting concession in 1987–88. *No concession*: right side restricted to the 407 winners who did not. Both series use a triangular kernel at the MSE-optimal bandwidth from the pooled RDD for that year ($h \approx 1.1$). Error bars are conventional 95% confidence intervals. Concession receipt among winners is endogenous; the split is illustrative, not causal identification.

PFL list-position and the loser-to-winner ratio. Figure A7 documents the within-coalition density asymmetry at the party level. Within the 25 states where PMDB and PFL ran on joint coalition lists, PFL candidates accumulated proportionally far more near-margin losers than winners relative to PMDB. At the MSE-optimal bandwidth ($h = 1.44$), PFL’s loser-to-winner ratio is 2.93 versus 1.64 for PMDB—a gap of 1.8× that is stable across all bandwidth choices. This party-level asymmetry is the direct source of the +14.4 pp discontinuity in PFL membership in the covariate-smoothness test (Table A2): the local polynomial detects the density gradient—proportionally more PFL candidates just below the threshold than just above—as a jump in PFL membership at zero. The ratio is structural: PFL, as the junior coalition partner, claimed fewer seats per list, concentrat-

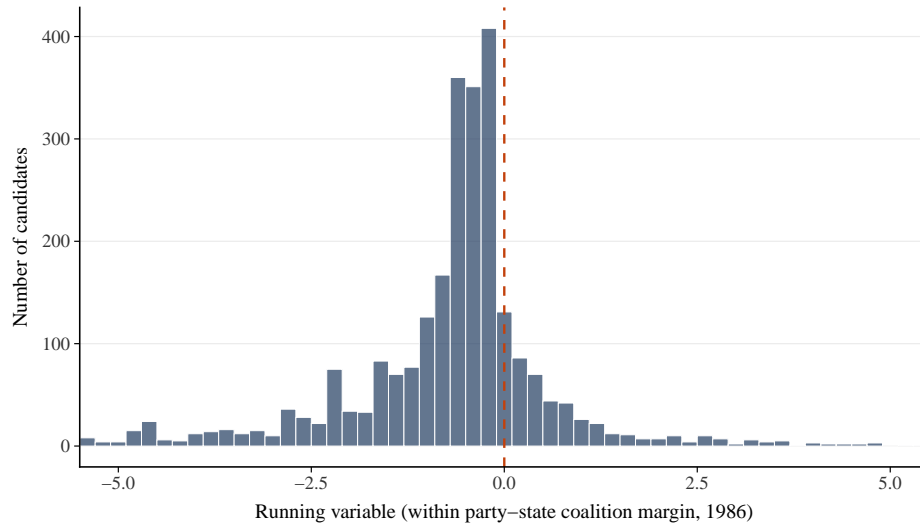


Figure A4: Distribution of the running variable (1986 within-coalition margin)
Notes: Companion to the RD design described in Section 5.1 of the main text. Histogram of the running variable (distance to the effective seat threshold within each party-state coalition) for all 2,614 candidates in the RD population. Vertical dashed line marks the cutoff at zero. Bins of width 0.2.

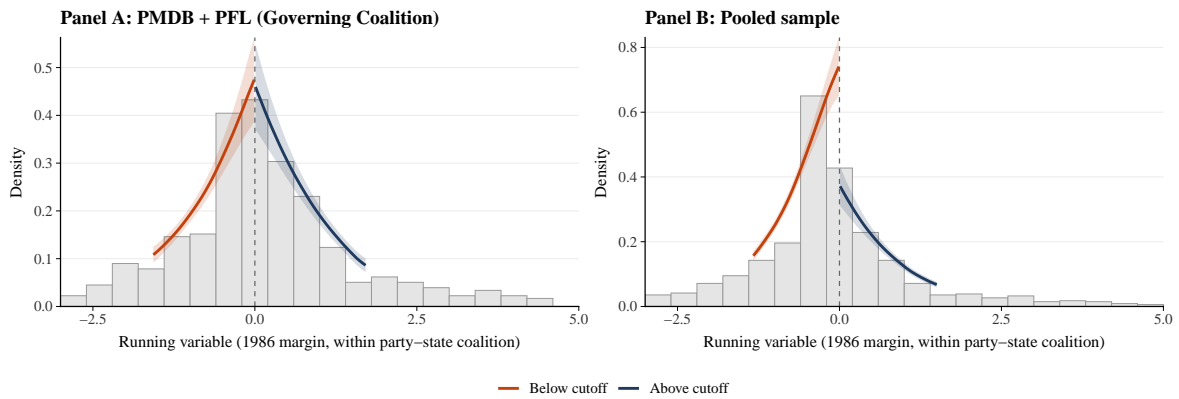


Figure A5: Local-polynomial density of the 1986 running variable at the cutoff
Notes: Panel A restricts to PMDB (Sarney's governing coalition); Panel B pools all parties. Curves are local-polynomial density estimates of Cattaneo et al. (2020) computed separately on each side of the cutoff, with 95% pointwise confidence intervals. The histogram in the background uses a fixed binwidth of 0.4. The bandwidth used for the visual smoothing is 1.4 times the MSE-optimal bandwidth returned by `rdensity`. Vertical dashed line marks the seat threshold.

Table A3: Archive-records sensitivity: three headline numbers

	Auto-match only	+ Motter archival
<i>Panel A: Main 1986 close-election RDD on 1987–88 concession receipt</i>		
TV or Radio	0.060** (0.028)	0.136*** (0.036)
Radio	0.044 (0.025)	0.131*** (0.035)
TV	0.031 (0.020)	0.048** (0.023)
<i>Panel B: Within-winner Loyalty analysis (state-by-party fixed effects)</i>		
Loyalty Index	0.062 (0.051)	0.130*** (0.027)
Non-Divisive Index	0.157** (0.065)	0.296*** (0.048)
Divisive Index	0.002 (0.041)	-0.037 (0.033)
Difference (non-div - div)	0.380*** (0.068)	0.454*** (0.051)

Notes: Re-estimation of the three headline results under two coding choices for concession receipt. “Auto-match only” uses the automated CNE-TSE matching procedure described in Section 4. “+ Motter archival” adds the additional cases recovered by Motter et al. (2019) from Ministry of Communications paper files, which is the coding used throughout the main text. Panel A is the main 1986 close-election RDD (equivalent to Panel A of Table 1) on three outcomes: any media concession, radio only, TV only. Panel B is the within-winner Loyalty analysis (equivalent to Column 2 of Table 3) with state-by-party fixed effects; the Difference row is the coefficient on the interaction $\text{Concession} \times \mathbb{1}\{\text{Institutional}\}$ in the stacked deputy-index regression. Triangular kernel, MSE-optimal bandwidth, nearest-neighbor variance in Panel A; state-by-party fixed effects with state-by-party cluster-robust standard errors in Panel B. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

ing its winners at the very margin of what shared lists could deliver and leaving many more PFL candidates just short of election.

Fistel heterogeneity as a continuous interactor. The state-level Fistel heterogeneity in Figure 5 splits the 27-unit state-level Fistel distribution at its median. A binary split loses statistical power and depends on the choice of cutoff. Table A4 re-estimates the same heterogeneity with Fistel density entered as a continuous (z-scored) interactor in a pooled local-linear regression that fully interacts the above-cutoff indicator, the running-variable

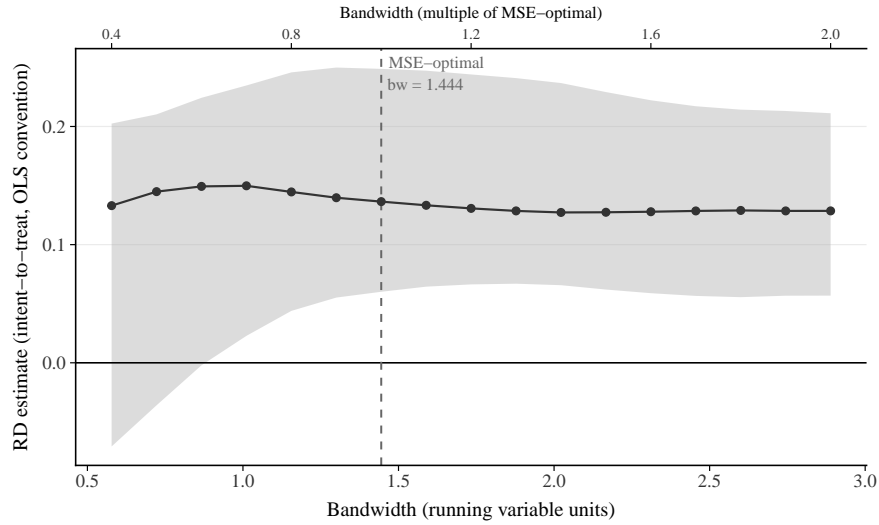


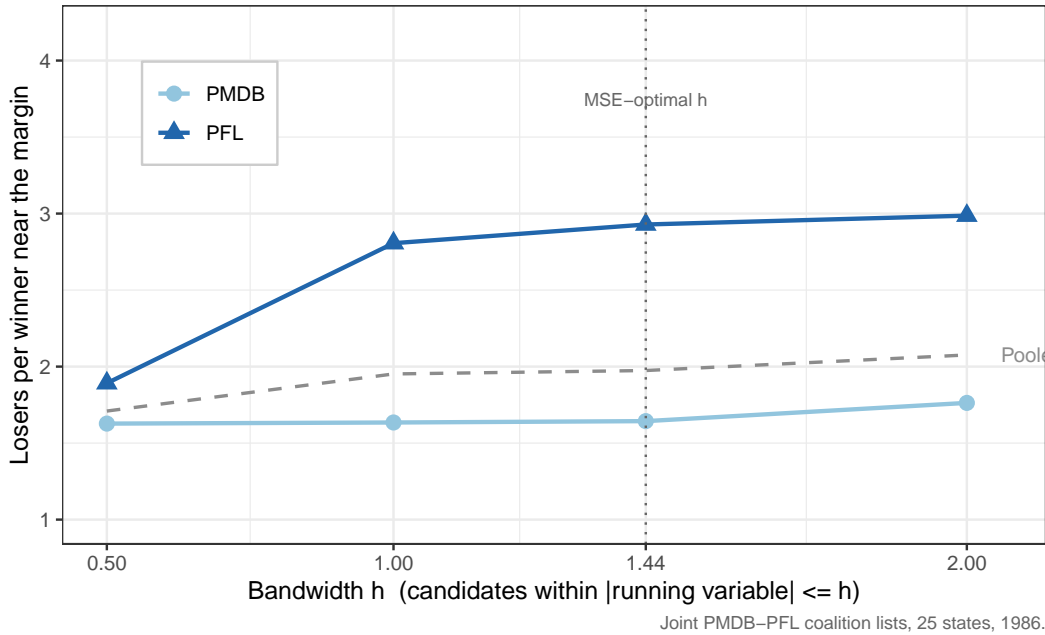
Figure A6: Bandwidth Sensitivity of the 1986 RDD Estimate

Notes: Companion to the bandwidth-sensitivity discussion in Section 5.1 of the main text (Panel A of Table 1). Each point reports the estimated RDD effect of narrowly winning the 1986 election on the probability of receiving a media concession, using a local linear specification and bandwidths ranging from $0.4\times$ to $2.0\times$ the CCT-optimal bandwidth. Shaded regions show 95% confidence intervals. The vertical dashed line indicates the CCT-optimal bandwidth.

polynomial, and the standardized Fistel measure. We exclude the Federal District from the sample for this exercise, since DF is an extreme outlier with only one municipality and therefore a mechanical Fistel-per-municipality of 24.

The reported $\beta_{\text{above} \times \widetilde{\text{Fistel}}}$ coefficient measures how much the RDD estimate changes per one-standard-deviation increase in state-level Fistel density. On elected-1994—the year at which the binary split previously produced a marginally significant difference—the continuous interaction is -0.032 ($p = 0.067$), directionally consistent with the binary-split finding and at comparable marginal significance. The implied RDD coefficient at the 25th, 50th, and 75th percentiles of the candidate-weighted Fistel distribution is 10.7, 8.2, and 7.8 percentage points respectively, displaying the expected monotone decline with market competitiveness. Other outcome years show small and statistically insignificant interactions, in line with the binary-split results in the main text. Two takeaways. First, the

Figure A7: Loser-to-winner ratio near the electoral margin, by party



Notes: Each point shows the ratio of candidates with a negative running variable (losers) to candidates with a positive running variable (winners) within bandwidth h , computed separately for PMDB (circles) and PFL (triangles). Sample restricted to the 25 states where both parties ran on the same coalition list in 1986 ($N = 1,142$). The dashed grey line shows the pooled PMDB+PFL ratio; the vertical dotted line marks the MSE-optimal bandwidth ($h = 1.44$). PFL's loser-to-winner ratio is approximately $1.8\times$ higher than PMDB's at the MSE-optimal bandwidth and stable across bandwidth choices, reflecting a structural feature of open-list proportional representation: PFL, as the junior coalition partner, claimed fewer seats per list and concentrated its winners at the margin, generating the density gradient that the covariate test detects as a jump in PFL membership.

continuous specification does not overturn the main-text finding—the Fistel heterogeneity is real in 1994 but modest and mostly absent in other years. Second, the continuous specification yields slightly sharper inference on the 1994 contrast (moving the p -value from roughly 0.10 under the binary split to 0.07 here), consistent with the expected power gain from avoiding an arbitrary cutoff.

Loyalty analysis separately for PMDB and PFL. The within-winner Loyalty analysis in the main text pools candidates from PMDB, the centrist anchor of Sarney's governing coalition, and PFL, the right-wing partner formed largely by former members of the military regime's ARENA. Table A7 re-estimates the four Loyalty-Index-family coefficients

Table A4: Fistel heterogeneity with a continuous interactor

	β_{above}	$\beta_{\text{above} \times \widetilde{\text{Fistel}}}$	Implied RDD at Fistel percentile		
	(at Fistel mean)	(per 1 SD)	$p_{25} = 1.17$	$p_{50} = 2.01$	$p_{75} = 2.15$
<i>Panel A: Dependent variable: $\mathbb{1}\{\text{Elected}_t\}$</i>					
1990	0.256*** (0.054)	0.029 (0.043)	0.249	0.271	0.275
1994	0.099*** (0.037)	-0.032* (0.017)	0.107	0.082	0.078
1998	0.087** (0.036)	0.002 (0.030)	0.087	0.088	0.088
2002	0.053* (0.031)	-0.026 (0.018)	0.059	0.039	0.036
<i>Panel B: Dependent variable: $\mathbb{1}\{\text{Ran}_t\}$</i>					
1990	0.418*** (0.065)	0.007 (0.056)	0.417	0.422	0.423
1994	0.074 (0.052)	-0.011 (0.023)	0.076	0.068	0.066
1998	0.098 (0.061)	0.030 (0.046)	0.090	0.114	0.118
2002	0.105** (0.045)	-0.043 (0.034)	0.115	0.082	0.076

Notes: Companion to the heterogeneous-persistence discussion in Section 5.4 of the main text (companion binary-split results in Figure 5). Pooled local-linear regression on the union bandwidth (MSE-optimal for the full sample) with a full three-way interaction between the above-cutoff indicator, the 1986 running variable, and the (z-scored) state-level Fistel-per-municipality density. Robust standard errors in parentheses. Implied RDD columns compute $\beta_{\text{above}} + \beta_{\text{above} \times \widetilde{\text{Fistel}}} \cdot z_p$ where z_p is the z-score of the p th percentile of the candidate-level Fistel distribution (column headers report the raw Fistel-per-municipality value at each percentile). Federal District is excluded as an extreme outlier. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A5: HLATE 2 Robustness: Donut RDD and Covariate Adjustment (Fistel State-Level Heterogeneity)

	1990	1994	1998	2002
<i>Panel A: Dependent variable: $\mathbb{1}\{Ran_t\}$</i>				
Baseline	-0.132 (0.119)	0.172* (0.098)	0.076 (0.097)	0.104 (0.086)
Eff. N (non-comp. + comp.)	2104	2139	2126	2056
Donut 0.10	-0.178 (0.147)	0.151 (0.114)	0.051 (0.115)	-0.023 (0.094)
Eff. N (non-comp. + comp.)	1909	2001	1995	1914
Donut 0.25	0.064 (0.227)	0.289 (0.177)	0.024 (0.196)	-0.111 (0.151)
Eff. N (non-comp. + comp.)	1490	1553	1442	1595
+ Covariates	-0.120 (0.120)	0.179* (0.097)	0.107 (0.097)	0.103 (0.086)
Eff. N (non-comp. + comp.)	2104	2139	2126	2056
<i>Panel B: Dependent variable: $\mathbb{1}\{Elected_t\}$</i>				
Baseline	-0.082 (0.092)	0.139* (0.084)	0.057 (0.072)	0.012 (0.058)
Eff. N (non-comp. + comp.)	2227	2126	2272	2063
Donut 0.10	-0.176* (0.107)	0.172* (0.104)	-0.006 (0.089)	-0.022 (0.058)
Eff. N (non-comp. + comp.)	1988	1964	2002	2135
Donut 0.25	-0.048 (0.169)	0.334** (0.142)	0.086 (0.116)	0.001 (0.101)
Eff. N (non-comp. + comp.)	1586	1546	1660	1646
+ Covariates	-0.066 (0.090)	0.146* (0.082)	0.063 (0.070)	0.011 (0.058)
Eff. N (non-comp. + comp.)	2227	2126	2272	2063

Notes: Each cell reports the non-competitive-minus-competitive difference (the coefficient on $\mathbb{1}\{\text{running} \geq 0\} \times \mathbb{1}\{\text{non-competitive}\}$) from a pooled local-linear regression on the union bandwidth. “Donut d ” drops observations with $|\text{running}| < d$; “+ Covariates” adds pre-determined controls (1986 vote share, 1982 election participation, PMDB dummy, PFL dummy) additively following Calonico et al. (2019). Triangular kernel with MSE-optimal bandwidth selected within each cell. Conventional standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A6: Fistel Heterogeneity: Bias-Corrected CIs and Minimum Detectable Effect

Panel A: Elected as federal deputy				
	1990	1994	1998	2002
$\hat{\beta}_{\text{non-comp.}}$	0.230**	0.205**	0.133*	0.067
Robust 95% CI	[0.040, 0.402]	[0.026, 0.379]	[-0.020, 0.293]	[-0.041, 0.181]
$\hat{\beta}_{\text{competitive}}$	0.302***	0.062	0.071*	0.049
Robust 95% CI	[0.182, 0.411]	[-0.013, 0.129]	[-0.005, 0.143]	[-0.023, 0.106]
Difference (non-comp. – comp.)	-0.082	0.139*	0.057	0.012
HC-robust SE	(0.092)	(0.084)	(0.072)	(0.058)
MDE (80%, $\alpha=0.10$)	0.230	0.209	0.180	0.143
Pooled BW	2.21	1.90	2.43	1.65
Panel B: Ran as federal deputy				
	1990	1994	1998	2002
$\hat{\beta}_{\text{non-comp.}}$	0.358**	0.264**	0.229**	0.197***
Robust 95% CI	[0.058, 0.591]	[0.060, 0.472]	[0.014, 0.403]	[0.055, 0.364]
$\hat{\beta}_{\text{competitive}}$	0.452***	0.061	0.119*	0.045
Robust 95% CI	[0.304, 0.567]	[-0.060, 0.148]	[-0.011, 0.214]	[-0.053, 0.115]
Difference (non-comp. – comp.)	-0.132	0.172*	0.076	0.104
HC-robust SE	(0.119)	(0.098)	(0.097)	(0.086)
MDE (80%, $\alpha=0.10$)	0.297	0.243	0.241	0.214
Pooled BW	1.80	2.00	1.91	1.62

Notes: Re-estimation of the Fistel state-level heterogeneity with bias-corrected inference. Panels A–B mirror Figure 5. Subsample betas $\hat{\beta}_{\text{non-comp.}}$ and $\hat{\beta}_{\text{competitive}}$ are estimated via `rdrobust` with triangular kernel and MSE-optimal bandwidth; reported intervals are the robust (bias-corrected) 95% CIs of Calonico et al. (2014). The Difference row is the coefficient on $\mathbb{1}\{\text{running} \geq 0\} \times \mathbb{1}\{\text{non-competitive}\}$ from a pooled local-linear regression on the union bandwidth with HC-robust standard errors (in parentheses). The MDE row reports the minimum detectable effect at 80% power and $\alpha = 0.10$ (two-tailed): $\text{MDE} = (z_{0.80} + z_{0.95}) \times \text{SE} = 2.487 \times \text{SE}$. At the 1994 horizon the MDE is 20.9 pp (Elected) and 24.3 pp (Ran), both exceeding the observed point estimates of +13.9 pp and +17.2 pp, confirming the split is underpowered to reject the null at conventional thresholds. The directional pattern— $\hat{\beta}_{\text{non-comp.}}$ remaining positive and significant in 1990–1998 while $\hat{\beta}_{\text{competitive}}$ fades after 1990—is consistent with the productive-political-capital framework. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

separately for each party. The Institutional-minus-Ideological differential—the paper’s preferred within-deputy measure—is positive and significant in both PMDB (+37.8 percentage points, $p = 0.001$) and PFL (+16.4 percentage points, $p = 0.010$). The PMDB estimate is roughly twice as large. This is consistent with the mechanism: PFL candidates were already ideologically closer to the Executive on the two institutional items (presidentialism and the five-year term), so the marginal effect of a concession on their institutional loyalty is smaller; PMDB candidates spanned a wider ideological range within the centrist coalition, so concessions could shift more of them on the institutional margin. On the Ideological Index, the concession-recipient coefficient is near zero for PFL (-0.005 , $p = 0.90$); for PMDB the coefficient is negative and marginally significant (-0.063 , $p = 0.077$), consistent with recipients voting slightly less in favor of left-opposition positions on ideological issues as well. The concentration of the response on the Institutional margin—where institutional loyalty has direct constitutional stakes—remains the dominant pattern in both parties.

Table A7: Loyalty analysis separately for PMDB and PFL

	PMDB (centrist gov party)	PFL (right-wing ally)
Loyalty Index	0.102*** (0.035)	0.112** (0.046)
Non-Divisive Index	0.273*** (0.070)	0.175** (0.065)
Divisive Index	-0.063* (0.034)	-0.005 (0.036)
Difference (non-div – div)	0.378*** (0.104)	0.164** (0.059)
Deputies	252	114
State-party clusters	25	26

Notes: Re-estimation of the four Loyalty-Index-family coefficients from Table 3 (Panel A Column 2 for the Loyalty Index; Panel B for Institutional, Ideological, and Difference) separately for candidates whose 1986 party was PMDB (centrist anchor of the governing coalition) and PFL (right-wing partner formed largely by former members of the military-regime ARENA). State-by-party fixed effects; standard errors clustered at the state-by-party level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A8: Wild-Cluster Bootstrap p -values for Roll-Call Headline Coefficients

	Coefficient (SE in parens)	Cluster-robust t -stat	Cluster-robust p -value	Wild-cluster bootstrap p
Loyalty Index (Panel A, Col 2 of Table 3)	0.130*** (0.027)	4.80	0.0000	0.0140
Non-Divisive Index (Panel B)	0.296*** (0.048)	6.21	0.0000	0.0000
Divisive Index (Panel B)	-0.037 (0.033)	-1.12	0.2677	0.1311
Difference non-divisive – divisive (Panel B)	0.454*** (0.051)	8.94	0.0000	0.0000

Notes: Wild-cluster bootstrap p -values for the four headline coefficients from the within-winner roll-call analysis (Table 3). The specification is Column 2 of Panel A (Loyalty Index) and Panel B (Institutional, Ideological, Difference), with state-by-party fixed effects and 95 state-by-party clusters. Bootstrap uses $B = 999$ Rademacher sign flips at the state-by-party cluster level on the unrestricted model residuals. The Institutional Index, Ideological Index, and the Institutional-minus-Ideological gap are unaffected by the small-cluster correction; the Loyalty Index p -value rises from < 0.001 to 0.014, remaining significant at the five-percent level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Placebo: close-election RDD for every federal election 1982–1998. A comprehensive check on the NCA-specificity of the main RDD is to re-run the same design on every federal Chamber-of-Deputies election from 1982 to 1998 with the outcome defined as the receipt of at least one radio or television concession (self or close relative) during the subsequent four-year legislative term. If marginal close-election winning generated broadcasting concessions as a generic feature of Brazilian politics, the pattern documented in Panel A of Table 1 should recur in other cycles. If it is specific to the NCA bargaining environment, only the 1986 election should produce a positive discontinuity.

To keep the comparison consistent across cycles, the outcome in every row of Table A10 and Figure A8 is constructed using *only* the automated CNE–TSE matching procedure described in Section 4. The archival records from Motter et al. (2019), which complement the automated match in the main text by recovering additional cases of indirect ownership registered under proxies, are *not* used here. The reason is logical: Motter et al. (2019) audited the 1987–88 NCA period specifically; he did not assemble a comparable enumeration for 1983–86, 1991–94, 1995–98, or 1999–2001. Including his additions only in the NCA row would mechanically inflate the 1986 estimate relative to the placebo rows and defeat

Table A9: Descriptive Statistics Across Estimating Samples

	Full RDD sample		Winners (roll-call)		Concession recipients	
	Mean	SD	Mean	SD	Mean	SD
<i>Running variable</i>						
1986 running variable (within-coalition margin, %)	-0.611	2.151	1.691	3.242	0.672	2.975
1986 absolute vote share (%)	0.926	2.139	3.400	3.855	2.364	3.432
<i>Political outcomes</i>						
Elected in 1986 NCA	0.186	0.389	1.000	0.000	0.548	0.499
Elected in 1982	0.173	0.378	0.461	0.499	0.372	0.485
Ran in 1982	0.288	0.453	0.509	0.500	0.437	0.497
Elected in 1990	0.110	0.313	0.419	0.494	0.286	0.453
Elected in 1994	0.057	0.231	0.202	0.402	0.121	0.326
Elected in 1998	0.053	0.224	0.183	0.387	0.136	0.343
<i>Party</i>						
PMDB in 1986	0.273	0.445	0.531	0.500	0.327	0.470
PFL in 1986	0.164	0.370	0.240	0.428	0.291	0.456
PDS in 1986	0.097	0.296	0.069	0.255	0.065	0.248
<i>Concession (treatment)</i>						
Received media concession 1987–88	0.076	0.265	0.225	0.418	1.000	0.000
Received radio concession 1987–88	0.062	0.240	0.194	0.396	0.809	0.394
Received TV concession 1987–88	0.023	0.149	0.067	0.251	0.296	0.458
<i>Concession (pre-1986)</i>						
Held media concession 1983–86	0.043	0.203	0.078	0.268	0.302	0.460
Held radio concession 1983–86	0.034	0.181	0.057	0.232	0.216	0.413
Held TV concession 1983–86	0.013	0.115	0.032	0.175	0.116	0.321
<i>Roll-call (winners only)</i>						
Loyalty Index (mean of 7 NCA roll-call supports)	–	–	0.653	0.221	–	–
Divisive Index (mean of 5 ideological votes)	–	–	0.651	0.171	–	–
Non-Divisive Index (mean of 2 institutional votes)	–	–	0.585	0.434	–	–
Observations	2,614		475		199	

Notes: Means and standard deviations for the variables used in the empirical analysis, reported separately for the three estimating samples: the *full 1986 RDD sample* (all 1986 Chamber-of-Deputies candidates with a non-missing within-coalition running variable); the *winner roll-call sample* (the subset of 1986 winners linked to the digitized NCA roll-call data used in the loyalty regressions); and the *concession-recipient sample* (candidates coded as having received a radio or TV concession in 1987–88). The running variable is the within-party-coalition-state distance to the effective seat threshold following Novaes (2024); the 1986 vote share is the candidate's absolute share within their state. Electoral outcomes come from the Tribunal Superior Eleitoral (TSE). Party membership is the 1986 ballot party. Concession variables are constructed by linking firm-partner records (CNE) and the Ministry of Communications broadcasting registry to individual politicians, following the matching procedure described in Section 4. The Loyalty, Ideological, and Institutional indices are the mean deputy-level support on the seven NCA roll-call topics (Loyalty) or their two sub-groups (Ideological: agrarian reform, right to strike, job stability, work hours, property rights; Institutional: presidential system and five-year term). Roll-call indices are defined only for the winner sample.

the purpose of the comparison. We therefore restrict every row to the automated match, which recovers a strict subset of the concessions Motter documents. This makes the entire placebo series a *lower bound* on the true difference between the NCA cycle and the other cycles: on the automated match alone, the 1986 estimate is +6.0 percentage points on the original 1987–88 two-year window ($p = 0.034$) and +8.1 percentage points on the four-year 1987–90 window reported below ($p = 0.019$); adding Motter’s archival additions raises the 1987–88 estimate to the +13.6 percentage points reported in Panel A of Table 1. For the placebo argument the automated-match magnitude is sufficient and, because Motter’s audit is absent from the comparison cycles, the honest one.

Table A10 and Figure A8 report the result. The 1986 row reproduces a positive and statistically significant effect on all three outcomes (+8.1 pp for any media, +5.8 for radio, +6.1 for TV). No other election in the sample produces a comparable discontinuity: the 1982 estimate is positive but not significant (+3.7 pp, $p = 0.13$ for any media), and the 1990, 1994, and 1998 estimates are null or slightly negative on all three outcomes, none statistically distinguishable from zero. An alternative hypothesis that the 1982 election produced concessions to buy votes against the 1984 Diretas Já amendment receives no support in these data: the 1982 point estimate on the 1983–84 pre-Diretas window is +1.3 pp ($p = 0.47$, not shown in the table). Two ancillary notes on interpretation. First, the bandwidths reported for the 1994 and 1998 rows are numerically smaller than those in the other rows because the running variable for those elections is expressed in decimal rather than percentage units; since `rdrobust` is scale-invariant at the cutoff, the coefficients are directly comparable. Second, the precision of each placebo row is sufficient to rule out effects of the size estimated in the NCA period. Together, the table and the figure make the NCA-specificity of the pattern visible at a glance.

Table A10: Placebo RDDs: Federal Elections 1982–1998 on Concessions During the Subsequent Term

Election (outcome window)	Any	Radio	TV
1982 (window 1983–1986)	0.037 (0.021)	0.027 (0.018)	0.001 (0.011)
Mean outcome	0.032	0.024	0.009
Eff. obs. (L/R)	857 / 399	856 / 399	803 / 359
Bandwidth (MSE)	2.21	2.19	1.63
RD population		1,587	
1986 (window 1987–1990)	0.080** (0.032)	0.058* (0.028)	0.061** (0.026)
Mean outcome	0.073	0.057	0.027
Eff. obs. (L/R)	1569 / 348	1541 / 343	1534 / 338
Bandwidth (MSE)	1.32	1.23	1.18
RD population		2,614	
1990 (window 1991–1994)	-0.017 (0.020)	-0.012 (0.018)	-0.007 (0.012)
Mean outcome	0.032	0.026	0.011
Eff. obs. (L/R)	1613 / 327	1710 / 342	1759 / 348
Bandwidth (MSE)	1.17	1.28	1.33
RD population		2,907	
1994 (window 1995–1998)	0.007 (0.023)	-0.020 (0.019)	0.016 (0.017)
Mean outcome	0.039	0.024	0.018
Eff. obs. (L/R)	885 / 215	760 / 181	836 / 200
Bandwidth (MSE)	0.01	0.01	0.01
RD population		1,430	
1998 (window 1999–2001)	-0.015 (0.013)	-0.009* (0.005)	-0.001 (0.012)
Mean outcome	0.022	0.008	0.015
Eff. obs. (L/R)	1362 / 322	1149 / 255	1324 / 310
Bandwidth (MSE)	0.01	0.01	0.01
RD population		2,602	

Notes: Each block corresponds to one federal election year. The running variable is the within-party-coalition-state margin for that election; the outcome is an indicator for receiving at least one concession (self or close relative) during the legislative term that follows, using the automated CNE–TSE matching procedure only (without the archival additions of Motter et al. 2019). Restricting the outcome to the automated match is essential for the placebo comparison, because Motter et al. (2019) audited the 1987–88 window specifically and not the placebo windows; any residual difference between the 1986 row and the other rows therefore represents a strict lower bound on the true NCA effect (adding Motter’s records to the 1987–88 window raises the main estimate from +6.0 pp to +13.6 pp, as reported in Panel A of Table 1). Rows within each block report, in order: the RD estimator (robust significance stars), the conventional standard error in parentheses, the sample mean of the outcome, the effective observations on the left and right of the cutoff, the MSE-optimal bandwidth, and the full RD population of candidates with a non-missing running variable. The bandwidth for the 1994 and 1998 rows is expressed in the native scale of those running variables, which is decimal rather than

Table A11: Sensitivity of the Within-Deputy Differential: Coefficient Stability and Cinelli–Hazlett Bounds

	Within-deputy differential (Non-Divisive – Divisive)
<i>Panel A: Coefficient stability</i>	
Uncontrolled coefficient ($\tilde{\beta}$)	0.454
Controlled coefficient ($\hat{\beta}$)	0.454
Uncontrolled R^2	0.107
Controlled R^2	0.219
<i>Panel B: Cinelli–Hazlett (2020) sensitivity</i>	
Robustness Value ($q = 1, \beta = 0$)	0.246
Robustness Value ($q = 1, \alpha = 0.05$)	0.191
Benchmark: $1 \times$ running variable, partial $R^2_{Y \sim Z X,D}$	5.25e-04
Benchmark: $1 \times$ pre-1986 radio, partial $R^2_{Y \sim Z X,D}$	1.54e-05
Benchmark: $1 \times$ pre-1986 TV (1985), partial $R^2_{Y \sim Z X,D}$	2.25e-04

Notes: The coefficient of interest is the within-deputy interaction $\text{Concession} \times \mathbb{1}\{\text{Institutional}\}$ in the stacked deputy-index regression that contrasts each deputy’s Institutional and Ideological index values; this is the +45.4 percentage-point (0.454) “Difference” row of Table 3 Panel B. Panel A documents coefficient stability: the controlled and uncontrolled estimates coincide at 0.454 because the within-deputy design already differences out deputy-level selection by construction—adding state-by-party fixed effects, the 1986 vote margin, and pre-1987 ownership controls is redundant to the within-deputy variation, so observables are mechanically uninformative about the direction of omitted-variable bias. For this reason, Oster’s (2019) δ^* statistic diverges by construction and is not reported; it would be misleading to interpret $\delta^* = \infty$ as evidence of unusual robustness in this setting. Panel B reports the Cinelli and Hazlett (2020) robustness value, which is the appropriate sensitivity measure because it explicitly targets confounders at any level, including confounders that vary at the issue-by-deputy level (which the within-deputy design does not rule out). $RV_{q=1} = 0.246$ means a confounder would have to explain roughly one quarter of the residual variance in both the treatment and the differenced outcome simultaneously to nullify the estimate. The benchmarks report the partial $R^2_{Y \sim Z|X,D}$ for the strongest observable covariates; the unobserved confounder would need to be roughly two orders of magnitude stronger than the strongest observable to nullify the effect.

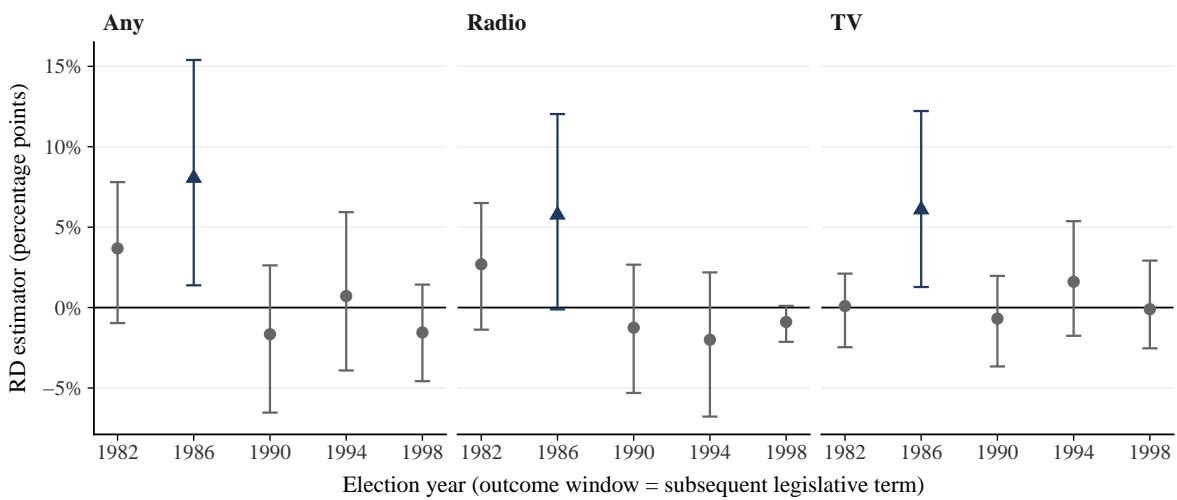


Figure A8: Placebo RDD estimates across federal elections, 1982–1998

Notes: Each point plots the RD estimator (conventional coefficient) of narrowly winning a federal Chamber-of-Deputies election on the probability of receiving at least one radio or television concession during the subsequent four-year legislative term (1999–2001 for the 1998 election, reflecting the endpoint of the outorga data). Bars show robust 95% confidence intervals bias-corrected following Calonico et al. (2014). The 1986 estimate is highlighted in navy; the remaining years (1982, 1990, 1994, 1998) are in grey. All outcomes are built from the automated CNE–TSE matching procedure *only*, without the archival additions of Motter et al. (2019), making the 1986–vs–other contrast a strict lower bound on the NCA-specific effect.

Online Appendix I

Historical Evidence

Not for Publication

B Contemporary Newspaper Evidence

Table B1: Representative Politician-Broadcasters, 1979–2000

Politician	UF	Main office(s)	Party (NCA)	Broadcasting outlet(s)
Antônio Carlos Magalhães	BA	Minister of Communications (1985–90); Governor; Senator	PFL	TV Bahia (Globo affiliate) and Rede Bahia
José Sarney	MA	President of Brazil (1985–90); Senator	PMDB	Sistema Mirante (TV, AM, FM, held through children)
Arnon de Mello / Fernando Collor	AL	Senator; Governor; President of Brazil (1990–92)	PMDB	Organizações Arnon de Mello (TV Gazeta, Rádio Gazeta FM)
Edison Lobão	MA	Senator; Minister of Mines & Energy	PFL	Sistema Difusora (Rádio and TV Difusora, Rádio Guajará, Rádio Imperatriz)
Jáder Barbalho	PA	Governor; Senator; Minister of Agrarian Reform	PMDB	Grupo RBA de Comunicação (RBA TV, Rádio Clube do Pará, Rádio Cabano)
Inocência Oliveira	PE	President of Chamber of Deputies (1993–95)	PFL	Rádio A Voz do Sertão (Serra Talhada)
Orestes Quécia	SP	Governor (1987–91); Senator	PMDB	Grupo Sol Panamby (Luzes da Ribalta AM, Nova Brasil FM network)
Marcondes Gadelha	PB	Minister of Communications (1989–90); Senator	PFL	Sistema Regional de Comunicação (FM Souza, AM Esperança)
João Alves Filho	SE	Governor; Minister of Interior	PFL	Rádio FM Jornal (Propriá, Aracaju, Estância)
Odacir Soares	RO	Senator	PFL	Rádio Clube de Rolim de Moura; Rádio Planalto; Rádio Sociedade Rondônia
Tasso Jereissati	CE	Governor; Senator	PMDB	Sistema Verdes Mares (TV Verdes Mares Globo affiliate, Rádio Verdes

This appendix presents contemporary newspaper material illustrating two central features of the institutional environment surrounding the National Constituent Assembly (NCA): (i) the political salience and discretionary nature of broadcast concessions, and (ii) the sharp decline in President José Sarney's popularity following the post-election collapse of the *Plano Cruzado*.

Figure B1 reproduces a front-page story from *Folha de S. Paulo*, published on November 15, 1988, describing the political allocation of radio concessions in the state of Bahia. The report describes how the Minister of Communications, Antonio Carlos Magalhães, distributed radio broadcasting licenses to political allies.

text discussion that the government entered the NCA with limited political capital and relied heavily on individualized legislative bargaining.



Figure B2: Newspaper headlines illustrating the post-election collapse of the *Cruzado* plan and the resulting decline in presidential popularity.

Notes: Compilation of headlines from *Tribuna da Imprensa*, *Jornal do Brasil*, and *Folha de S. Paulo* documenting the rapid deterioration of President Sarney's approval ratings following the November 1986 elections. The stories document the resurgence of inflation and internal divisions within the governing coalition that increased the Executive's reliance on individualized legislative bargaining during the NCA.

C Contentious Votes in the 1987–88 National Constituent Assembly

This appendix documents the procedure used to assemble and classify the voting records of the 1987–88 National Constituent Assembly (NCA), and provides background on the key legislative issues used in the empirical analysis. We relied on the official historical archive maintained by the Brazilian Chamber of Deputies, including the “25 anos da Constituição” digital collection.²⁸ All vote tallies, roll-call sheets, and government statements in this appendix come directly from these primary sources.

Data construction and coding procedure

The NCA voting records are preserved in heterogeneous formats, including scanned PDF documents, typed lists, and transcribed minutes. To convert these materials into a structured dataset, we implemented a multi-step workflow:

1. **Digitization via OCR.** Original roll-call sheets were processed using Optical Character Recognition (OCR) tools. Manual inspection ensured accuracy in names, vote options, and abstention codes.
2. **Name standardization.** Deputy names were standardized using a Stata/R cleaning routine that harmonized spelling variations, abbreviations, married names, and typographical inconsistencies. These standardized names were merged with the 1986 electoral registry to uniquely identify each deputy.
3. **Vote classification.** Each voting session was examined to extract the possible voting options (e.g., *Sim*, *Não*, *Abstenção*, *Obstrução*). When different abstention types

²⁸Available at: <https://www2.camara.leg.br/atividade-legislativa/plenario/discursos/escrevendohistoria/25-anos-da-constituicao-de-1988/momentos-marcantes-da-constituente>.

appeared, they were harmonized to a single “abstain” category.

4. **Government position coding.** Using official transcripts of the 1987–88 National Constituent Assembly (NCA), we identified the Executive’s stated position for each roll-call vote. Government floor leaders—particularly Deputy Mário Covas—publicly announced whether the Executive supported or opposed each amendment or motion. Based on these declarations in the plenary records, each roll-call was coded as “government *FOR*” or “government *AGAINST*.”
5. **Deputy-level loyalty index.** For each deputy i , topic t , and roll-call v , we coded:

$$Loyalty_{ivt} = \begin{cases} 1, & \text{if pro-government vote,} \\ 0, & \text{if anti-government vote.} \end{cases}$$

Abstentions were coded according to the “expected action” implied by the government’s stated preference (e.g., an abstention counted as opposition when the government needed an affirmative vote). For each deputy–topic pair, we then averaged $Loyalty_{ivt}$ across all votes v belonging to topic t :

$$Loyalty_{it} = \frac{1}{N_t} \sum_{v=1}^{N_t} Loyalty_{ivt},$$

yielding a measure of alignment between 0 and 1 for each deputy and each substantive issue debated in the Assembly.

The resulting dataset contains a harmonized panel of deputies \times key topics, enabling the empirical analysis reported in Section 5.

<p>O Sr. Roberto Freire — Sr. Presidente, peça a palavra, pela ordem.</p> <p>O SR. PRESIDENTE (Ulysses Guimarães) — Tem a palavra, pela ordem, o nobre Constituinte Roberto Freire.</p> <p>O SR. ROBERTO FREIRE (PCB — PE. Sem revisão do orador.) — Sr. Presidente, defendendo os direitos dos trabalhadores por uma jornada de quarenta horas, o Partido Comunista Brasileiro votará “sim”.</p> <p>O Sr. José Maria Eymael — Sr. Presidente, peça a palavra, pela ordem.</p> <p>O SR. PRESIDENTE (Ulysses Guimarães) — Tem a palavra, pela ordem, o nobre Constituinte José Maria Eymael.</p> <p>O SR. JOSÉ MARIA EYMAEL (PDC — SP. Sem revisão do orador.) — Sr. Presidente, o PDC foi co-autor da proposta de 44 horas semanais, visando a um equilíbrio. Dentro do espírito do acordo do primeiro turno, encaminho contrariamente a votação.</p> <p>O Sr. Nelson Jobim — Sr. Presidente, peça a palavra, pela ordem.</p> <p>O SR. PRESIDENTE (Ulysses Guimarães) — Tem a palavra, pela ordem, o nobre Constituinte Nelson Jobim.</p> <p>O SR. NELSON JOBIM (PDMB — RS. Sem revisão do orador.) — Sr. Presidente, o texto básico faz parte de algo que foi construído pelo PMDB na Comissão de Sistematização e que se reproduziu no primeiro turno. Portanto, o PMDB fica com o Relator, mantém a coerência de sua posição e vota “não”.</p> <p>O Sr. Artur da Távola — Sr. Presidente, peça a palavra, pela ordem.</p> <p>O SR. PRESIDENTE (Ulysses Guimarães) — Tem a palavra, pela ordem, o nobre Constituinte Artur da Távola.</p> <p>O SR. ARTUR DA TÁVOLA (PSDB — RJ. Sem revisão do orador.) — Sr. Presidente, a questão é aberta no PSDB.</p> <p>O SR. PRESIDENTE (Ulysses Guimarães) — Passamos à votação.</p>	<p>A Emenda foi rejeitada.</p> <p>VOTARAM OS SRS. CONSTITUINTES:</p> <p>Presidente Ulysses Guimarães — Abstenção</p> <p>Abigail Feitosa — Sim</p> <p>Acival Gomes — Sim</p> <p>Adauto Pereira — Não</p> <p>Ademir Andrade — Sim</p> <p>Adhemar de Barros Filho — Sim</p> <p>Adolfo Oliveira — Sim</p> <p>Adroaldo Streck — Não</p> <p>Adyson Motta — Não</p> <p>Aécio de Borba — Não</p> <p>Aécio Neves — Abstenção</p> <p>Afif Domingos — Não</p> <p>Afonso Sancho — Não</p> <p>Albano Franco — Não</p> <p>Albérico Cordeiro — Não</p> <p>Alceni Guerra — Abstenção</p> <p>Aldo Arantes — Sim</p> <p>Alfredo Campos — Não</p> <p>Aloysio Chaves — Não</p> <p>Aluizio Campos — Não</p> <p>Alysson Paulinelli — Não</p> <p>Amaury Müller — Sim</p> <p>Ângelo Magalhães — Não</p> <p>Anna Maria Rattes — Sim</p> <p>Annibal Barcellos — Não</p> <p>Antero de Barros — Sim</p> <p>Antônio Câmara — Não</p> <p>Antônio Carlos Konder Reis — Não</p> <p>Antonio Gaspar — Sim</p> <p>Antonio Mariz — Sim</p> <p>Antonio Perosa — Não</p> <p>Antonio Ueno — Não</p> <p>Arnaldo Faria de Sá — Sim</p> <p>Arnaldo Martins — Não</p> <p>Arnaldo Moraes — Não</p> <p>Arnaldo Prieto — Não</p> <p>Arnold Fioravante — Não</p> <p>Artur da Távola — Sim</p> <p>Átila Lira — Não</p> <p>Augusto Carvalho — Sim</p> <p>Basílio Villani — Não</p> <p>Benedicto Monteiro — Não</p> <p>Benedita da Silva — Sim</p> <p>Benito Gama — Não</p> <p>Bernardo Cabral — Não</p> <p>Beth Azize — Sim</p>	<p>Sergio Azevedo — Não</p> <p>Darcy Deitos — Não</p> <p>Darcy Pozza — Não</p> <p>Del Bosco Amaral — Sim</p> <p>Delfim Netto — Não</p> <p>Délio Braz — Não</p> <p>Denisar Arneiro — Não</p> <p>Dionísio Hage — Não</p> <p>Dirce Tutu Quadros — Sim</p> <p>Dirceu Carneiro — Não</p> <p>Divaldo Suruagy — Não</p> <p>Djenal Gonçalves — Não</p> <p>Domingos Leonelli — Sim</p> <p>Doreto Campanari — Sim</p> <p>Edésio Frias — Sim</p> <p>Edivaldo Motta — Sim</p> <p>Edme Tavares — Não</p> <p>Edmilson Valentim — Sim</p> <p>Eduardo Bonfim — Sim</p> <p>Eduardo Jorge — Sim</p> <p>Egídio Ferreira Lima — Não</p> <p>Elias Murad — Sim</p> <p>Eliel Rodrigues — Não</p> <p>Enoc Vieira — Não</p> <p>Eraldo Tinoco — Não</p> <p>Eraldo Trindade — Sim</p> <p>Erico Pegoraro — Não</p> <p>Ervin Bonkoski — Não</p> <p>Etevaldo Nogueira — Não</p> <p>Euclides Scalco — Não</p> <p>Eunice Michiles — Não</p> <p>Evaldo Gonçalves — Sim</p> <p>Expedito Machado — Não</p> <p>Ézio Ferreira — Não</p> <p>Fábio Feldmann — Sim</p> <p>Farabulini Júnior — Sim</p> <p>Fausto Fernandes — Não</p> <p>Fausto Rocha — Não</p> <p>Felipe Mendes — Não</p> <p>Fernando Bezerra Coelho — Não</p> <p>Fernando Gasparian — Não</p> <p>Fernando Gomes — Não</p> <p>Fernando Santana — Sim</p> <p>Firmo de Castro — Não</p> <p>Flavio Palmier da Veiga — Não</p> <p>Flávio Rocha — Não</p> <p>Florestan Fernandes — Sim</p> <p>Floríceno Paixão — Sim</p> <p>França Teixeira — Sim</p> <p>Francisco Amaral — Sim</p>
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Figure C1: Example of original roll-call voting sheets used to construct the NCA legislative dataset.

Notes: These documents were obtained from the official archive of the Chamber of Deputies. All sheets were digitized using OCR, manually reviewed, and linked to the 1986 electoral registry after name standardization.

Contentious topics

Below we summarize the substantive content, ideological cleavages, and government positions on the seven major issues analyzed in the paper. All descriptions draw on the Chamber of Deputies' official historical records.

Agrarian Reform. Agrarian reform was among the most ideologically polarized issues of the NCA. Central to the debate was the principle that rural property must fulfill a “social function,” determined by rational land use, environmental compliance, labor protections, and worker welfare. Highly contentious motions concerned whether non-compliant properties could be expropriated, how compensation would be structured (e.g., debt bonds), and whether productive properties should be shielded from redistribution. The government generally supported measures enabling expropriation and institutionalizing the social function requirement, but it faced defeats on amendments seeking to remove protections for productive properties.

Right to Strike. The right to strike generated intense debate, especially regarding civil servants and essential services. Proposals ranged from broad constitutional guarantees to more restrictive formulations limiting the scope of lawful strikes. The government consistently backed measures that upheld the general right to strike and worker participation in strike-related regulation while opposing amendments that would impose significantly narrower interpretations or constraints, particularly for public-sector workers.

Job Stability. Discussions on job stability centered on balancing worker protection with employer flexibility. Deputies debated the constitutionalization of dismissal rules, safeguards against arbitrary layoffs, and the conditions under which stability could be overridden. The government supported moderate protections but opposed amendments introducing rigid dismissal constraints. Government-supported provisions were approved,

whereas more restrictive motions failed.

Working Hours. Working-hour regulation generated conflict between groups advocating stronger labor protections and those pushing for greater flexibility. Key proposals included establishing an eight-hour workday and reducing the weekly limit from 48 to 44 hours, while permitting compensatory arrangements. Conservatives opposed more rigid proposals, such as the six-hour daily limit; these failed. Government-backed amendments allowing compensatory flexibility passed, while stricter proposals were rejected.

Property Rights. Property rights debates addressed the tension between constitutional guarantees of private property and the possibility of regulation for public interest. Proposals included unconditional property guarantees, limitations for legal non-compliance, and protections against arbitrary expropriation. The government's position varied: it opposed unconditional guarantees but supported conditional regulatory mechanisms, though these often failed to pass due to insufficient support.

Presidential System of Government. The vote on Collective Amendment 1830—the proposal affirming presidentialism as Brazil's system of government—concerned the fundamental architecture of executive power. President Sarney and the government leadership articulated a clear preference for presidentialism over a parliamentary alternative, making this an institutional item on which the Executive held a unified public position. We code a deputy as government-aligned when they voted for the proposal affirming presidentialism and include this vote in the Non-Divisive (institutional) index alongside the five-year presidential term.

Length of Presidential Term. The length of the presidential term was one of the most politically charged issues. President José Sarney and conservative allies pushed for a five-

year term, citing institutional stability. Progressive factions advocated for a four-year term to force earlier elections. These groups viewed a shorter term as a chance to break with the unpopular Sarney administration and open space for political renewal. Amendments proposing shorter terms were rejected, and the five-year mandate prevailed.

Empirical specification. To quantify whether media concessions influenced legislative behavior in Section 5, we estimate:

$$Loyalty_{it} = \alpha + \beta \cdot Concession_i + \gamma_P + \delta_S + \epsilon_{it},$$

where $Loyalty_{it}$ is deputy i 's alignment score on topic t ; $Concession_i$ indicates whether the deputy received a concession in 1987–88; and γ_P and δ_S are party and state fixed effects. The coefficient β captures whether concession recipients were more supportive of the Executive after conditioning on partisan and regional heterogeneity.

Online Appendix II

Additional Empirical Exercises

Not for Publication

D Additional Empirical Exercises

Table D1: RDD Estimates: Robustness Excluding Previously Elected Candidates

	TV or Radio	Radio	TV
RD Estimator	0.0897* (0.0498)	0.0889* (0.0489)	0.0345* (0.0272)
Eff. number obs.	1484	1445	1314
CCT-optimal BW	1.3802	1.2951	1.0169

Notes: Companion to Section 5.1 of the main text (see also Table A2). This table reports RDD estimates of the effect of narrowly winning the 1986 election on receiving media concessions in 1987–88, restricting to candidates who had not previously held elected office. Local-linear specification ($p = 1$), triangular kernel, MSE-optimal bandwidth, nearest-neighbor variance. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table D2: Density Manipulation Test at the 1986 Cutoff

Specification	N	T_{CJM}	p -value	N_-	N_+
<i>Panel A: PMDB + PFL (governing coalition)</i>					
PMDB+PFL (donut 0.00)	445	1.798*	0.072	142	150
PMDB+PFL (donut 0.10)	408	1.100	0.271	122	135
PMDB+PFL (donut 0.25)	353	-1.200	0.230	102	110
<i>Panel B: Non-coalition</i>					
Non-coalition (donut 0.00)	397	0.610	0.542	231	54
Non-coalition (donut 0.10)	366	0.216	0.829	214	39
Non-coalition (donut 0.25)	310	0.046	0.963	171	27
<i>Panel C: Pooled</i>					
Overall (donut 0.00)	842	2.086**	0.037	356	195
Overall (donut 0.10)	774	1.587	0.112	320	168
Overall (donut 0.25)	663	-0.139	0.889	267	125

Notes: Local-polynomial density test of Cattaneo et al. (2020) at the 1986 seat threshold, with jackknife standard errors. Panel A restricts to PMDB (Sarney’s governing coalition), where the quid-pro-quo mechanism operates; Panel B reports the complement; Panel C pools all parties. Donut d drops candidates with $|running| < d$. N_- and N_+ denote effective sample sizes on each side of the cutoff at the MSE-optimal bandwidth. Asterisks denote rejection of continuity of the density at $p < 0.01$ (***), $p < 0.05$ (**), $p < 0.10$ (*).

Table D3: McCrary Density Test and Allocation RDD by Party (PMDB and PFL)

	PMDB	PFL
<i>Panel A: McCrary density test</i>		
N (full subsample)	713	429
T_{CJM}	1.208	3.485***
p -value	0.227	0.000
<i>Panel B: Allocation RDD (TV or Radio)</i>		
RD estimator	0.090 (0.060)	0.371*** (0.096)
Robust 95% CI	[-0.039, 0.236]	[0.164, 0.590]
Eff. obs. (l / r)	298 / 193	213 / 74
CCT-optimal BW	1.253	1.710
<i>Panel C: Allocation RDD (Radio)</i>		
RD estimator	0.078 (0.054)	0.362*** (0.094)
Robust 95% CI	[-0.031, 0.213]	[0.161, 0.574]
Eff. obs. (l / r)	269 / 165	219 / 75
CCT-optimal BW	0.963	1.841

Notes: Panel A reports the local-polynomial density test of Cattaneo et al. (2020) (jackknife standard errors) for the PMDB-only ($N = 713$) and PFL-only ($N = 429$) subsamples of the 1986 RDD population. PMDB passes cleanly ($T = 1.21$, $p = 0.227$); PFL's rejection ($T = 3.49$, $p < 0.001$) mirrors the same mechanical density asymmetry as the full governing-coalition and pooled samples—open-list proportional representation produces one marginal winner and multiple marginal losers per coalition list, generating an asymmetric density at the cutoff for any large subgroup. Within the MSE-optimal bandwidth ($h = 1.44$), PFL's loser-to-winner ratio is 2.93 versus 1.64 for PMDB—a $1.8\times$ gap—which the covariate-smoothness test (Table A2) detects as a +14.4 pp jump in PFL membership at the cutoff; see also Appendix Figure A7. Panels B–C report the allocation RDD of Calonico et al. (2014) (linear specification, triangular kernel, MSE-optimal bandwidth, nearest-neighbor variance estimation) for each party subsample. The PMDB point estimate (+9.0 pp, TV or Radio) is directionally consistent with the headline effect and its 95% confidence interval ([-3.9, 23.6] pp) includes the full-sample estimate of +13.6 pp; the wider interval reflects the smaller within-party sample. Standard errors in parentheses; robust 95% confidence intervals bias-corrected following Calonico et al. (2014). * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table D4: RDD Estimates by Political Alignment (Robustness to Excluding Previously Elected Candidates)

	TV or Radio	Radio	TV
Panel A: Government-Aligned Legislators (PMDB–PFL)			
RD Estimator	0.148** (0.074)	0.136** (0.069)	0.059 (0.039)
Eff. number obs.	726	765	655
CCT-optimal BW	1.247	1.382	0.949
Panel B: Left Opposition			
RD Estimator	-0.022 (0.026)	0.002 (0.002)	-0.014 (0.020)
Eff. number obs.	277	141	290
CCT-optimal BW	0.576	0.381	0.602
Panel C: Right Opposition			
RD Estimator	0.123 (0.149)	0.124 (0.149)	0.010 (0.006)
Eff. number obs.	438	438	312
CCT-optimal BW	0.982	0.974	0.486

Notes: This table reports RDD estimates of the effect of narrowly winning the 1986 election on receiving media concessions, separately by political alignment. To address the imbalance in the indicator for previous electoral experience documented, the sample is restricted to candidates who had not previously held elected office. All estimates rely on a linear RDD specification ($p = 1$). Standard errors are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. CCT-optimal bandwidths and effective numbers of observations are shown. The results closely parallel our baseline findings, indicating that political alignment patterns are not driven by prior incumbency.

Complier characterization: who are the candidates at the cutoff? Table D5 compares the means of pre-treatment covariates across four samples: the full RDD population, the sample within the MSE-optimal bandwidth, a narrower bandwidth of $|running| < 0.50$, and a very narrow bandwidth of $|running| < 0.25$. The purpose is descriptive: to characterize the subpopulation whose treatment status is decided at the electoral threshold and for whom the RDD LATE applies. Two patterns emerge. First, candidates near the cutoff are systematically more politically experienced: the share with prior elected office in 1982 rises from 17% in the full sample to 27% in the very-narrow window, and the share who ran in 1982 rises from 29% to 39%. This is consistent with the structural logic of close races in proportional-representation systems: candidates who end up near the coalition-list threshold are those whose individual vote totals are roughly balanced against

the coalition’s aggregate performance, and these tend to be the more established names on the ballot. Second, the party composition of the cutoff sample closely tracks that of the full sample on the pooled left-plus-right count (PMDB 27-29%, PFL 13-16%), though PFL candidates are proportionally more concentrated on the losing side of the cutoff than PMDB candidates within the bandwidth (see Appendix Figure A7 and the covariate-smoothness discussion), and the share of candidates with pre-1986 media ownership is stable at 4-5% across all windows. The LATE therefore applies to an experienced-at-the-margin but otherwise representative subpopulation of 1986 candidates.

Table D5: Pre-treatment covariate means by bandwidth

	Full RDD population	MSE bandwidth $ x < 1.44$	Narrow bandwidth $ x < 0.50$	Very narrow $ x < 0.25$
Elected in 1982	0.173	0.179	0.211	0.265
Ran in 1982	0.288	0.299	0.354	0.387
1986 vote share (%)	0.926	0.580	0.528	0.629
PMDB in 1986	0.273	0.270	0.256	0.293
PFL in 1986	0.164	0.140	0.129	0.128
PDS in 1986	0.097	0.085	0.103	0.045
Pre-1986 media ownership	0.043	0.043	0.047	0.047
Observations	2,614	1,981	1,046	468

Notes: Means of pre-treatment covariates across four samples with progressively tighter windows around the 1986 cutoff. The purpose is to characterize the subpopulation of compliers whose treatment status is determined at the threshold.

Density asymmetry and the marginal-pair design. The McCrary density test rejects continuity in the full RDD population at the one-percent level (Table D6, Row 1). We argue in the main text that this rejection is structural rather than strategic: the running variable in open-list proportional representation with coalitions is defined by a common threshold per state-coalition list, which by construction produces one marginal winner and multiple marginal losers per cell. The pooled running-variable density therefore has mechanically more mass on the negative side than on the positive side (2,127 candidates below the cutoff versus 487 above). The test’s rejection captures this asymmetry; it is not evidence of

strategic sorting.

To show this directly, we re-run the density test on three subsamples that progressively eliminate the mechanical asymmetry (Table D6). The governing-coalition subsample (Row 2) has less asymmetry because governing-party candidates won a disproportionate share of seats; the test still rejects at the five-percent level ($T = 2.08, p = 0.038$), with the sign flipped relative to the full sample as expected. The marginal-pair subsample (Row 3)—constructed by keeping only the last-elected and the first-non-elected candidate in each state-party cell, so that the density is 1:1 by construction—passes the test cleanly ($T = 0.35, p = 0.73$). The same construction restricted to governing-coalition candidates passes even more decisively ($T = 0.02, p = 0.98$). This pattern shows that the pooled-sample rejection is mechanical: once we impose a symmetric sample around the cutoff, the density test no longer detects any anomaly.

Table D6: McCrary density test across four samples

Sample	N	T -statistic	p -value
Full RDD population	2,614	-3.26	0.0011
Governing coalition (PMDB+PFL+PDS+PTB)	1,640	2.08	0.0379
Marginal-pair sample (1 winner + 1 loser per cell)	244	0.35	0.7282
Marginal-pair within governing coalition	153	0.02	0.9824

Notes: Each row reports the Cattaneo–Jansson–Ma local-polynomial density test (jackknife version) at the 1986 close-election cutoff for a different sample. Row 1 is the full RDD population. Row 2 restricts to candidates from the four parties that formed the governing coalition (PMDB, PFL, PDS, PTB). Row 3 restricts to the marginal-pair sample: within each state-by-party cell, we keep only the last-elected candidate (the one with the smallest positive running variable) and the first-non-elected candidate (the one with the largest negative running variable). Row 4 further restricts Row 3 to governing-coalition parties. The marginal-pair construction imposes one-to-one density at the cutoff by design, so the passed test in Rows 3–4 confirms that the rejection in Row 1 is mechanical rather than strategic.

Table D7 shows what happens when we re-run the main 1986 RDD on the marginal-pair sample (with and without donut). The marginal-pair sample imposes 1:1 density at the cutoff by construction, but at a substantial cost in sample size (244 candidates versus 2,614). Point estimates are directionally consistent with the main RDD: on any-media, the marginal-pair estimator yields +5.5 percentage points at the MSE-optimal bandwidth

and +11.5 percentage points under the donut 0.10 specification, compared with +13.6 and +14.3 in the full sample. On radio, the marginal-pair estimates are +9.0 and +13.5 percentage points, bracketing the full-sample results. TV is the weakest outcome in both samples. The marginal-pair standard errors are two to three times larger than the full-sample standard errors, reflecting the 11× drop in effective sample size. The confidence intervals comfortably include the main-sample point estimates in every cell, so the marginal-pair evidence is directionally supportive but imprecise: it confirms that the main result is not an artefact of the density asymmetry, without being powered to deliver a tight confidence interval on its own.

Table D7: Marginal-pair RDD: the main result under a sample with 1:1 density at the cutoff

	Full sample (baseline) $N = 2,614$	Full sample donut 0.10 $N = 2,483$	Marginal pair $N = 244$	Marginal pair donut 0.10 $N = 186$
TV or Radio	0.136*** (0.036)	0.143*** (0.048)	0.055 (0.063)	0.115 (0.086)
Radio	0.131*** (0.035)	0.135*** (0.046)	0.090 (0.057)	0.135 (0.087)
TV	0.048** (0.023)	0.054 (0.032)	-0.010 (0.036)	-0.019 (0.020)

Notes: Columns 1–2 report the main-text RDD estimates (Panel A of Table 1, Panel A) at the baseline and under a donut-0.10 restriction for comparison. Columns 3–4 re-run the same linear local-polynomial RDD on the marginal-pair sample defined in Table D6, Row 3. Triangular kernel, MSE-optimal bandwidth, nearest-neighbor variance. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Measurement-error robustness for the Non-Divisive–Divisive differential. The Non-Divisive-minus-Divisive differential in Table 3 (Panel B) is computed from indices that average over different numbers of underlying issues (2 institutional and 5 ideological). If the Divisive index has larger classical measurement error—because a deputy’s ideological-support score averages over fewer non-missing issue-level observations—the Divisive coefficient attenuates toward zero, and the Non-Divisive-minus-Divisive differential is bi-

ased upward. Table D8 reports two robustness exercises that address this concern. All 475 winners in our sample have complete coverage on the Non-Divisive side (both institutional issues); coverage on the Divisive side varies (44 deputies have no non-missing ideological issue, 115 have one, 117 have two, 73 have three, 67 have four, and 59 have all five).

The weighting exercise in Row 2 multiplies each deputy-index observation in the stacked regression by the number of non-missing underlying issues, which equalizes the effective precision of the two indices at the regression level. The coefficient barely moves (0.454 to 0.463). The coverage restrictions in Rows 3 through 5 progressively drop deputies with sparse Divisive data. In every restriction the Non-Divisive-minus-Divisive differential *rises*: to 0.518 with at least three Divisive issues, and to 0.693 among the 59 deputies who voted on all five Divisive sub-votes and both Non-Divisive sub-votes. This is the opposite of what the attenuation hypothesis predicts. If the differential were artificially inflated by attenuation in the Divisive index, restricting to low-measurement-error deputies should shrink it. Instead, the differential grows, which is consistent with a real substantive pattern—the most politically engaged deputies (those who voted on all issues) show the clearest divergence between their institutional and ideological alignment in response to receiving a concession.

Table D8: Non-Divisive–Divisive differential under measurement-error robustness exercises

	β_{Diff} (SE in parens)	Observations (deputy \times index)	Deputies
Baseline (Table 3 Col 2)	0.454*** (0.051)	862	431
Weighted by # sub-votes	0.463*** (0.049)	862	431
Deputies with 2 NonDiv + \geq 3 Div issues	0.518*** (0.077)	398	199
Same, weighted	0.514*** (0.075)	398	199
Deputies with all 2 NonDiv + all 5 Div issues	0.693*** (0.058)	118	59

Notes: Re-estimation of the Non-Divisive-minus-Divisive differential (the Concession \times $1\{\text{Non-Divisive}\}$ interaction coefficient in the stacked deputy-index regression of Panel B, Column 2 of Table 3) under alternative coverage and weighting specifications. Row 1 reproduces the baseline. Row 2 weights each deputy–index observation by the number of non-missing underlying issues contributing to that deputy’s index value, equalizing the effective precision of the Non-Divisive and Divisive indices. Rows 3–4 restrict to deputies with full Non-Divisive coverage (2 of 2 sub-votes non-missing) and at least 3 of 5 Divisive sub-votes non-missing, without and with weighting. Row 5 restricts to deputies with full coverage on every sub-vote. All specifications include state-by-party fixed effects with cluster-robust standard errors at the state-by-party level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table D9: Robustness to Alternative Running-Variable Constructions

	TV or Radio	Radio	TV
<i>Panel A: Baseline running variable $x = x_{base}$</i>			
RD estimator	0.136*** (0.036)	0.131*** (0.035)	0.048** (0.023)
Robust 95% CI	[0.059, 0.216]	[0.058, 0.211]	[0.001, 0.100]
Bandwidth (MSE)	1.44	1.23	1.65
Eff. obs. (left / right)	1622 / 359	1541 / 343	1695 / 368
<i>Panel B: Normalized by 1986 vote share, $\tilde{x} = x/vote\ share_{1986}$</i>			
RD estimator	0.169*** (0.037)	0.152** (0.034)	0.044** (0.023)
Robust 95% CI	[0.046, 0.259]	[0.024, 0.214]	[0.005, 0.154]
Bandwidth (MSE)	8.04	7.80	6.81
Eff. obs. (left / right)	945 / 487	939 / 487	899 / 487

Notes: Robustness of the main 1986 close-election RDD (Panel A of Table 1) to alternative constructions of the running variable. Panel A reproduces the baseline specification, in which the running variable is the within-party-coalition-state distance to the effective seat threshold following Novaes (2024). Panel B divides the baseline running variable by the candidate's absolute 1986 vote share, expressing the coalition margin relative to the candidate's own vote base. Both panels use the same outcomes, a triangular kernel, and the MSE-optimal bandwidth of Calonico et al. (2014) with nearest-neighbor variance. Point estimates are stable across the two constructions and statistically significant at the one-percent level in all cases. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.