The Bureaucratic Politics of Networks: How Patronage Shapes Intergovernmental Collaboration

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Abstract

How does patronage—the political appointment of bureaucrats—affect the coordination and joint delivery of policies? Research has examined patronage's effects on bureaucratic performance, but mostly within hierarchical, top-down policymaking. However, in many domains, governance depends on horizontal networks of intergovernmental collaboration. I argue that patronage's effect on collaboration depends on the fit between the resources appointees carry and the demands of the positions they fill. Patronage facilitates coordination when appointees' political capital can be deployed for brokerage and negotiation but undermines it when loyalty-driven hiring displaces technical expertise and organizational stability. To test this argument, I analyze the network of environmental collaboration agreements among Colombian public agencies using hierarchical Exponential Random Graph Models. The evidence shows that managerial patronage fosters collaboration, especially along co-partisan lines, while professional-level patronage inhibits it. These findings underscore patronage's contingent role in governance networks and the importance of bureaucratic politics for collaborative policy delivery.

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1 Introduction

The assignment of public jobs on political grounds, referred to as patronage (Panizza et al. 2019), has been a central target of public sector reform worldwide because of its potentially deleterious effects on governance. By selecting loyal over competent personnel, political principals can increase their sway over bureaucratic decision-making (Gallo and Lewis 2011; Hollibaugh et al. 2014), facilitating preferential or illicit uses of public resources (Brierley 2020; Bussell 2019; Grindle 2012). Political control of bureaucratic hiring has consequently been associated with poorer government performance (Colonnelli et al. 2020; Lewis 2007; Rauch and Evans 2000) and increased public sector corruption (Dahlström et al. 2012; Meyer-Sahling et al. 2018; Oliveros and Schuster 2017). More recent scholarship has complicated this picture, showing that patronage can also enhance governance outcomes when politicians have the right incentives, boosting economic performance and improving service delivery (Jiang 2018; Toral 2024a).

However, these debates have developed around the largely implicit assumption that governance is organized hierarchically, with political principals directing bureaucratic agents within discrete organizational boundaries. This assumption poorly describes how governance increasingly works in practice, and especially in domains defined by common-pool resources such as environmental and natural resource management. In these settings, effective governance depends not only on isolated bureaucracies executing top-down mandates, but also on horizontal collaboration across jurisdictions and levels of government. Arrangements that depend on voluntary cooperation between autonomous government entities, such as shared services programs, memoranda of understanding, and joint venture agreements, have become widespread and deemed necessary to overcome the political and administrative fragmentation brought about by decentralization (Ansell and Gash 2008; Lubell et al. 2010; Ostrom 1990). How patronage shapes these horizontal modes of governance, where interdependent public organizations must cooperate voluntarily without clear relationships of political subordination, has been entirely sidelined.

In this paper, I develop a theoretical framework and new empirical evidence on how patronage shapes the emergence of collaborative policy delivery arrangements among government agencies. These arrangements entail collective action dilemmas that require aligning diverse organizational interests, negotiating resource dependencies, and building the institutions that facilitate joint problem-solving (Agranoff and McGuire 2001; Ostrom 1990; Provan and Kenis 2008). Whether and how such collaborations emerge depends on the incentives, relational resources, and technical capacities available within the participating organizations, all of which are shaped by how political incumbents staff their bureaucracies.

The central claim is that the effect of patronage on intergovernmental collaboration is governed by the functional fit between the resources that appointees carry and the demands of the positions they occupy. Politicians do not pursue uniform appointment strategies: they tend to place skilled, well-connected individuals in positions consequential for organizational performance while using lower-level posts to reward partisan supporters (Kopecký et al. 2016; Panizza et al. 2019). At managerial ranks, where the coordination demands center on brokerage, negotiation, and the ability to commit organizational resources, politically embedded appointees can be genuine assets because their cross-boundary networks, system-level institutional knowledge, and career incentives that reward visible coordination lower the transaction costs of interorganizational engagement. Importantly, because the political capital that managerial appointees deploy is rooted in party networks and reciprocal partisan obligations, the collaboration-facilitating effects of managerial patronage should be concentrated along co-partisan lines. Shared partisan affiliation reduces defection risk, lowers transaction costs, and activates the very channels through which appointees' relational capital operates most effectively (Fischer and Sciarini 2016; Kopecký et al. 2016).

In contrast, at professional and operational ranks, where the coordination demands center on technical expertise, procedural continuity, and implementation capacity, loyalty-driven hiring can introduce a functional mismatch: appointees who lack domain knowledge displace the specialized personnel on whom collaboration depends and introduce the organizational churn that erodes the institutional memory and credible commitments on which sustained cooperation rests. The effect of patronage on collaboration is therefore not uniformly positive or negative, but conditional on where it penetrates the bureaucratic hierarchy.

I test these arguments by analyzing all environmental collaborative projects among public agencies in Colombia between 2018 and 2024. Using text analysis methods, I identified and classified all public collaborative agreements ($N = 4{,}716$) signed between organizations from all levels of government, and matched them to fine-grained employer–employee public records containing information on the universe of civil servants in the country. Because tie formation in collaborative networks exhibits second- and third-order dependencies, I use Exponential Random Graph Models (ERGMs) with a hierarchical extension to account for regional heterogeneities and the embeddedness of collaboration dynamics in the national network. Consistent with the functional fit argument, I find that managerial patronage increases the likelihood of collaborative ties ($\approx$19% higher odds per standard deviation), while professional-level patronage decreases it ($\approx$30% lower odds). Further interaction analyses suggest that the collaboration premium associated with managerial appointments is concentrated among co-partisan partners and is largest when professional-level bureaucrats are organizationally stable.

This paper makes three contributions. First, it brings the politics of bureaucracy into conversation with the network and collaborative governance literatures, which have largely treated bureaucratic actors as background context rather than as politically constituted agents whose composition shapes horizontal coordination (Biesbroek and Lesnikowski 2018; Biesbroek et al. 2018). I argue that the politics of bureaucratic appointments structure the incentives, capacities, and relational resources with which agencies engage as collaborative partners. Second, it extends the patronage literature to the more recent domain of horizontal, decentralized policymaking, where the mechanisms through which patronage operates differ fundamentally from those in hierarchical settings (Brierley 2021; Brierley et al. 2022; Dahlström and Lapuente 2022). By specifying the positional logic of functional fit, the paper moves beyond documenting heterogeneous effects toward explaining why they arise. Third, it contributes to the political economy of environmental governance by showing that the interplay between the political and organizational dimensions of the bureaucracy is an important determinant of whether polycentric governance systems can coordinate around the common-pool problems they are designed to manage (Dijkstra and Fredriksson

2010; Fredriksson and Wollscheid 2014; Hu et al. 2021).

2 Theory

The central argument in this paper is that the way patronage shapes intergovernmental collaboration depends on where in the bureaucratic hierarchy it penetrates and the resources appointees can deploy. Appointments at different ranks carry systematically different resources that either match or mismatch the functional demands of collaboration in the positions they fill. When the fit is strong, patronage can become a network resource; when it is poor, it becomes a liability.

Bureaucratic politics scholarship has established that patronage is not a uniform practice of rewarding loyalty at the expense of quality. At upper levels of the hierarchy, appointees tend to be party professionals or programmatic technocrats selected for their capacity to design policy, coordinate across agencies, and navigate institutional complexity (Bersch et al. 2017; Kopecký et al. 2016). At lower levels, appointments more often serve reward functions, compensating supporters whose individual contribution to organizational performance is seen as marginal (Panizza et al. 2019). Incumbents consistently calibrate appointment strategies to the institutional significance of positions, investing in competence and connectivity at managerial levels while using lower ranks to sustain partisan commitments (Brierley 2021; Hollibaugh et al. 2014; Panizza et al. 2018). The following subsections elaborate how this dual appointment logic either enhances or limits collaborative capacity: managerial political capital can facilitate collaboration when deployed where it can be activated, while clientelistic hiring at operational ranks displaces the technical expertise and organizational stability on which collaborative projects depend.

2.1 Political Capital and Network Facilitation

The idea that politically embedded actors can serve as coordination resources has deep roots in the comparative politics literature. Classic accounts of state-building document how patron–client networks operated as informal coordination mechanisms linking national elites with subnational authorities in fragmented institutional contexts (Ang 2020; Geddes 1996; Grindle 2012; Huntington 1968). This has been the case of, for instance, Brazil's politicized but effective developmental state (Schneider 1991) and Colombia's clientelistic bureaucracies that substituted for weak state presence in peripheral regions (Leal Buitrago and Davila 1990). Recent work has extended these insights with causal evidence. Jiang (2018) shows that Chinese cities with politically connected leaders experience faster economic growth through patron–client coordination, while Jiang and Zhang (2020) demonstrate that such connections secure privileged access to national transfers. Toral (2024a) provides causal evidence from Brazil that patronage improves service delivery by creating "upward embeddedness" that reinforces mutual trust and enables more effective monitoring. These relational dynamics suggest that patronage connections can function as governance instruments structuring communication and coordination across fragmented policy arenas (Peters and Bianchi 2020).

Patronage studies have typically concentrated on the perspective of the patron, but appointed personnel are also strategic political actors whose incentives and behaviors within networks shape governance processes (Cornell and Grimes 2023; Oliveros 2021). I argue that there are at least three theoretical pathways through which appointees may lower barriers to intergovernmental cooperation: their distinctive network position, their system-level knowledge, and their career incentives.

First, political appointees occupy a distinctive structural position in governance networks. Research on networking patterns within government shows that politicians and political actors tend to be more active than career bureaucrats in building ties with external actors across organizational boundaries, particularly with actors in business, the community sector, and other levels of government (Alexander et al. 2011). Appointees can thus function as boundary-spanners: their networks of political connections span across domains and jurisdictions, endowing them with relational capital that career bureaucrats lack. This cross-boundary connectivity is precisely the resource that collaboration demands as intergovernmental networks depend on social capital that reduces uncertainty and the risks of discoordination or defection (Bodin 2017; Fischer and Sciarini 2016; Lubell

2013; Ostrom 1990), and politically embedded actors are well placed to supply it, deploying their position and political capital to search for and secure partners (Weible 2011).

Second, appointees typically accumulate system-level institutional knowledge—that is, knowledge about the governance system as a whole, its actors, rules, and available resources (see Vantaggiato and Lubell 2022)—that complements the domain-specific expertise of tenured staff. Because of their relatively frequent turnover across positions and agencies, managerial appointees can become governance generalists with experience across multiple organizational settings. Public administration research has demonstrated that such managerial mobility promotes innovation and policy adoption by aligning preferences across organizational boundaries and reducing transaction costs, since generalists know the characteristics of other actors and the workings of the system (Chen and Yi 2024; Huang and Berry 2021; Teodoro 2009, 2010). Strategically placed appointees, such as party professionals and programmatic technocrats who combine political networks with policy expertise, can thus serve as information brokers, channeling knowledge between organizations and providing the intelligence about partners, priorities, and available resources that collaboration requires (Bersch et al. 2022; Panizza et al. 2019).

Third, facilitating coordination aligns with the career incentives of political appointees themselves. Their professional trajectories depend on their patrons' political success and on their own reputation within the political–administrative system (Cornell and Grimes 2023; Oliveros 2021). Widening and strengthening professional networks is therefore not merely a byproduct of their work but a deliberate career strategy as visible collaboration with other agencies helps build reputation, demonstrates competence, and improves future appointment prospects (Ingold and Leifeld 2016; Teodoro 2010, 2013). In this sense, the individual-level logic of career advancement reinforces the organizational-level logic of coordination, making appointees willing (and not merely able) network facilitators.

An important nuance to these mechanisms is that, because the political capital appointees deploy is typically rooted in party networks, shared political principals, or reciprocal political obligations, the collaboration-facilitating effects of patronage should be especially pronounced along partisan lines. A well-established finding in network governance research is that political homophily—the tendency of politically aligned actors to collaborate—is a powerful structuring force in policy networks (Gerber et al. 2013; Henry 2023; Song et al. 2018). Collaboration with political allies reduces the risks of defection, facilitates informal bargaining, and lowers transaction costs because shared partisan affiliation generates baseline trust and a common political vocabulary (Fischer and Sciarini 2016; Kopecký et al. 2016). Political appointees are well placed to activate these homophilic channels because their access to party networks, knowledge of who holds office where, and ability to leverage their patron's authority all operate most effectively within partisan boundaries. If political capital is the primary resource appointees contribute to collaboration, its effect should therefore be concentrated among ties linking organizations whose appointments are made by the same or aligned parties. The corollary is that where collaboration requires bridging politically opposed organizations, the very partisan capital that facilitates coordination among allies may become a liability—signaling competing loyalties, raising suspicion, and inhibiting the emergence of cross-partisan ties.

2.2 The Costs of Politicization for Coordination

However, the relational capital of political appointees may not translate into collaborative behavior if patronage simultaneously erodes two bureaucratic attributes that underpin collaborative networks: specialized policy knowledge and organizational stability.

2.2.1 The loyalty–competence trade-off

Patronage literature has paid significant attention to the "ally principle," according to which politicians prefer to hire loyal servants to guarantee alignment and direct accountability (Dahlström and Lapuente 2022). Staffing the bureaucracy with loyal servants who lack policy-domain knowledge and do not have safeguards to "speak truth to power" implies a general reduction in the technical capacity of the bureaucracy to produce and process specialized policy knowledge. High levels of patronage also discourage investment in specialized expertise and can reduce the motivation of tenured bureaucrats to devote effort to their work, as career prospects and professional autonomy diminish (Fuenzalida and Riccucci 2019; Gallo and Lewis 2011; Lewis 2008).

This matters for collaboration because intergovernmental cooperation demands precisely such specialization. Knowledge is the "currency of collaboration" (Emerson et al. 2012, 16). It encompasses sharing, generating, and refining expertise to develop common solutions to collective problems (Agranoff and McGuire 2001; Ansell and Gash 2008; Klijn and Koppenjan 2012). In many policy domains, such as natural resource management, policymaking operates around wicked problems where complex domain interlinkages, cross-jurisdictional dynamics, and contested knowledge accompany high uncertainty. Intergovernmental networks benefit from high levels of technical expertise and field-specific experience that reduce the risks of failure-prone experimentation and increase the chances of policy learning (Leach et al. 2014; Vantaggiato and Lubell 2022). Experts play a crucial role in defining policy objectives, reducing potential coordination failures (Calvert 1992), and enhancing the quality of deliberation in policy forums by keeping discussions anchored to pre-established issues and oriented toward concrete actions (Ansell and Gash 2008). This policy knowledge also provides decision-making legitimacy, a resource typically guaranteed by a cadre of meritocratically selected personnel that belong to the same epistemic community (Haas 1992; Kolvani and Nistotskaya 2025; Mavrot and Sager 2018).

When patronage erodes this technical base, organizations become less attractive and less capable collaboration partners. In policy domains where technical credibility is a precondition for joint action, actors may avoid engaging with organizations that lack domain knowledge, as the risks of coordination failure and defection increase (Ansell and Gash 2008; Berardo and Scholz 2010). The relational capital of appointees, however valuable for brokerage and negotiation, cannot compensate for an organizational deficit in the specialized expertise that collaboration requires to move from search to agreement to implementation.

2.2.2 Bureaucratic turnover

Increased bureaucratic turnover is a well-established consequence of patronage. Election winners routinely reshape the bureaucracy upon taking office, disrupting teams, displacing job-specific experience, and degrading service delivery (Akhtari et al. 2022; Brassiolo et al. 2020; Toral 2024b). In politicized administrations, political transitions intensify this churn as politically motivated personnel leave anticipating conflicts with incoming incumbents, and elected officials refill public ranks with their own supporters (Brassiolo et al. 2021; Doherty et al. 2019). The heightened turnover produced by patronage disrupts organizational learning and institutional memory (Bagchi and Chakrabarti 2021; Pollitt 2009; Rao and Argote 2006; Stark and Head 2018).

This is particularly problematic in horizontal governance contexts, because collaboration is inherently dependent on continuous experimentation, learning, and mutual adaptation through repeated interactions (Gerlak et al. 2019; Howlett et al. 2017; Rittelmeyer et al. 2024; Siciliano 2017; Vantaggiato 2019). When turnover erodes the technical capacity and stability that organizational learning requires (Corbett et al. 2018; Geys et al. 2023), the specialized expertise accumulated within agencies is lost, alongside the standardized procedures and bureaucratic routines on which collaboration depends. Network governance theory has consistently argued that institutional stability is a prerequisite for governance systems to survive and mature. Stability enables shared knowledge, mutual learning, and the credible commitments on which sustained collaboration rests (Lubell 2013; McGinnis et al. 2020). Patronage-driven bureaucracies exhibit higher rates of institutional memory loss and greater decision-making uncertainty for potential partners (Pollitt 2009), making both the formation of new collaborative ties and the continuity of existing ones less likely.

As discussed above, the rotation of individual managers across positions can contribute to coordination by endowing mobile administrators with system-wide knowledge, cross-organizational networks, and professional reputation (Teodoro 2009, 2013; Yi et al. 2018). However, when turnover occurs at scale, reshaping entire organizational cadres rather than rotating individual leaders, the aggregate effect shifts from network facilitation to institutional destabilization. The distinction is between selective mobility, in which individual leaders carry knowledge and connections across organizational boundaries, and systemic churn, in which the wholesale replacement of personnel erodes the organizational foundations on which collaboration depends. The former enriches the relational capital available for coordination; the latter undermines the institutional continuity that makes coordination possible.

2.3 Synthesis: Bureaucratic Hierarchy, Coordination, and Functional Fit

The opposing mechanisms map directly onto the dual appointment logic identified above. The core insight is one of functional fit: patronage shapes collaboration through the interaction between the resources appointees bring and the demands of the positions they fill.

At managerial levels, coordination needs center on discretion, institutional navigation, and the ability to identify partners, negotiate agreements, and commit organizational resources (Agranoff and McGuire 2001; Scott and Thomas 2017). These are functions for which the political networks, system-level knowledge, and brokerage skills of strategically appointed individuals are well suited. Managerial appointees can leverage their cross-boundary connectivity to search for and secure collaboration partners, deploy their system-wide knowledge to reduce transaction costs of interorganizational engagement, and draw on career incentives that reward visible networking and coordination. Importantly, the partisan capital embedded in these appointments implies that their primary channels to lower defection risk are realized among politically aligned organizations.

At professional and operational levels, the coordination needs are fundamentally different: technical expertise, procedural continuity, and the capacity to design, implement, and monitor collaborative projects (Corbett et al. 2018; Emerson et al. 2012). Appointments at these ranks, which typically follow clientelistic dynamics, introduce a functional mismatch. Appointees who lack domain knowledge displace the specialized personnel on whom collaboration depends and, even if well-connected politically, cannot easily leverage relational resources from positions that afford little discretion or institutional authority. When this displacement occurs at scale, it compounds into the systemic churn that erodes organizational learning, institutional memory, and the capacity to sustain credible commitments over time (Akhtari et al. 2022; Pollitt 2009).

3 Case Selection and Institutional Context: Environmental Governance Networks in Colombia

To test the theoretical framework developed above, I study collaborative governance networks in the environmental domain in Colombia. The choice of context reflects two distinct case selection logics—one concerning the policy domain, the other the country.

Environmental governance represents a paradigmatic policy domain (Flyvbjerg 2006) for building theory about the bureaucratic politics of intergovernmental networks. The collective action tensions of governance are particularly salient in this domain because the goods at stake (e.g., watersheds, forests, biodiversity, climate regulation) are archetypal common-pool resources whose ecological boundaries rarely align with administrative jurisdictions (Bodin 2017; Ostrom 1990), making effective governance structurally dependent on coordination across fragmented, polycentric systems. Furthermore, environmental resources are deeply politically contested. Property rights, land use, extraction permits, and pollution regulation directly implicate distributive conflicts and electoral constituencies, making the partisan and relational capital embedded in bureaucratic appointments potentially central to how governance processes unfold (Dijkstra and Fredriksson 2010; Duit et al. 2016; Fredriksson and Wollscheid 2014). If patronage shapes intergovernmental collaboration anywhere, the effects should be most theoretically legible in the environmental domain.

Within this domain, Colombia represents a typical case (Seawright and Gerring 2008) of countries governing environmental policy under fragmented authority and mixed bureaucratic personnel systems. A pioneer in Latin American environmental decentralization, Colombia devolved significant environmental authority to subnational actors through the Constitution of 1991, extended through successive legal reforms over the past three decades. The resulting governance structure distributes environmental authority across four main levels: ten national-level environmental authorities, 27 *departamentos* acting as primary regional environmental authorities, 34 subregional autonomous environmental corporations (*Corporaciones Autónomas Regionales*) spanning multiple municipalities and responsible for monitoring, enforcement, and permit issuing across jurisdictional lines, and 1,103 municipalities, with six formally constituted metropolitan areas acting as inter-municipal environmental authorities.

The highly decentralized governance architecture, the heterogeneous distribution of state capacity, and the prevalence of cross-boundary ecosystems make intergovernmental collaboration a functional necessity for the governance of the environment. In response, Colombian governments have widely experimented with collaborative forms of environmental management during the last decades. Joint actions between public agencies are formalized through *convenios interadministrativos*, which require the mutual and voluntary commitment of resources by all parties, making them negotiation-intensive and costly signals of genuine collaborative intent. They form part of Colombia's public procurement system and are subject to strict legal monitoring and fiscal oversight, closely resembling the intergovernmental agreements widely used by US agencies. These instruments provide a legally comparable, administratively costly, and therefore credible unit of analysis for studying collaborative behavior across the entire national system.

Regarding bureaucratic institutions, Colombia's public sector combines professionalization of the civil service career with pervasive politicization (Ayala-García et al. 2022; Sanabria-Pulido and Leyva 2023). This coexistence of meritocratic and political appointment tracks is a stable institutional equilibrium broadly characteristic of Latin American and other emerging democracies (Cortázar Velarde et al. 2014; Kopecký et al. 2016). Career civil servants are selected through the *concurso público de méritos*, a competitive merit-based process centrally administered by the national government that applies to all tenured career positions across all levels of government. Simultaneously, elected officials at the municipality and *departamento* level retain extensive discretion to make at-will appointments (*libre nombramiento y remoción*), usually for managerial and high-level positions, though these extend to lower bureaucratic ranks as well. The resulting variation in bureaucratic composition is substantial. The share of at-will political appointees ranges from under 10% to near-total dominance across agencies, varying across ranks, jurisdictions, and even municipalities within the same metropolitan area. I leverage this organizational-level variation to study how patronage profiles shape the likelihood of intergovernmental collaboration.

In sum, Colombia combines a highly decentralized and polycentric environmental governance structure with a degree of bureaucratic politicization frequent among middle-income democracies, reflecting the modal conditions under which many Global South states currently govern environmental policy. The findings are thus most directly applicable to political systems characterized by meaningful environmental decentralization, mixed personnel systems, and reliance on horizontal policymaking mechanisms—conditions increasingly common across Latin America, sub-Saharan Africa, Eastern Europe, and Asia-Pacific (Kopecký et al. 2016; Meyer-Sahling et al. 2018; Moon and Hwang 2013).

4 Data and Measurement

4.1 Data

I leverage several administrative datasets to identify the collaborative networks and bureaucratic characteristics of public agencies in Colombia. The primary source is the full record of contracts signed by public entities between 2018 and 2024 in the national procurement system, SECOP II. Each entry contains details including contract type, contracting parties, monetary value, and a textual description of the agreement's purpose. From nearly four million contracts, I identified all *convenios interadministrativos* related to environmental and natural resource management using a two-stage procedure.

The first stage applied keyword queries based on terms from Colombian public law and administrative practice that typically denote inter-administrative agreements. The keyword set was compiled through fieldwork and interviews with public lawyers and experienced bureaucrats, then expanded iteratively from contracts already identified, yielding an initial dataset of 31,064 contracts signed between 2016 and 2024. Because this search was intentionally broad to maximize recall, the second stage used Large Language Models (LLMs) to distinguish environmental collaborative agreements from unrelated contracts. Contract descriptions were classified using an iterative prompt-engineering procedure applied to the Mistral Large 2411 model, validated against a manually labelled ground-truth sample of 2,000 randomly selected contracts. The final prompt was a few-shot specification that defined the environmental domain, enumerated relevant subsectors, and provided positive and negative exemplars, achieving an F1 score of 0.985 on the full validation set. I corrected residual misclassifications manually. Full details of the classification procedure, prompt specifications, and performance evaluation across model and temperature settings are reported in Appendix A1 in the Supporting Information (SI; Tables A1–A3).

The final dataset consists of 4,716 environmental collaboration agreements signed across all levels of government. From these data, I construct a national, undirected network of environmental collaborations, where nodes are public agencies and ties are defined as formal bilateral agreements between them. Agencies include municipalities, *departamentos*, national ministries, public universities, research centers, development agencies, and utility companies, among others. These agreements span all relevant dimensions of environmental governance and cover a wide range of collaborative activities, from a joint initiative between two municipalities to improve natural resource management through community-level programs, to a bilateral agreement between a regional environmental authority and a local government to design flood and landslide mitigation infrastructure. Importantly, I also include the universe of agencies with a legal mandate in environmental governance that did not sign any agreements during the period, representing them as isolates, to prevent bias from conditioning the analysis only on observed collaborations.

The resulting network contains 1,804 organizations (nodes) and 1,949 ties (edges), with 592 agencies (33%) as isolates. Network density is very low at 0.0012, meaning only a small fraction of all possible ties is realized. Agencies have an average of 4.32 agreements ($SD = 12.82$), a median of 2, and a strongly right-skewed degree distribution (skewness = 10.98), indicating that most agencies participate in very few collaborations while a small number act as hubs. The clustering coefficient is 0.033, reflecting minimal transitivity overall and suggesting that shared partners do not routinely convert into closed triangles. Together, these features describe a sparse network with pronounced preferential attachment dynamics, consistent with common empirical findings in polycentric environmental governance settings (Bodin and Crona 2009). Figure 1 depicts a subsample of the network including organizations from the three largest *departamentos*, illustrating both the hub-and-spoke structure and the tendency for intra-regional clustering.

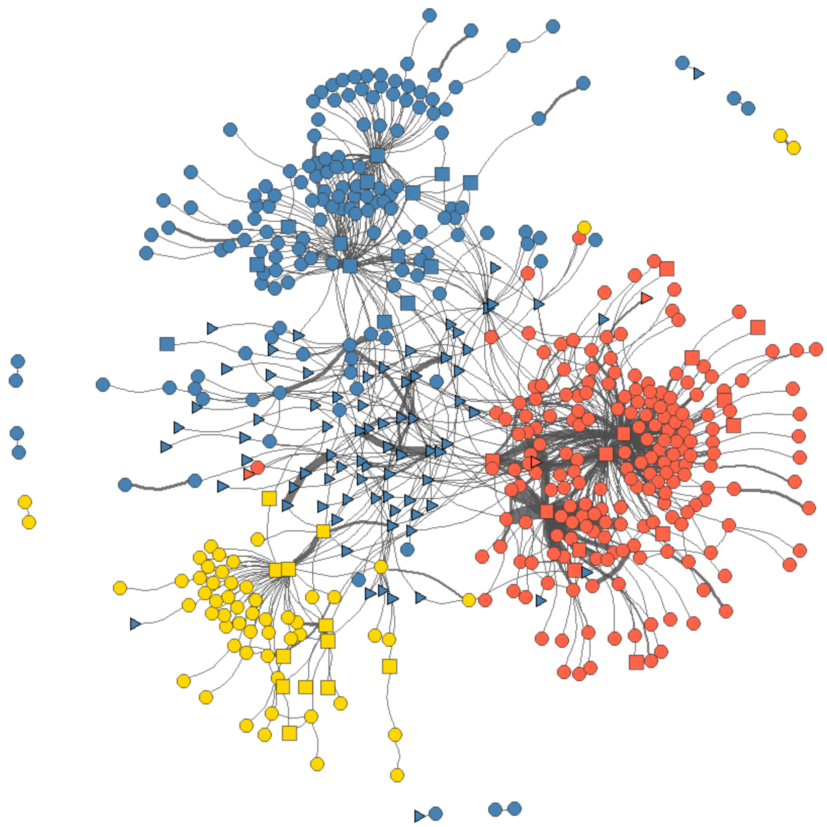

Figure 1: Environmental Collaboration Network (Top three *departamentos*). The graph shows a subnetwork of organizations from the three largest *departamentos* in Colombia: Cundinamarca (including Bogotá, blue), Antioquia (red), and Valle del Cauca (yellow). Node shapes indicate the tier of government—local (circle), regional (square), and national (triangle). Edge thickness reflects the number of collaborations between two organizations (thicker = more collaborations). Isolate nodes were removed from the graph to improve clarity (but not from the analysis).

I match the network data with the Public Employment Information and Management System (SIGEP), which records detailed information on all public employees, including demographics, education, tenure, salary, contract type, and prior experience. The SIGEP data was published and made fully available only for 2022, which constrains the analysis to be cross-sectional. Missing values in organizational covariates are addressed using denoising autoencoders, a machine learning approach to multiple imputation developed by Lall and Robinson (2022), and all variables are then aggregated to the agency level. Appendix A2 in the SI reports the imputation procedure and performance diagnostics (Figure A1).

4.2 Measurement and Variables

Patronage. The main independent variable is patronage, measured as the share of an agency's personnel holding at-will (*libre nombramiento y remoción*) positions—appointments made at the discretion of an elected official without competitive merit selection. This operationalization maps directly onto the extent to which an agency's workforce is politically rather than meritocratically staffed.

Given the granularity of the personnel records, I disaggregate this measure by bureaucratic rank. Managerial patronage captures the share of directive and top positions filled through at-will appointment, including subnational cabinet secretaries (*secretarios de despacho*), agency directors, and subdirectors. Professional patronage refers to the share of mid- and low-rank professional and operational positions filled through at-will contracts. These include personnel such as engineers, legal advisors, field inspectors, GIS technicians, and budget analysts, whose primary function is to design, implement, and monitor policy rather than to direct the agencies. Summary statistics reported in Appendix A4 in the SI (Tables A4–A8) illustrate the substantial cross-agency variation in each measure. Overall patronage ranges from 0 to 100%, with managerial positions exhibiting considerably higher average politicization (mean: 75%) than professional-level ones (mean: 18%).

Bureaucratic attributes. Three measures of bureaucratic attributes relevant to the theory are included: technical capacity, bureaucratic stability, and average public sector experience. Technical capacity is proxied by the share of employees with postgraduate education, a consistent predictor of domain-specific expertise widely used in comparative research (Bednar 2024; Hollibaugh 2018). Bureaucratic stability is measured as the mean expected total tenure of an agency's personnel. The key challenge is that cross-sectional administrative records capture each employee's entry date but not their exit, so using observed tenure alone would systematically underestimate organizational continuity due to right-censoring. To address this, I estimate Kaplan–Meier survival curves stratified by contract type and use them to project each employee's expected remaining service conditional on time already served. These projections are averaged at the agency level, yielding a measure of expected workforce stability rather than merely observed seniority. Full technical details are reported in Appendix A3 in the SI (Figures A2–A3). Finally, public sector experience is proxied by the average years of service of staff across their entire careers in the public sector. All three measures are computed at the agency level and also disaggregated by bureaucratic rank.

Controls. Several controls drawn from the governance networks literature are included. Agency size, measured as the log of total employees, captures organizational capacity as a well-established driver of collaboration (Krause et al. 2021; Siciliano and Wukich 2016; Vantaggiato 2019). The statutory *categoría* of each jurisdiction (a legal classification based on population and fiscal revenues) is included as a set of indicators, functioning both as an institutional marker of jurisdictional standing and as a proxy for local fiscal capacity. To capture environmental pressure, I include the number of people affected by climate-related events such as floods, droughts, or landslides (inverse sine transformed to reduce skewness while preserving zeros). Geographic distance between agencies (inverse sine transformed) controls for the transaction cost reduction that proximity affords (Baldwin et al. 2018; Lubell et al. 2002). All continuous variables are z-standardized to mean 0 and $SD = 1$. Variable correlations are presented in Figure A4 in the SI.

Homophily terms. Four homophily terms reflecting institutional and political structuring forces are included (McPherson et al. 2001). A government tier homophily term tests whether agencies preferentially partner across administrative levels. An organization type homophily term accounts for functional similarity. A *departamento* homophily term captures the tendency for intra-regional collaboration driven by shared ecological pressures, common institutional histories, and regionally bounded environmental mandates. Most central to the theoretical argument is political homophily, which codes two agencies as aligned if their staff are appointed by incumbents from the same political party, testing the well-established finding that shared partisan affiliation reduces defection risk and lowers coordination costs in policy networks (Gerber et al. 2013; Henry 2023; Song et al. 2018).

Endogenous network terms. A distinctive feature of collaborative networks is that tie formation is not independent across dyads: whether two agencies collaborate is shaped not only by their individual attributes but also by the broader structure of relationships in which they are embedded. Standard regression approaches cannot accommodate this interdependence, which is why the ERGM framework (explained below) is employed (Cranmer et al. 2017; Scott and Ulibarri 2019). Two structural terms are included.

The first is a geometrically weighted degree term (GWDEG, decay = 1), which models the skewed degree distribution evident in the descriptive statistics. Substantively, agency degree reflects accumulated reputation, administrative visibility, and demonstrated capacity for managing inter-organizational relationships. Including this term prevents differences in organizational prominence from inflating the estimated effects of patronage. It also controls for the large number of isolates.

The second is a geometrically weighted edgewise shared partner term (GWESP, decay = 0.5), which captures the propensity of two organizations to collaborate when they share a common partner (i.e., triadic closure). In governance networks, triadic closure reflects trust diffusion and risk reduction: a shared partner provides both actors with information about each other's reliability and creates mutual accountability (Berardo and Scholz 2010). Including this term controls for the fact that some agencies are more likely to collaborate simply because they operate in denser relational clusters, and prevents patronage estimates from being confounded by the tendency of politically aligned agencies to cluster in the same regional networks or environmental subdomains.

4.3 Statistical Approach

Relational data violate the independence assumptions of standard regression, producing biased estimates and over-optimistic p-values. I therefore employ Exponential Random Graph Models

(ERGMs), a family of stochastic models that estimate the probability of observing a given network as a function of sufficient statistics summarizing both structural endogenous dependencies and exogenous covariate effects (Cranmer and Desmarais 2010; Snijders et al. 2006). Formally, for an observed network $\mathbf{Y}$ with adjacency entries $Y_{ij} \in \{0,1\}$, the distribution is given by

$$\Pr(\mathbf{Y} = \mathbf{y} \mid \theta) = \frac{\exp\bigl(\theta^{\top}\mathbf{s}(\mathbf{y})\bigr)}{\kappa(\theta)}, \tag{1}$$

where $\mathbf{s}(\mathbf{y})$ includes statistics for endogenous process and covariate effects, θ is the corresponding parameter vector, and $\kappa(\theta)$ is the normalizing factor ensuring a proper probability mass function, approximated using simulation-based methods. Conceptually, ERGMs can be read as logistic regressions of tie formation, where the probability of a collaboration between two agencies depends on the changes in the specified network statistics (Desmarais and Cranmer 2012).

The national scope of the data introduces a key modelling challenge. The standard ERGM assumes homogeneity across actors once covariates are included, which is unrealistic here because agency behavior is strongly conditioned by the *departamento* in which organizations are embedded. *Departamentos* generate systematic differences in tie formation through variation in administrative capacity, subnational political dynamics, and traditions of coordination. Ignoring this structure risks attributing regional heterogeneity to endogenous network processes and inducing bias in the estimates (Duxbury 2023). Including *departamento* fixed effects is equally problematic because *departamentos* vary enormously in size (from 3 to 227 agencies), so fixed effects would leave estimation in sparsely populated regions relying on minimal within-region variation, producing unstable estimates and misestimation of the structural network terms.

To address this, I adapt the iterative procedure proposed by Kevork and Kauermann (2022) to introduce group-level random effects into the ERGM framework, estimating hierarchical mixed ERGMs (mERGMs) with *departamento*-level random intercepts. This captures the shared collaboration propensity of agencies within the same region while preserving the national scope of analysis through partial pooling. Region-specific intercepts are treated as draws from a common distribution, allowing *departamentos* with little information to borrow strength from those with more, while still capturing systematic regional heterogeneity. Letting g_i denote the *departamento* of agency i, the probability of the observed network conditional on both structural parameters and regional random effects is

$$\Pr(\mathbf{Y} = \mathbf{y} \mid \theta, \alpha) = \frac{\exp\Big(\theta^\top \mathbf{s}(\mathbf{y}) + \sum_{i<j}(\alpha_{g_i} + \alpha_{g_j})\, y_{ij}\Big)}{\kappa(\theta, \alpha)}. \tag{2}$$

This is an extension of the canonical form in Equation 1. Here, α is a g-dimensional vector of group-specific random effects assumed to follow $\alpha \sim \mathcal{N}(\mathbf{0}, \sigma_\alpha^2 \mathbf{I}_g)$, where σ_α^2 is the variance and $\mathbf{I}_g$ represents a g-dimensional identity matrix. From Equation 2 we can obtain the model for each tie Y_{ij} conditional on all other ties in the network, $\mathbf{Y}_{-ij}$, as follows:

$$\log \frac{\Pr(Y_{ij} = 1 \mid \mathbf{Y}_{-ij}, \theta, \alpha)}{\Pr(Y_{ij} = 0 \mid \mathbf{Y}_{-ij}, \theta, \alpha)} = \theta^\top \Delta_{ij} \mathbf{s}(\mathbf{y}) + \alpha_{g_i} + \alpha_{g_j}, \tag{3}$$

where $\Delta_{ij}\mathbf{s}(\mathbf{y})$ is the vector of change statistics associated with toggling the edge between agencies i and j. By additively including the regional random effects for each node, this formulation captures unobserved regional heterogeneity while still modelling endogenous dependencies through standard ERGM terms. Following Kevork and Kauermann (2022), random effects are estimated via Laplace approximation and supplied as fixed offsets to the ERGM likelihood step, estimated by MCMLE. The two steps alternate until convergence. Full derivation, estimation details, model stability diagnostics, and comparisons against standard ERGM specifications are reported in Appendices A5–A7 in the SI (Figures A5–A8; Tables A9–A10).

The baseline models establish the direct associations between patronage and collaborative tie formation. Interaction terms between patronage measures and key bureaucratic attributes (i.e., stability, technical capacity, and public sector experience) are then added to probe whether the effect of patronage varies with the organizational conditions that the theory identifies as either enabling or constraining the conversion of political capital into collaborative capacity. This is necessarily an indirect test, as proper identification of the causal pathways would require individual-level data on appointee behavior, network positions, and information flows beyond the scope of the present analysis. The interaction models should therefore be read as evidence of theoretical coherence rather than causal decomposition.

Relatedly, a note on inferential scope is warranted. Patronage levels are not randomly assigned, and no institutional discontinuity provides a clean source of exogenous variation in bureaucratic composition. The analyses should therefore be understood as estimating systematic associations between patronage profiles and collaborative behavior, rather than average treatment effects in the experimental sense. Remaining limitations and the conditions under which stronger identification would be possible are discussed below.

5 Results

5.1 Do Political Appointees Differ across Ranks and from Permanent Bureaucrats?

Before turning to the network models, I report baseline differences in bureaucratic attributes between political appointees and permanent career staff across managerial and professional ranks (Figure 2), and the within-rank differences (Figure 3).

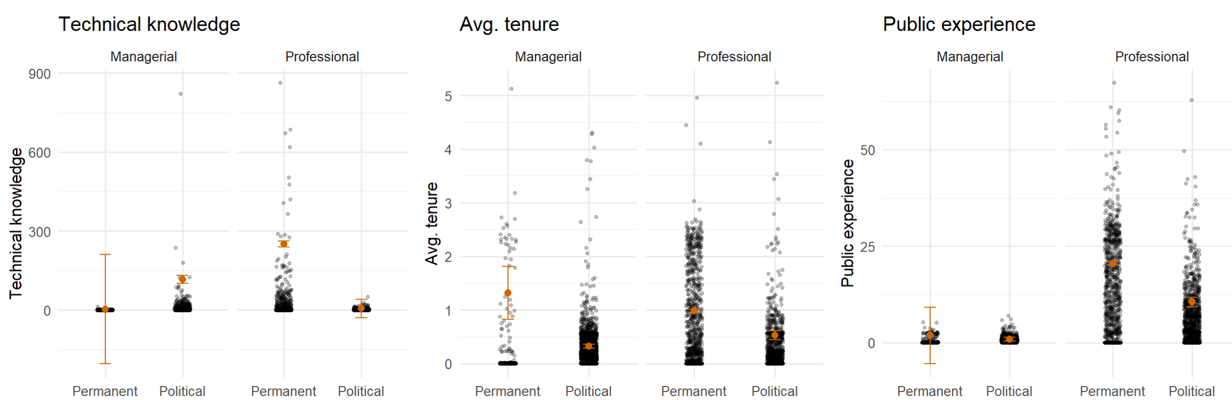

Figure 2: Organization means by rank and type of appointment. Estimated group means and 95% CIs in orange. CIs that exclude 0 indicate statistically significant effects.

The patterns are consistent with the dual appointment logic outlined above. At the professional rank, politically appointed staff are systematically less qualified than their permanent counterparts. They hold postgraduate degrees at lower rates, have accumulated less public sector experience, and exhibit shorter expected tenures. At the managerial rank, the picture is more nuanced. Appointees show shorter expected tenures than permanent staff, consistent with the heightened turnover that political appointments produce, but no significant differences emerge in postgraduate qualifications or public sector experience, suggesting that incumbents exercise greater selectivity in filling top posts. Whether and how these contrasting profiles translate into differences in collaborative behavior is the question the network models address.

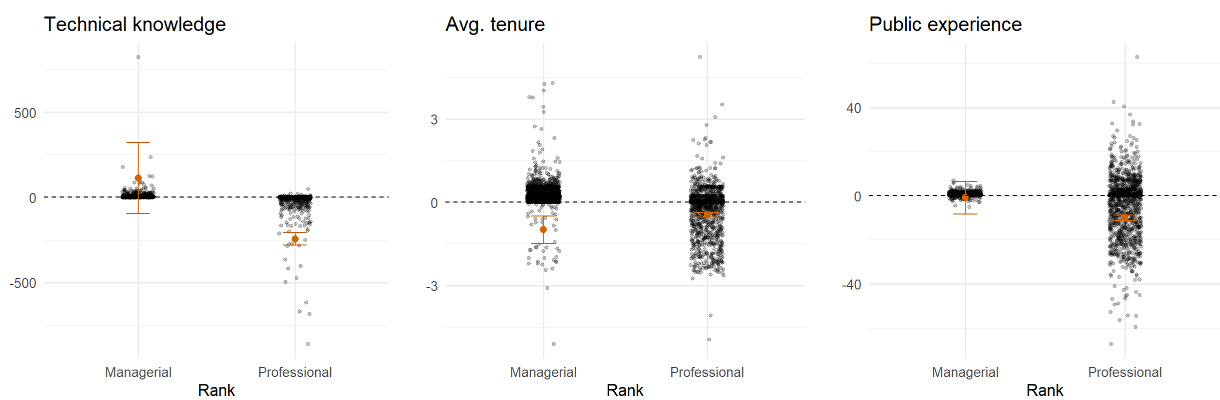

Figure 3: Within-rank differences (patronage appointments vs. permanent staff). Estimated group means and 95% CIs in orange. CIs that exclude 0 indicate statistically significant effects. Points below the zero line indicate higher values of the attribute for permanent career staff.

5.2 Main Network Models

Table 1 reports the mERGM estimates across four model specifications. Coefficients are interpretable as log-odds of tie formation and can be exponentiated to obtain multiplicative effects on the odds of a collaboration. Model 1 presents aggregate patronage and bureaucratic attributes; Models 2 and 3 disaggregate these by managerial and professional rank; Model 4 includes both sets simultaneously.

Table 1: Main effects: mixed Exponential Random Graph Models with *departamento* random effects

	Model 1	**Model 2**	**Model 3**	**Model 4**
Aggregate measures				
Patronage	−0.008 (0.040)			
Technical capacity	−0.184*** (0.035)			
Stability	−0.902*** (0.048)			
Public experience	1.046*** (0.050)			
Managerial-rank measures				
Man. patronage		0.155*** (0.035)		0.175*** (0.035)
Man. technical capacity		−0.253*** (0.037)		−0.199*** (0.038)
Man. stability		−0.503*** (0.040)		−0.512*** (0.042)
Man. public experience		0.424*** (0.032)		0.412*** (0.032)
Professional-rank measures				
Prof. patronage			−0.354*** (0.037)	−0.359*** (0.039)
Prof. technical capacity			−0.167*** (0.040)	−0.212*** (0.040)
Prof. stability			−0.102 (0.071)	0.038 (0.072)
Prof. public experience			0.128 (0.067)	0.044 (0.068)
Controls	✓	✓	✓	✓
AIC	26,241	26,715	26,122	22,557
Adjusted AIC	16,829	15,756	15,588	15,427
$\sigma_{\text{Dept.}}$	2.74	2.62	2.56	2.61
Pseudo-ICC	0.22	0.21	0.21	0.21
REs	32	32	32	32

Note: $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$. Coefficients are the conditional log-odds change in the probability of a tie. Standard deviations in parentheses. All continuous variables are z-standardized to mean 0 and $SD = 1$. $\sigma_{\text{Dept.}}$ is the standard deviation of the estimated region-level random effects on the log-odds scale. Pseudo-ICC is the share of residual variance in tie formation attributable to between-region clustering; Appendix A6 in the SI explains in detail how these parameters are calculated. AIC is the standard Akaike Information Criterion as calculated with the `ergm` package. The adjusted AIC is computed via network simulation with a penalty for the effective degrees of freedom of the group random effects; Appendix A7 explains this adjustment and compares the results (Table A10). Table A9 presents VIFs for the variables in the fully saturated Model 4 to assess collinearity. Full control coefficients, including the baseline model, are reported in Table A11.

Main patronage effects. In Model 1, the aggregate patronage coefficient is small and indistinguishable from zero, suggesting no net association with collaboration. Disaggregation across ranks in Models 2–4 reveals sharp and opposing rank-specific patterns that the aggregate obscures. Managerial patronage is consistently positive: a one-standard-deviation increase in the share of politically appointed managers corresponds to approximately 19% higher odds of a collaborative tie ($\exp(0.175) \approx 1.19$, $p < 0.001$). Professional patronage is consistently negative: the equivalent increase at the operational rank corresponds to roughly 30% lower odds ($\exp(-0.359) \approx 0.70$, $p < 0.001$). Organizations with larger shares of political appointments at the top show higher collaborative engagement, while those where patronage penetrates operational ranks show the opposing pattern. These estimates are stable whether the rank-level measures are entered separately or jointly, and are robust to the inclusion of the full control set.

Bureaucratic attributes. Technical capacity is negatively associated with collaboration across all specifications ($\exp(-0.2) \approx 0.84$, $p < 0.001$), and managerial stability shows a similar negative association ($\exp(-0.51) \approx 0.60$, $p < 0.001$), while public sector experience is a strong positive correlate at the managerial level ($\exp(0.41) \approx 1.51$, $p < 0.001$). Neither stability nor experience reaches significance at the professional rank. These patterns are worth interpreting carefully. Everything else constant, organizations with deeper in-house expertise appear to be less inclined to form collaborative ties, consistent with the notion that less of what collaboration would potentially provide is missing internally. The managerial patterns point to a complementary dynamic. Managers with broader public sector career histories may carry the institutional knowledge and interpersonal networks that correspond to lower search and negotiation costs of collaboration, and that association is most pronounced where managerial rotation is higher—a pattern that would account for the opposing signs on experience and stability.

The null effects at the professional rank further reinforce the functional-fit argument. The rotation premium implied by the negative managerial stability coefficient does not extend to professional staff, whose positions afford less discretion to translate mobility into interorganizational connections. This null direct effect also suggests that professional-level stability is not in itself a structural prerequisite for collaboration, even if, as shown below, it may condition the degree to which managerial political capital predicts tie formation. In parallel, public sector experience is not clearly associated with collaboration at the professional level, which does not grant enough discretion to deploy it for establishing ties.

Controls. Full control coefficients are reported in Table A11 in the SI. The endogenous structural terms behave as expected: the large positive GWDEG coefficient (≈3.9–4.1, $p < 0.001$) confirms pronounced preferential attachment, and the negative GWESP coefficient (≈ -0.13, $p < 0.001$) indicates that collaborations tend toward open, non-redundant configurations rather than closed triads. The homophily terms confirm that collaboration is strongly structured by regional and partisan proximity: agencies within the same *departamento* are roughly fourteen times more likely to form ties ($\exp(2.62) \approx 13.8$, $p < 0.001$), and co-partisan organizations are more than four times as likely to collaborate as cross-partisan ones ($\exp(1.53) \approx 4.6$, $p < 0.001$). Agencies are less likely to partner with organizations of the same government tier ($\exp(-0.34) \approx 0.71$, $p < 0.001$) or organizational type ($\exp(-1.78) \approx 0.17$), reflecting a preference for complementary over redundant mandates. Among exogenous controls, organizational size is the strongest predictor ($\exp(1.69) \approx 5.4$, $p < 0.001$), collaboration probability decreases monotonically with jurisdictional deprivation, and geographic distance is associated with roughly halved odds ($\exp(-0.73) \approx 0.48$, $p < 0.001$). Counter to expectations, climate-event exposure is negatively associated with collaboration ($\exp(-0.95) \approx 0.39$, $p < 0.001$), suggesting that high-risk agencies may face capacity constraints that reduce their attractiveness as partners.

5.3 Interaction Models

The main models establish that managerial and professional patronage are associated with collaborative tie formation in opposing directions. The interaction models address a further question: under what conditions is patronage associated with collaboration? The evidence points to a four-part story: professional patronage is consistently associated with reduced collaboration; managerial patronage is associated with a collaboration premium only when a minimum stable bureaucratic platform is present; for managerial patronage, the operative resource appears to be political capital rather than domain expertise; and that capital is most strongly associated with tie formation across same-party lines.

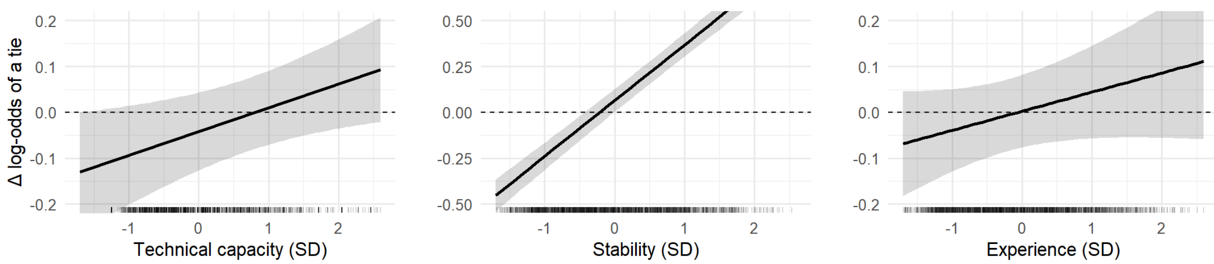

Figure 4: Marginal effects of aggregate patronage. Marginal effect of patronage on the log-odds of collaboration conditional on bureaucratic attributes (standardized). Lines show log-odds effects; shaded areas are 95% delta-method CIs; rugs show the distribution of the moderator. The Patronage × Technical capacity interaction term is significant at the 0.05 level; Patronage × Stability is significant at the 0.001 level; Patronage × Public experience is not significant at conventional levels. Table A12 in the SI presents the full regression results.

Figure 4 plots the marginal association of aggregate patronage with the log-odds of collaboration conditional on each bureaucratic attribute. The most striking interaction is with stability: the association is near-zero or negative at low levels of organizational stability but becomes positive and significant once agencies are at or above the mean on expected tenure—patronage is associated with a collaboration premium only where organizational churn does not dominate. The interaction with technical capacity runs in the same direction, though more modest and marginally significant, suggesting that a minimum technical base also conditions the conversion of political connections into actionable partnerships. The interaction with public sector experience is positive but imprecise.

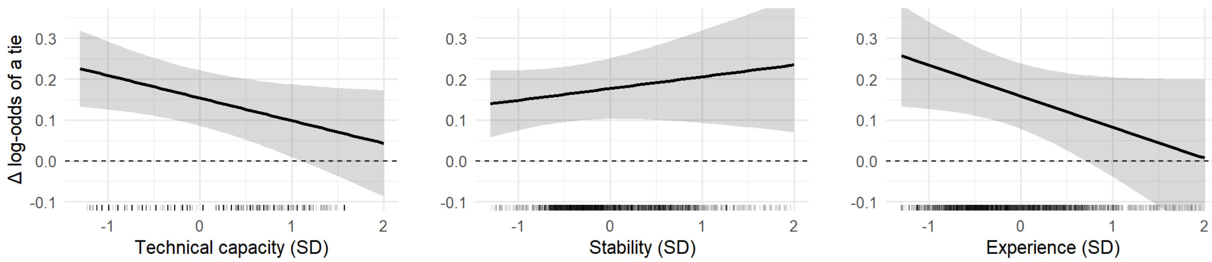

Figure 5: Marginal effects of managerial-level patronage. Marginal effect per 1-SD of managerial patronage on the log-odds of collaboration conditional on bureaucratic attributes (standardized). Lines show log-odds effects; shaded areas are 95% delta-method CIs; rugs show the distribution of the moderator. Only the Managerial patronage × Managerial technical capacity interaction term is significant at the 0.05 level. Table A13 in the SI presents the full regression results.

Figure 5 disaggregates these interactions to the managerial level. The dominant interaction here is with managerial technical capacity, but the sign is negative: the collaboration premium associated with politically appointed managers is largest where managerial specialized knowledge is thinner. This substitution pattern is consistent with the interpretation that where in-house technical expertise is already high, career staff drive collaborative engagement on that basis; where it is lower, political connections fill that structural gap. Managerial stability and public sector experience show weak or non-significant interactions, further consistent with the pathway running through relational capital rather than organizational continuity. At the professional level, no interactions reach conventional significance thresholds (Figure 6). The negative association between professional patronage and collaboration holds regardless of the bureaucratic resources present, consistent with operational-level appointments introducing a functional misalignment that better-endowed organizations do not neutralize.

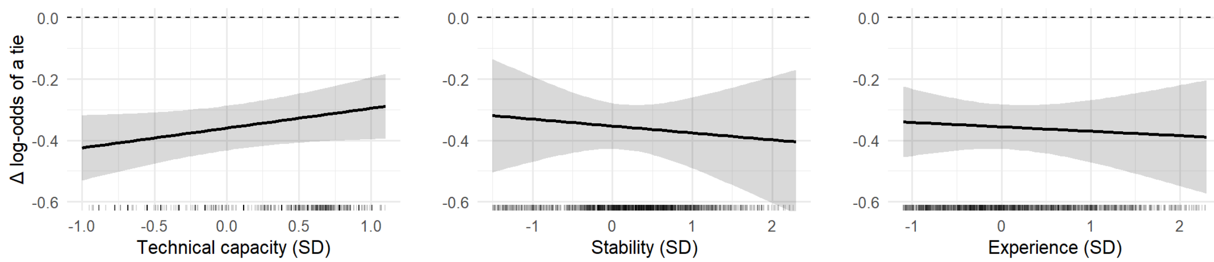

Figure 6: Marginal effects of professional-level patronage. Marginal effect per 1-SD of professional patronage on the log-odds of collaboration conditional on bureaucratic attributes (standardized). Lines show log-odds effects; shaded areas are 95% delta-method CIs; rugs show the distribution of the moderator. None of the interactions are statistically significant at the conventional levels. Table A14 in the SI presents the full regression results.

If political capital is the primary resource managerial appointees contribute, I suggested that the collaboration-enhancing effect of managerial patronage should be concentrated along partisan lines. Figure 7 provides direct evidence of this pattern. Only the managerial patronage $\times$ political homophily interaction is positive and statistically significant; neither the aggregate nor professional-level interactions approach significance. A one-standard-deviation increase in managerial patronage is associated with a 7% reduction in the odds of cross-partisan collaboration ($\exp(-0.07) \approx 0.93$) but a 51% increase in the odds of same-party ties ($\exp(-0.07+0.48) \approx 1.51$). Organizations with politically connected managers show selectively higher coordination among co-partisans and little or no corresponding pattern across partisan divides.

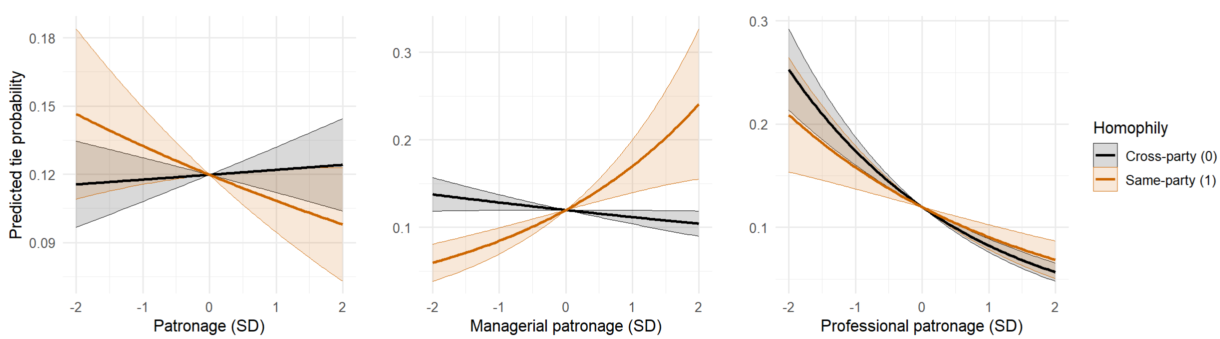

Figure 7: Marginal effects of patronage by political homophily. Marginal effect of patronage on predicted tie probability by political homophily. Curves anchored at the overall tie probability (network density = $p = 0.12\%$). Ribbons show 95% delta-method CIs for slope uncertainty. Only the interaction for Managerial patronage $\times$ political homophily is significant at the conventional levels. Table A15 in the SI presents the full regression results.

The aggregate-level interaction with stability (Figure 4) is stronger and emerges more cleanly than its managerial-level counterpart (Figure 5), hinting that the relevant stabilizing base lies not in managerial staff but in the professional corps beneath them. Interacting managerial patronage with professional-level stability directly, Figure 8 shows that a one-standard-deviation increase in the average expected tenure of professional-rank bureaucrats (approximately 5.4 years) is associated with roughly 11% stronger collaboration odds for a given level of managerial patronage ($\exp(0.102) \approx 1.11$). This cross-rank pattern aligns with the organizational balancing thesis: the collaboration premium of managerial political appointments is largest where the professional corps beneath them is stable and predictable (Krause et al. 2006).

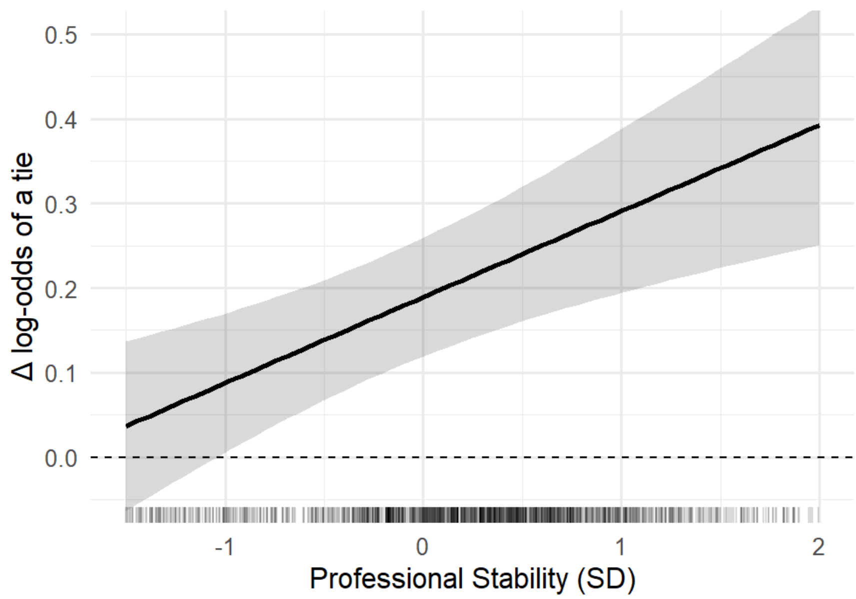

Figure 8: Marginal effects of managerial patronage conditional on professional-level stability. Line shows log-odds effects; shaded areas are 95% delta-method CIs; rugs show the distribution of the moderator. Table A16 in the SI presents the full regression results.

5.4 Alternative Explanations

The most obvious concern is that the managerial patronage effect reflects organizational capacity rather than the political capital of appointees. Three features of the analyses work against this interpretation. The models directly control for agency size, jurisdictional fiscal category, and the distribution of technical qualifications and public sector experience, so the patronage coefficients are estimated net of the capacity dimensions most plausibly driving both hiring practices and collaborative engagement. More fundamentally, a capacity confound would need to operate in opposite directions at different ranks to simultaneously explain why managerial patronage increases collaboration while professional patronage decreases it. Most decisively, a capacity argument predicts that well-connected managers should facilitate collaboration broadly, yet the collaboration premium is concentrated almost entirely among co-partisan dyads and is negligible or negative across partisan lines.

A more problematic alternative concerns reverse causality: politicians may strategically place well-connected managers in agencies already embedded in collaborative networks rather than those managers generating new ties. Two features partially constrain this interpretation. The ERGM degree distribution term absorbs variation attributable to pre-existing hub status, so the patronage coefficients are not simply capturing prior centrality. More tellingly, the partisan concentration of the effect is difficult to reconcile with strategic placement, which would predict a general rather than a homophilically structured collaboration premium. Fully ruling out reverse causality would nonetheless require a temporal analysis tracking changes in appointments alongside tie formation—an identification strategy that the cross-sectional structure of the personnel records places beyond the scope of the present analysis.

6 Concluding Remarks

The political dynamics of the agencies responsible for deciding and implementing collaborative governance arrangements have received remarkably little attention in the governance literature (Morrison et al. 2019), despite growing recognition that horizontal coordination has become indispensable for managing multilevel systems and cross-jurisdictional policy problems (Bodin 2017). In contexts where cross-jurisdictional collaboration is a functional necessity (e.g., environmental management, climate adaptation, and common-pool resource governance) the bureaucratic politics of collaborative capacity deserves direct analytical investigation rather than being taken as background institutional context.

I have argued that intergovernmental collaboration is shaped by the political mechanisms embedded in bureaucratic hierarchies, and that patronage is an important and underappreciated part of that story. The central claim is that how patronage shapes collaboration depends on where it penetrates the hierarchy and on the fit between the resources appointees carry and those the organization needs to engage as a collaborative partner. The empirical evidence from environmental collaboration networks in Colombia lends support to these theoretical arguments: managerial patronage can facilitate collaboration when incumbents place politically embedded actors in positions where their relational capital can be deployed for brokerage and negotiation, while patronage at operational ranks undermines coordination by displacing the technical expertise and organizational continuity on which collaborative capacity depends. The cross-rank complementarity finding extends this logic further, showing that the collaboration premium associated with managerial political capital is largest when a stable professional corps provides the execution capacity beneath it. The paper thus carries into the terrain of network and collaborative governance the classic bureaucratic politics thesis that government performance is improved by an organizational balance between political and autonomous staff (Krause et al. 2006; Lewis 2008; Peters and Bianchi 2020; Svara 2001).

Two clarifications and open research agendas follow from this framing. The first concerns the upstream political motivations behind appointments. The argument advanced here is agnostic about the incentives that lead politicians to make the appointments they do. That is, whether incumbents place well-connected managers for electoral reasons, party coordination purposes, or programmatic policy goals is orthogonal to the central claim, which concerns the bureaucratic mechanism through which those appointments shape collaborative behavior once in post. The open question is under what political conditions incumbents optimize for a collaboration-enhancing functional fit. Electoral competition and credible accountability institutions may pressure incumbents to invest in collaborative capacity by selecting managers whose political capital is deployable for governance rather than merely for rent extraction (Brierley 2020; Geddes 1996; Grindle 2012; Toral 2024a). The nature of the policy problem (e.g., whether cross-jurisdictional coordination is electorally visible and attributable) likely conditions how much incumbents care about the functional consequences of their appointment strategies, and party organizational strength may further determine whether the partisan channels that managerial appointees activate are dense enough to sustain collaboration.

The second clarification concerns collaboration quality rather than formation. Whether the bureaucratic politics mechanisms that enable collaborative ties lead to effective environmental governance, or instead reproduce political collusion without substantive policy benefits, is not self-evident from the results. Patronage-enabled networks risk being politically coherent but technically shallow if homophilic channels are used to prioritize political visibility over substantive policy adaptation and learning. This points to a broader question the collaborative governance literature has largely sidestepped: whether facilitating collective action, regardless of the mechanism through which it emerges, is always conducive to virtuous governance arrangements. By treating tie formation as intrinsically desirable and setting aside the political incentives behind joint action, the literature has underestimated the potential pathologies that collaborative mechanisms can generate in highly politicized contexts, where the boundary between coordination and collusion is difficult to observe and easier to cross.

A third direction concerns identification. The present analyses establish systematic associations between patronage profiles and collaborative behavior, but the observational design means that causal claims remain tentative: patronage levels are not randomly assigned, and no institutional feature of the Colombian system provides a clean discontinuity for exogenous variation. Three strategies would allow future work to move closer to identification. Electoral transitions, which produce sharp, politically driven reshuffles of bureaucratic personnel, offer variation in patronage levels that is plausibly exogenous to pre-existing collaboration networks. Linking annual snapshots of bureaucratic composition to network evolution over time would further allow the temporal ordering of appointments and tie formation to be directly observed rather than inferred from cross-sectional co-variation. Finally, individual-level data on appointee contacts, meeting records, and interorganizational networks would enable the relational capital mechanism to be tested directly, rather than through the organizational-level interaction patterns used here.

All in all, this paper speaks to a fundamental feature of governance in fragmented, polycentric settings: coordination is achieved through politics as much as through institutional design, and the political logics that define bureaucratic structure can, under the right conditions, serve as its infrastructure. The challenge for scholars and practitioners is not to wish away the political dimensions of public organizations but to understand when and how political mechanisms interact with the dynamics of collective action. In decentralized settings where administrative fragmentation makes horizontal coordination a structural necessity and mixed bureaucratic systems are the modal institutional form, that understanding is crucial to assess the pitfalls of current state reforms and envision how to design new ones.

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