Coordination vs. Competition: Local Government Centralization and Metropolitan Growth

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Abstract

This paper examines how local government organization affects metropolitan growth. My focus is on local government centralization, or the degree to which public decision-making authority is dispersed across governments. I provide a novel measure of centralization using the distribution of local government employment across jurisdictions within U.S. metropolitan statistical areas (MSAs). Because local government constitutes only part of the broader institutional framework operating in metropolitan areas, the relationship between local government organization and growth may be endogenous. To address this, I use an instrumental variables strategy based on post office closures following the introduction of Rural Free Delivery in 1902. In both the ordinary least squares and IV specifications, I find that more centralized MSAs experience greater population and private employment growth. Examining potential mechanisms that link centralization to growth, I find that centralized MSAs employ more local government workers, collect more revenue, and are less likely to host data centers, a highly tax-sensitive form of mobile capital. The findings suggest that, at the metropolitan scale, internalizing cross-jurisdictional externalities may be more important for growth than increased competition between governments.

1 Introduction

An ongoing debate in public economics centers on how the organization of government shapes public-good provision. One way organization may have an effect is by altering the incentives and constraints governments face when choosing policy. A key dimension of the debate, emphasized by Oates (1972), is whether public decision-making authority should be *centralized* into a few large jurisdictions or *decentralized* across many smaller ones. An important tension concerns the role of competition between governments, giving rise to two distinct conceptual frameworks. The *competition* framework emphasizes the benefits of decentralization by focusing on the information and incentives of public officials, as in Tiebout (1956), Besley and Case (1995), and Hayek (1945). In contrast, the *coordination* framework emphasizes how externalities can turn competition between governments into a "race to the bottom." Central to the coordination framework are the tax-competition models developed in Wilson (1986), Zodrow and Mieszkowski (1986), and Wildasin (1988).[1]

While the debate occurs across all levels of government, it is particularly salient at the local level.[2] In this paper, I estimate the effect of centralization on metropolitan growth using cross-sectional variation across metropolitan statistical areas (MSAs) in the United States. Using the Annual Survey of Public Employment and Payroll, I provide a new measure of local government centralization using the distribution of local government employment across individual governments within the same MSA. I find that centralization increases both population and employment growth across multiple time horizons using both ordinary least squares and an instrumental variables strategy based on the Postal Service's introduction of Rural Free Delivery in 1902. I then turn to the mechanisms behind centralization's positive effect on growth by examining local government employment, tax revenue, and data center siting across MSAs, as data centers provide a close empirical analog to the mobile capital found in tax-competition models.

The questions surrounding local government centralization are relevant to policy. Many countries have actively restructured local governments in recent decades, often through mergers that can be viewed as discrete increases in centralization. Between 2008 and 2017, fifteen European countries implemented large-scale municipal reforms (Swianiewicz 2018). Japan undertook a similar consolidation in the early 2000s, reducing its number of municipalities from 3,232 to 1,719 (Miyazaki 2014). Toronto merged its six core municipalities into a single government unit in 1999 (Slack and Bird 2013). In the United States, voters have repeatedly

[1] See Wilson (1999) for a review of both frameworks.

[2] Two other notable arenas concern decentralizing authority from national governments to states, and centralizing national authority to a supranational institution.

considered local government consolidation, including 166 city–county merger referenda since 1815 (Hall, Matti, and Zhou 2020), alongside numerous smaller merger and annexation decisions. Because these reforms represent discrete increases in centralization, evaluating their effects or conducting ex ante counterfactual analysis depends critically on the behavioral framework used to characterize the strategic environment faced by local governments.

Growth is particularly useful for distinguishing between the two competing frameworks because they generate opposing predictions. The competition framework emphasizes the benefits of intergovernmental competition and predicts that more decentralized MSAs will grow faster. The coordination framework instead emphasizes the costs of externalities across governments and predicts that more centralized MSAs will grow faster. Although the relative importance of these opposing forces is unobservable, growth captures their net effect on the attractiveness and economic performance of a location. The estimated effect of centralization on growth therefore distinguishes between the reduced-form predictions of the two frameworks, while the mechanism outcomes examined later help explain why.

I confront two main empirical challenges in estimating the effect of centralization. The first is measurement, because centralization is latent. I develop a new measure of centralization using the Annual Survey of Public Employment and Payroll, which allows me to focus on how local government employment is distributed across government units within an MSA.[3] Even within this narrower focus, there are many possible ways to measure centralization using local government employment. This is closely related to the broader problem of measuring industrial concentration. I build on the insight and methodology of Ahern, Kong, and Yan (2024) by constructing a family of centralization statistics using local government employment shares within each MSA, similar to how they create a family of concentration statistics from industry market shares.[4] I then aggregate these statistics into an index that captures a common component that I call centralization.[5]

The second empirical challenge concerns endogeneity. Endogeneity can be difficult to define in this setting as many variables that initially appear to be potential sources may instead be channels through which centralization affects growth. For example, higher rates of crime likely reduce growth prospects, but centralization may directly affect the efficiency of the lo-

[3]My strategy is similar to, but distinct from, that of Goodman (2013). He focuses on how local government employment is distributed across types of local government and measures employment shares for municipalities, counties, school districts, and special districts.

[4]For additional discussion of concentration and diversity measures in economics and other fields, see Patil and Taillie (1982), Hill (1973), Hannah and Kay (1977), Maturo (2018), and Nocke and Schutz (2025).

[5]I show in the appendix that this aggregation has no effect on the qualitative conclusions of this paper. I reproduce all results using the centralization statistic that minimizes the absolute t-statistic of the estimated effect. The qualitative results do not change, and in many cases the point estimate is larger.

cal police force. Including crime rates in specifications then risks controlling for a potentially important channel through which centralization affects growth. Similar arguments apply to human capital if centralization affects the efficiency of schooling (Hoxby 2000); to infrastructure if centralization affects public investment; and to housing prices if centralization affects land-use regulation and zoning. These variables are therefore plausibly post-treatment: controlling for them would risk conditioning on outcomes of centralization rather than isolating pre-existing sources of omitted-variable bias (Angrist and Pischke 2009). While this view narrows the scope of the endogeneity problem, it does not eliminate it.

One remaining source of endogeneity stems from the information lost when abstracting to the index. My measurement of centralization captures the concentration of local government employment across legally distinct governments but abstracts from government type, vertical versus horizontal fragmentation (Goodman 2019), functional categories within employment, political organization, and internal structure. MSAs with the same measured centralization may distribute employees differently across counties, municipalities, school districts, and special districts; concentrate particular functions in specialized governments; adopt different formal governance rules;[6] and organize individual governments into many, potentially competing, departments. If these unmeasured organizational features affect growth and correlate with measured centralization, they create omitted-variable bias.

A second source of endogeneity stems from the fact that local governments constitute only one part of the broader institutional environment active in a metropolitan area. Civil-society and business organizations, non-profits, social rules and norms, and programs administered by state and federal governments may independently affect growth. Because these complementary institutions may develop jointly with, or in response to, local government organization, their influence is difficult to separate empirically.[7] Standard estimation approaches may therefore conflate the effect of centralization with the effects of omitted institutions, producing biased estimates.

To address these concerns with an IV strategy, I require an instrument that generates variation in centralization while remaining orthogonal to the omitted determinants of the outcomes I study. Burns (1994) argues that the modern organization of local government reflects a historical process of annexation and defensive incorporation. Existing cities seeking to expand had to annex surrounding territory. Where that territory was already settled, residents

[6] An important example would be legislative restrictions on local governments' ability to adjust tax rates without direct electoral approval, commonly associated with the "tax revolts" of the 1970s. See Mullins and Wallin (2004) for additional details.

[7] The joint determination of these two classes of institutions is likely nontrivial. Centralization may increase demand for highly localized institutions, while larger centralized governments may crowd them out.

could block annexation by defensively incorporating a new local government. Repeated annexation attempts therefore generated additional local governments and, over time, a more decentralized metropolitan structure.

The instrument exploits variation in the initial conditions of this process. Under this account, metropolitan areas containing more historical settlement nodes should have experienced more potential annexation conflicts and therefore more defensive incorporation. Because the number of historical settlement nodes is not directly observed, I proxy for it using the number of post offices that closed between 1902 and 1910 following the rollout of Rural Free Delivery.

Rural Free Delivery (RFD) was implemented nationally in 1902. Before its rollout, residents of rural areas typically collected mail from local post offices whose viability depended on the postal activity generated by surrounding communities. After RFD made free mail delivery available to rural residents, many of these post offices closed. Because RFD extended the 1863 Free City Delivery program to rural areas (United States Postal Service 2024), the resulting closures were disproportionately concentrated outside existing cities. The number of post office closures following RFD therefore provides a proxy for the number of historical settlement nodes, which in turn shifts the initial conditions governing the subsequent process of annexation and defensive incorporation. The validity of the instrument therefore requires that the historical rural settlement variation captured by post office closures affects modern metropolitan outcomes only through its effect on the organization of local government.

Across both OLS and the IV strategy, I find that greater centralization increases growth. I measure growth in two ways. First, I use detailed population data from the National Cancer Institute's Surveillance, Epidemiology, and End Results (SEER) Program to measure MSA population every decade from 1977 to 2017, choosing years to coincide with the ASPEP data. Second, I measure private employment over corresponding years using the Business Dynamics Statistics. Across both measures, I find that greater centralization increases growth.

While the positive effect of centralization on growth lends empirical support to the coordination framework, further work is required to determine the mechanisms behind the effect. The coordination framework emphasizes the role of externalities, but also leaves open the possibility that growth could be driven by economies of size (Fox and Gurley 2006). The remaining empirical exercises examine a series of outcomes that help distinguish among the channels linking local government centralization to growth.

I first ask if centralization promotes growth through a narrowly defined economies of size mechanism. If economies of size enabled MSAs to become more technically efficient, then the per-unit cost of public-good provision would decrease. More centralized MSAs would then

be able to provide the same public good bundle with fewer resources or expand provision at the same total cost. Either case can promote growth. However, I find that more centralized MSAs employ more local government workers and collect more tax revenue. These findings contradict the narrow version of the economies of size mechanism in which greater centralization promotes growth by passing cost savings on to constituents.

In contrast, the higher local government employment and higher tax revenue collection in more centralized MSAs align with the predictions of tax competition models. In decentralized MSAs, taxing authority is fragmented across jurisdictions, intensifying competition for a common mobile tax base. In these models, individual local governments fail to internalize two externalities. First, the tax base attracted by lowering their own tax rates comes partially from neighboring local governments. Second, the tax base lost by increasing their tax rates is partially directed toward their neighbors. If decentralized MSAs do not internalize these externalities to the same extent that centralized MSAs are able to, aggregate tax revenue collection can fall (Brueckner and Saavedra 2001). With less revenue, local government employment can fall as well. By internalizing these externalities, more centralized MSAs can sustain higher revenue collection and hire a greater number of workers.

While tax competition matches the observed empirical patterns, it is not decisive. An extended economies of size mechanism could still fit the evidence under certain conditions. If economies of size lower the unit cost of public goods, the lower costs may induce a rebound effect (Saunders 1992; Gillingham, Rapson, and Wagner 2016), leading to increased input use despite greater technical efficiency. This rebound can explain the greater local government employment observed in more centralized MSAs. The increase in tax revenue imposes a stronger requirement: public-good provision must expand enough that total expenditure rises even as unit costs fall. One plausible way for this to occur is through the extensive margin. Projects or services that are too costly for decentralized governments may be feasible if centralized MSAs are more technically efficient.

To partially distinguish between tax competition and this extended economies of size mechanism, I examine location decisions made by highly mobile, tax-sensitive firms. Tax competition predicts that these firms should be more likely to locate in decentralized MSAs, where a larger number of local governments compete to attract a common mobile tax base. The extended economies of size mechanism does not generate this prediction: if centralization promotes growth by lowering the cost of public-good provision, there is no reason for the effect of centralization to reverse for these firms in particular. As a concrete case, I examine where firms choose to construct data centers and find that more centralized MSAs are less likely to host them. Given the positive effect of centralization on aggregate employment

growth, the negative effect on data centers cuts against economies of size as the dominant mechanism but follows naturally from tax competition. Relative to most other firms, data centers are highly mobile and tax-sensitive; the benefits they generate accrue primarily to the host jurisdiction; and their costs are dispersed throughout the MSA through higher electricity prices, at least in the short run (Alvarez et al. 2026). These features make data center siting decisions especially sensitive to the coordination failures among local governments that underlie tax competition.[8]

As a useful analogy, a firm deciding where to construct a data center resembles a *seller* choosing among potential auctions. More decentralized MSAs, with a larger number of local governments, resemble auctions with more competing bidders. Just as additional bidders increase seller surplus in standard auction settings, greater competition among local governments can lower the effective tax or regulatory cost faced by the firm. Data centers may therefore locate disproportionately in decentralized MSAs even if centralized MSAs experience greater aggregate employment growth. The result suggests that tax competition is sufficiently salient to affect firm location decisions and may also help explain Slattery (2025)'s finding that firms capture most of the surplus generated when governments compete to attract them into their own jurisdiction.

Taken together, the positive relationship between centralization and growth suggests that the coordination framework better describes local government behavior than the competition framework in metropolitan areas. The mechanism evidence likewise points more strongly toward centralization increasing growth by alleviating coordination failures than by exploiting economies of size, although the evidence cannot fully rule out a role for the latter mechanism. Two features of local government may help explain why coordination failures are particularly important in this setting. First, interactions between local governments occur in close spatial proximity, potentially amplifying spillovers and externalities. Second, local governments may select into providing goods and services over which they have a comparative advantage relative to private provision, and that comparative advantage is likely greatest for activities characterized by externalities (Sinn 1997). Together, these features make externalities especially salient and create scope for competition across decentralized jurisdictions to inefficiently suppress public-good provision. Centralization can therefore promote growth by internalizing these externalities and expanding the provision of growth-enhancing public goods.

[8]The incentives created by MSA-wide negative externalities may explain the heated political battles over data centers. Even if all jurisdictions in an MSA would prefer not to have a data center, a worse outcome for a jurisdiction is to experience the negative effects, such as higher electricity bills, without the tax compensation that comes from the data center locating within their own jurisdiction. See Ingberman (1995) and Oates and Schwab (1988) for more details.

This work contributes to several literatures, beginning with research on political mergers. Because political mergers generally increase centralization, estimating the effect of centralization is central to evaluating these reforms. Existing studies, largely based on case studies and quasi-experimental designs, examine a wide range of institutional contexts, including Israel (Reingewertz 2012), China (Tang and Hewings 2017), Denmark (Blom-Hansen et al. 2016), Japan (Miyazaki 2018), and Iceland (Karlsson and Eythórsson 2022). Recent syntheses emphasize the highly heterogeneous and often contradictory results across settings and outcomes (Zheng, Sun, and De Witte 2023). In contrast, I estimate the effect of centralization across relatively homogeneous U.S. metropolitan areas using an instrumental-variables methodology. By abstracting from discrete merger events, this approach mitigates concerns about institutional heterogeneity and endogenous selection[9] that complicate merger-based analyses.

This paper also relates to the broader literature on government centralization and institutional performance. A large literature studies how the allocation of fiscal and political authority between national and subnational governments affects economic growth and institutional outcomes (Weingast 1995; Davoodi and Zou 1998; Treisman 2000; Rodden 2004; Iimi 2005; Baskaran and Feld 2013). This literature yields mixed predictions and empirical results: some work emphasizes the growth benefits of decentralization through stronger intergovernmental competition (Weingast 1995; Iimi 2005), while other work finds zero or negative effects of fiscal decentralization (Davoodi and Zou 1998; Baskaran and Feld 2013). One interpretation is that the relevant tradeoffs differ by scale. At the national level, decentralization may intensify competition across large and geographically separated jurisdictions, where spillovers are less salient. At the metropolitan level, however, the dominant concern may instead be the internalization of externalities among closely linked governments, making greater centralization more effective.

This paper fits most closely within a strand of the literature examining within-country variation in local government structure and centralization. Examples include Hoxby (2000) on educational performance, Stansel (2005) on population growth, Hatfield and Kosec (2013) and Hatfield and Kosec (2019) on wage growth and pollution, Hatfield, Kosec, and Rodgers (2024) on housing prices, Hobbs, Millsap, and Stansel (2019) on economic development, and Bianchi, Giorcelli, and Martino (2023) on labor market outcomes. However, existing studies rely on a wide range of centralization measures and sources of identifying variation that do

[9]Endogenous selection refers to the possibility that governments choosing to merge have the strongest unobserved reasons to do so. Existing studies use the idea that many merger waves are initiated by a federal government to plausibly overcome this difficulty. However, the secular trends that motivate the federal government to pursue merger policy may also affect policy evaluations and limit the external validity of their findings, especially if performance statistics are measured only over a short time horizon.

not map cleanly to local government centralization as measured here. Differences in how centralization is defined and identified may therefore help explain the divergence between prior findings and the results of this paper.

Favilukis and Song (2026) is likely the work most closely related to this paper and examines the effect of centralization on the intensity of zoning restrictions. They measure centralization using a Herfindahl–Hirschman index based on the residential land area of zoning jurisdictions and measure zoning restrictions using minimum lot sizes from Song (2025). To address endogeneity, they instrument centralization using the same land-based HHI, but measured using the jurisdictional boundaries of 1900. They find that greater centralization reduces zoning intensity[10] and develop a model in which coordination failures lead decentralized MSAs to impose excessively restrictive zoning. This mechanism closely parallels the one studied here: metropolitan decentralization can generate externalities that individual governments fail to internalize. Zoning may therefore be an important channel through which centralization affects metropolitan growth.

The remainder of the paper proceeds as follows. Section 2 presents the background and data on local government organization, including the measurement of centralization and the outcomes I study. Section 3 outlines the empirical strategy and describes the RFD instrument used for identification. Section 4 reports the results and discusses their implications for understanding the mechanisms of centralization. Section 5 concludes.

2 Background and Data

The main unit of analysis in this study is the metropolitan statistical area (MSA), as defined by the Office of Management and Budget.[11] An MSA consists of a group of counties organized around an urban core with substantial economic integration, as measured by commuting flows. MSAs are organized around principal cities, which are typically larger central cities. This definition provides a useful approximation to the relevant market for local government interactions. Concretely, I treat the MSA as the primary strategic environment in which local governments respond to one another's policies.

[10] Equivalently and in line with their framing, greater decentralization increases zoning intensity.

[11] MSAs constitute a middle ground between three commonly used geographic definitions. Core Based Statistical Areas (CBSA) include both Metropolitan Statistical Areas and Micropolitan Statistical Areas. Metropolitan Statistical Areas contain at least one urban area with a population of 50,000 or more, while Micropolitan Statistical Areas contain an urban area with a population between 10,000 and 49,999. Combined Statistical Areas (CSA) aggregate two or more adjacent CBSAs with substantial economic integration.

I focus on metropolitan areas rather than central cities alone because of their growing importance, especially in contemporary urban economies. As of 2025, roughly 85% of the United States population lives in a metropolitan area. Within metropolitan areas, 61% of people live in the suburbs outside the principal cities that anchor them. Post–World War II employment dispersion produced large suburban employment centers, often described as "edge cities," that contain substantial office, retail, and fiscal activity outside the historic central business district (Garreau 1991). As a result, central-city boundaries increasingly fail to capture the functional scale at which local governments interact. Glaeser and Kahn (2001) show that employment in U.S. metropolitan areas is highly dispersed, with only a small share of metropolitan employment located near the traditional city center. Similarly, the urban spatial-structure literature treats polycentricity as a central feature of modern urban economies rather than a deviation from the monocentric model (Anas, Arnott, and Small 1998; Giuliano and Small 1991).

The median MSA contains two counties, while 159 of the 378 MSAs consist of a single county. The Atlanta–Sandy Springs–Roswell MSA contains 29 counties, more than any other MSA. Because commuting patterns and urbanization evolve over time, MSA boundaries change across decades. To maintain a consistent geographic unit, I follow the literature and fix MSA boundaries at their 2017 definitions and use these boundaries to harmonize all historical variables. This allows changes over time to be interpreted within a consistent geographic footprint.[12]

I measure local government organization using the 1987, 1997, 2007, and 2017 waves of the Annual Survey of Public Employment and Payroll (ASPEP). The ASPEP is a full census of local governments conducted every five years and reports employment for each individual government unit. For each unit, employment is classified across 32 functional areas, such as policing, schools, and public works. In this study, I do not use this additional variation and describe each government unit only by its total employment.[13] The only alteration I make to the data is to drop employment in hospitals. By assigning each government unit to its respective MSA, I am able to describe each MSA as a vector of employment levels, which I use to construct the local government centralization indices discussed in the next section.

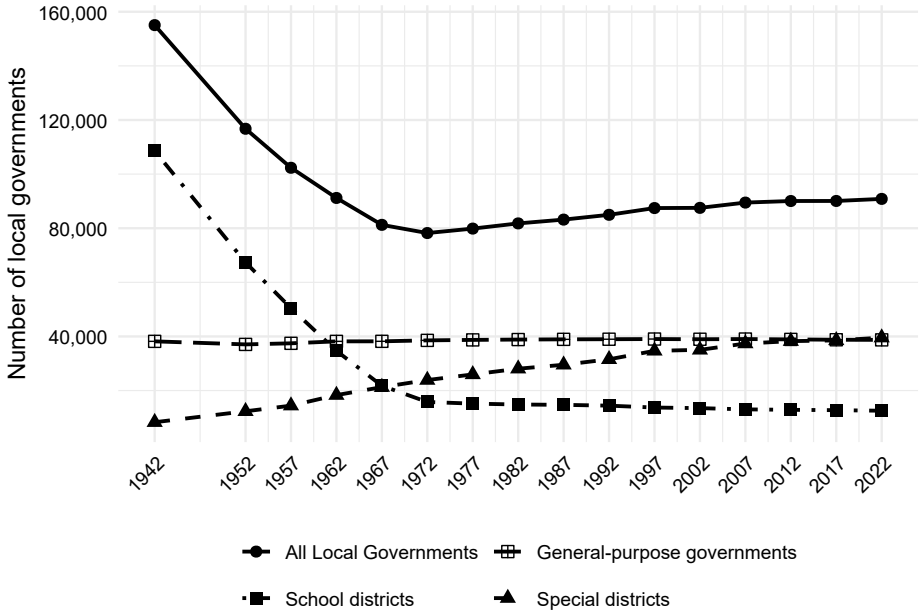
[12]This harmonization choice works against finding a positive effect on growth. The MSAs most affected by fixed 2017 boundaries are those that grew the most. Including governments in the added counties tends to decrease measured centralization, thereby attributing additional growth to more decentralized MSAs.

[13]Under this view, MSA m can be fully described by a $N_m \times 32$ matrix in which the columns capture employment across functional categories and the rows describe each government unit. I collapse this matrix by summing across the columns. Variation in how employment is distributed across functional categories, corresponding to summing over rows, may be an interesting source of variation, but here it can be treated as a source of endogeneity.

My empirical strategy treats local government organization as a persistent institutional feature rather than one that adjusts rapidly over time.[14] One possible source of persistence is the relative stability of transportation technology since the completion of the interstate highway system. Given commuting costs, the geographic scale of an MSA is likely bounded by the maximum distance residents are willing to commute.[15]

Figure 1 provides aggregate evidence that the institutional structure of U.S. local government stabilized around 1970. The y-axis reports the number of local governments, while the x-axis reports the year. The solid line with circles shows the total number of local governments; the dot-dashed line with squares shows school districts; the long-dashed line with squares shows general-purpose governments, including municipalities, townships, and counties; and the dashed line with triangles shows special districts.

Figure 1: Number of Local Governments Over Time



Notes: The figure plots the number of local governments in the United States over time. The solid line with circles shows the total number of local governments; the dot-dashed line with squares shows school districts; the long-dashed line with squares shows general-purpose governments, including municipalities, townships, and counties; and the dashed line with triangles shows special districts.

From the beginning of the series in 1942 until approximately 1972, the total number of local

[14] For empirical evidence on the inflexibility of political boundaries, see Epple and Romer (1989).

[15] Work-from-home arrangements may permit greater spatial dispersion within urban areas, as argued by Ramani and Bloom (2021). More recent work instead emphasizes that the main effect of work-from-home is migration toward lower-cost MSAs (Bick et al. 2024).

governments declined sharply as school districts consolidated.[16] After 1972, however, the total remained relatively stable. The number of general-purpose governments changed little, while the decline in school districts slowed substantially. The modest increase in the total number of governments after 1972 was driven primarily by the creation of special districts.

However, the increase in special districts overstates the corresponding change in the employment-based organization of local government. In 2022, special districts accounted for only 6.1% of local government employment, and 45% employed no workers, operating mostly as single-purpose taxing districts.[17] Figure 1 therefore suggests that the broad institutional structure of local government has remained stable since at least the early 1970s, while Table 1 below provides more direct evidence that the relative centralization of individual MSAs is highly persistent. Together, these patterns motivate the maintained assumption that the organization measured from 1987 to 2017 reflects a durable institutional structure extending throughout the period studied.

2.1 Local Government Centralization Index

Centralization refers to the degree to which decision-making authority is concentrated within a system of governments. A centralized system entrusts authority to one or a few large governments, while a decentralized system disperses authority across many smaller governments. In this section, I discuss how I measure variation in centralization across MSAs. To do so, I construct five local government centralization indices: one for each ASPEP survey year and one using the joint variation across all survey years. The indices are constructed by first obtaining a family of centralization statistics from the ASPEP employment data, then residualizing the statistics on population and land area, and finally aggregating the residualized measures into an index.

As discussed in the introduction, centralization is latent and can be measured in many different ways. Ahern, Kong, and Yan (2024) discuss a similar problem and show how to map industry market shares into a family of concentration statistics. I use this insight to build my measure of centralization by working with a vector of local government employment shares within an MSA. Let MSA m in survey year t be described by the vector s_{mt}, whose entries correspond to the share of employment accounted for by each government unit in the MSA. Define a family of centralization statistics as

[16] Fischel (2010) argues that these closures had limited effects because most involved one-room school districts.

[17] For additional discussion of taxing districts and a model of taxation with overlapping jurisdictions, see Ruggieri (2024).

$$ {}^{q}\lambda(s_{mt}) = \left(\sum_{j=1}^{|s_{mt}|} s_{jmt}^{q} \right)^{\frac{1}{1-q}}, \quad q \in [0, \infty). $$

Using this notation, ${}^{0}\lambda$ measures the number of governments, ${}^{2}\lambda$ is the inverse HHI, and $\lim_{q\to\infty} {}^{q}\lambda$ approaches the inverse one-government concentration ratio.[18] A convenient interpretation of these statistics is as "effective numbers" or "number-equivalent measures" (Adelman 1969; Schmalensee 1977; Kelly 1981), which measure how many symmetric governments it would take to match the centralization implied by the underlying vector. By varying the free parameter q, I can obtain a family of centralization statistics for each MSA. For low q, the statistics weight all local governments more uniformly, emphasizing richness. For high q, they concentrate weight on the largest local governments, emphasizing unevenness. In this setting, there is no theoretical reason for choosing one statistic over another. For that reason, I evaluate 50 statistics at evenly spaced points from the family of centralization statistics for each MSA and survey year.[19]

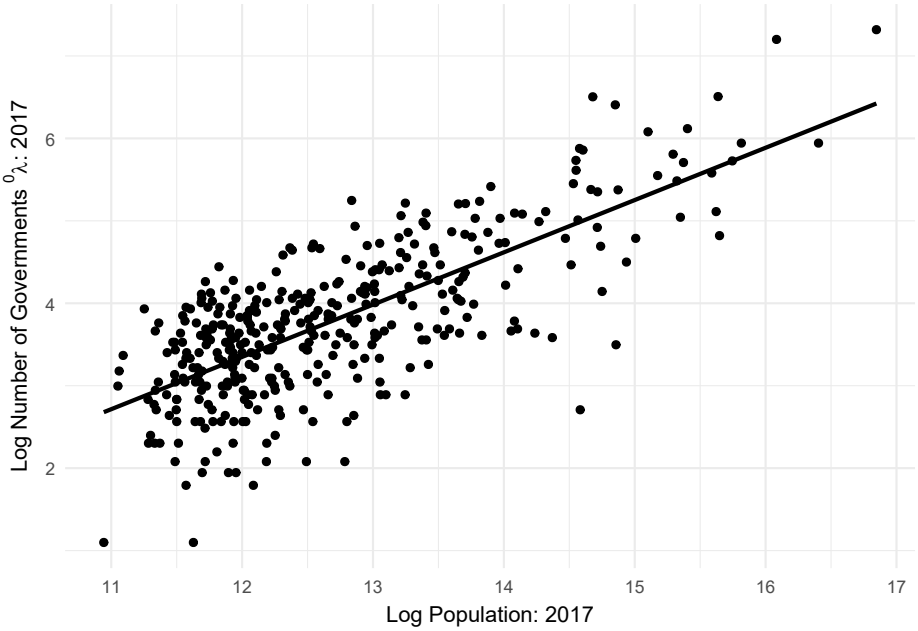
Directly comparing the sequence of effective numbers for each MSA can be misleading because it does not account for differences in MSA population or land area. Due to this, the same centralization statistics can represent different levels of centralization depending on population.[20] Similarly, land area may capture geographic differences across MSAs that affect the interpretation of centralization. Figure 2 highlights this dependence. The horizontal axis plots log MSA population in 2017 and the vertical axis plots the log number of governments in 2017.

[18] It is easy to show that $\lim_{q\to 1} {}^{q}\lambda$ is the Shannon entropy $\exp(-\sum_j s_j \log(s_j))$ and that for an MSA with equal employment across its $|s|$ government units, ${}^{q}\lambda$ is constant across q and equal to $|s|$.

[19] I use $u = \frac{q}{1+q}$ to map the domain into the unit interval and then choose points evenly in the interval.

[20] Fifty local governments in a lower-population MSA likely implies decentralization, while fifty governments in a high-population MSA likely indicates centralization.

Figure 2: Number of Governments and Population



Notes: The figure plots the log number of local governments against log MSA population in 2017 for the 378 MSAs in the sample. Each point is an MSA, and the solid line is the OLS fit.

To account for this dependence, I residualize all statistics on the MSA's log population and log land area, producing a vector of residualized statistics. For each of the 50 centralization statistics indexed by i, I estimate

$$\log^{q(i)} \lambda(s_{mt}) = \alpha_i + \beta_{iP} \log P_{mt} + \beta_{iL} \log L_m + \epsilon_{im},$$

where P_{mt} and L_m are the population and land area in MSA m during survey year t. Rearranging and centering yields

$${}^{q(i)}\tilde{\lambda}_{mt} = \log \frac{{}^{q(i)}\lambda(s_{mt})}{P_{mt}^{\beta_{iP}} L_m^{\beta_{iL}}}$$

As mentioned above, the numerator of this expression can be interpreted as the effective number of governments under centralization statistic i, while the denominator estimates the expected value of the statistic conditional on population and land area. The log-ratio therefore captures deviations from the level of local government centralization predicted by metropolitan scale and is orthogonal to those measures by construction.

The final step of the process is to aggregate the residualized statistics into a single index. An arbitrary index can be written as

$$C_{mt} = \sum_{i=1}^{50} w_i \cdot {}^{q(i)}\tilde{\lambda}_{mt}.$$

To construct the index I take to the data, I must select weights. I choose the weights using principal component analysis because the goal is to measure the common centralization signal contained within the family, not to privilege any specific statistic. PCA provides a disciplined way to combine these measures by extracting the dominant dimension of variation they share across MSAs. The index I use in the analysis is the standardized first principal component. Within each survey year, the index explains over 80% of the joint variation, and the pooled index, which combines each MSA's centralization statistics across survey years, explains 77% of the variation, indicating a substantial common component.

An alternative empirical strategy would be to examine variation in centralization within each MSA over time. However, I find that the relative ordering of MSAs by centralization is highly persistent. In Table 1, the standardized rank of the 2017 centralization index is regressed on the standardized ranks of the 2007, 1997, and 1987 indices in columns (1)–(3), respectively. Because both variables are standardized ranks, the estimated coefficient is equivalent to the Spearman rank correlation. Perfect persistence in the ordering of MSAs would correspond to a coefficient of one.

Table 1: Rank Persistence in Local Government Centralization

	Dependent variable:		
	Centralization Index: 2017		
	(1)	(2)	(3)
Centralization Index: 2007	0.987***		
	(0.008)		
Centralization Index: 1997		0.966***	
		(0.013)	
Centralization Index: 1987			0.943***
			(0.017)
Observations	378	378	378
R^2	0.974	0.933	0.889

*p<0.1; **p<0.05; ***p<0.01

Notes: Each column reports an OLS regression of the standardized rank of the 2017 centralization index on the standardized rank of the centralization index from the year listed in the row label. The estimated coefficient is therefore equivalent to the Spearman rank correlation. The sample includes 378 MSAs. Standard errors are reported in parentheses.

The results indicate a high degree of persistence. MSAs that rank as more decentralized in earlier periods generally remain more decentralized in later periods. This stability in

relative centralization motivates the cross-sectional approach used to estimate the effect of centralization and further supports treating metropolitan institutional structure as highly persistent over the study period.

2.2 Outcomes

2.2.1 MSA Population and Employment Growth

I construct MSA-level population variables using county-level data from the National Cancer Institute's Surveillance, Epidemiology, and End Results (SEER) Program, aggregated to 2017 MSA definitions. Population is measured for 1977, 1987, 1997, 2007, and 2017. In the shortest horizon (2007–2017), average MSA population growth is 7.3%, while over the longest horizon (1977–2017) it is 55.2%. To account for these compounded differences, I standardize each growth outcome to have mean zero and standard deviation one. This standardization allows the effect of centralization on growth to be compared across time periods.

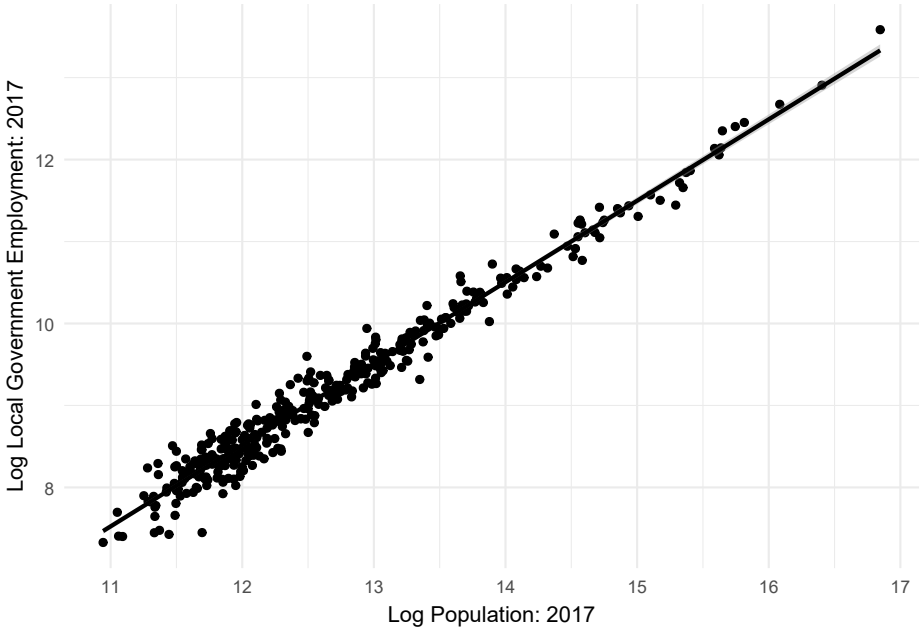
I construct employment measures using the Business Dynamics Statistics dataset. Because these data begin in 1978, I use employment levels for 1978, 1987, 1997, 2007, and 2017. Employment growth is measured analogously as the log difference between employment in 2017 and employment in a base year.

2.2.2 Local Government Employment and Finances

Local government employment is measured using the same ASPEP data used to construct the centralization index. Total employment is defined as the sum of employment across all government units within an MSA.

Figure 3 plots log local government employment against log population for 2017.

Figure 3: Local Government Employment and Population



Notes: The figure plots log local government employment against log MSA population in 2017 for the 378 MSAs in the sample. Each point is an MSA; the solid line is the OLS fit; and the shaded region is the 95% confidence interval.

The strong log-linear relationship motivates the log-log empirical specification used in the analysis.

To study government finances, I use data from the Census of Governments, which complements the ASPEP data by providing financial information for each government unit, allowing me to measure both total revenue and tax revenue. When constructing these data, I do not drop units with zero employment. This ensures that revenue from jurisdictions such as single-purpose taxing districts is included.

A major difference between the local government employment data and the financial data is the unit of measurement. Local government employment is measured in workers, whereas finances are measured in dollars. This creates a measurement issue because the value of a dollar varies significantly across geographic areas (Moretti 2013). To account for differences in local price levels, I use the ratio of BEA income to BDS employment as a proxy for average wages.[21]

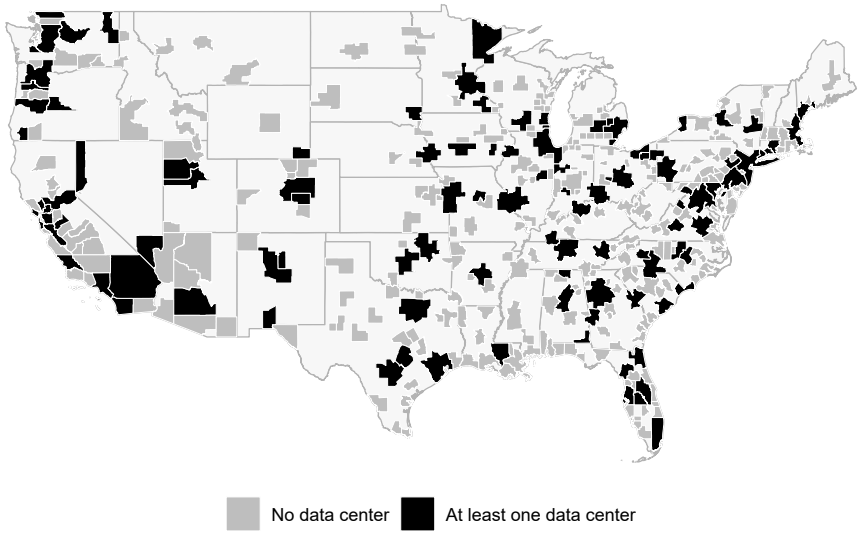
[21] The estimated effects are qualitatively similar with or without this wage control.

2.2.3 Data Centers

The final outcome is the location of data center facilities. I use data from the U.S. Department of Energy's IM3 Open Source Data Center Atlas (Mongird et al. 2025), which reports county locations for 1,242 data center facilities as of 2025. I aggregate these data to the MSA level and construct an indicator for whether an MSA hosts at least one data center. Approximately 27% of MSAs have at least one facility.

Figure 4 displays the geographic distribution of data centers across MSAs. MSAs shaded in black host at least one data center, while those in gray do not.

Figure 4: Data Center Locations



Notes: The map reports whether an MSA hosts at least one data center facility. MSAs shaded in black host at least one facility, and MSAs shaded in gray do not. Data center locations are aggregated to 2017 MSA definitions.

2.2.4 Additional Covariates

The empirical specifications discussed in the next section regress the outcomes above on the centralization index. In these regressions, I include the following sets of covariates. The first set includes geographic controls such as land area, surface-water area, Census Division fixed effects, an indicator for the presence of a port in the MSA, and an index of natural amenities specifically constructed to predict growth (McGranahan 1999). A second set of

controls includes an indicator for the largest government type (county, city, school, or special district) in the MSA and an indicator for whether the MSA is located in the Rust Belt.[22] Finally, I include the number of years between the establishment of the first post office in the MSA and 1902. This variable provides a clean measure of metropolitan age, and controls for features of the Rural Free Delivery instrument detailed below.

3 Empirical Strategy

This section presents the empirical specifications used to estimate the effect of centralization on the outcomes discussed above. The first subsection introduces the estimating equations, while the second details the instrumental-variables strategy and associated first-stage robustness checks.

3.1 Empirical Specifications

This subsection details the empirical specifications I use to estimate the effect of centralization on population and employment growth, local government employment levels, local government finances, and the location of data centers. The value of the centralization index in MSA m from survey year t is denoted by C_{mt}. In the growth equations, I use the centralization index constructed from the joint variation across years and drop the t index. Throughout, I allow for an unobserved determinant of the outcome, ξ_m, that may be correlated with the centralization index. If so, OLS estimates confound the effect of centralization with the effect of the unobserved determinant. I address this issue using the IV strategy described below.

3.1.1 MSA Population and Employment Growth

The population and employment growth equations have the same form. Let Y_{mt} denote either population or employment in MSA m at time t, and let T denote the terminal year, 2017. The specifications come from a multiplicative growth process given by

$$Y_{m,T} = Y_{m,t}^{\beta} \cdot \exp(\bar{g}_T + \epsilon_{mT}),$$

[22] The results remain qualitatively similar when I use the state-based definition of the Rust Belt from Alder, Lagakos, and Ohanian (2023) or omit Rust Belt fixed effects.

where β is a parameter that governs how growth scales with base-year population or employment, $\bar{g}_T$ is the average growth rate across all MSAs, and ϵ is an MSA-specific growth shock. Separating the average national growth from MSA idiosyncratic growth helps highlight what the specification estimates. Suppose that between t and T, all MSAs experienced the same proportional growth after accounting for scale, meaning $\epsilon_{mT} = 0$ for all MSAs m. In this case, there is no variation in growth for centralization to explain. In the data, MSAs do not experience proportional growth, providing meaningful variation in ϵ_{mT} for centralization to explain.

To get to an estimating equation, I take the logarithm and move one unit of $Y_{m,t}$ to the left-hand side, yielding

$$\log \frac{Y_{m,T}}{Y_{mt}} = \bar{g}_T + (\beta - 1) \log Y_{m,t} + \epsilon_{m,T}.$$

To create the empirical specification that I take to the data, I decompose the error term into the effect of centralization, the effect of geographic covariates G_m, metropolitan age A_m, the remaining fixed effects λ_m, the unobserved determinant of the outcome, and a mean-zero idiosyncratic error term as follows:

$$\epsilon_{m,T} = \gamma C_m + \delta G_m + \psi A_m + \lambda_m + \xi_m + u_{m,T}.$$

Substituting the error-term decomposition yields

$$\log \frac{Y_{m,T}}{Y_{mt}} = \bar{g}_T + (\beta - 1) \log Y_{m,t} + \gamma C_m + \delta G_m + \psi A_m + \lambda_m + \xi_m + u_{m,T}.$$

I estimate four equations for each outcome, varying only the base-year index t.

The coefficient of interest is γ, as it represents the effect of increasing centralization. Because both growth and the centralization index are standardized, γ measures the change in growth, in standard-deviation units, caused by a one-standard-deviation increase in centralization. Because the centralization index has no natural scale, only the sign and magnitudes relative to other estimates are identified. Positive rescaling of the index changes the coefficient's magnitude but not its sign. Multiplying the index by a negative value reverses its orientation, turning it into a decentralization index and reversing the coefficient's sign.[23] With a common orientation and scale, the signs and relative magnitudes of the estimates can be compared across specifications.

[23] A unit increase in centralization is equivalent to a unit decrease in decentralization.

3.1.2 Local Government Employment

Motivated by Figure 3, I model local government employment primarily as a function of population, with centralization explaining residual variation. I estimate five specifications: one for each of the four survey years and a pooled specification. Let $Y_{m,t}$ denote total local government employment and $P_{m,t}$ denote population in MSA m at time t. The equation I take to the data is

$$\log Y_{m,t} = \alpha + \gamma C_{m,t} + \beta \log P_{m,t} + \delta G_m + \psi A_m + \lambda_m + \xi_m + u_{m,t}.$$

This estimating equation is similar to the growth equations. However, because I observe a static outcome, local government employment in survey year t, I can use the survey-year-specific centralization index. I estimate the equation separately for each of the four ASPEP survey years and also estimate a pooled specification with survey-year fixed effects. Because of my assumption of a fixed institutional setting throughout the study period, I do not include MSA-level fixed effects in the pooled specification.

3.1.3 Local Government Finance

The specifications I use to estimate the effect of centralization on total revenue and tax revenue largely mirror the specifications for local government employment. Aside from the change in the dependent variable, the only alteration is the inclusion of the log ratio of income to employment. Denoting either total revenue or tax revenue in survey year t as Y_{mt} and the ratio of income to employment as W_{mt}, the estimating equation is

$$\log Y_{m,t} = \alpha + \gamma C_{m,t} + \beta \log P_{m,t} + \phi \log W_{m,t} + \delta G_m + \psi A_m + \lambda_m + \xi_m + u_{m,t}.$$

3.1.4 Data Center Locations

The final empirical specification concerns the location of data centers across MSAs. Because data center locations are observed only once rather than as a panel, they are not indexed by time. Let $D_m = 1$ if MSA m contains a data center and $D_m = 0$ otherwise. I estimate a linear probability model of the form

$$D_m = \alpha + \gamma C_m + \beta \log P_m + \delta G_m + \psi A_m + \lambda_m + \xi_m + u_m.$$

I estimate two additional data center specifications. The first includes private employment growth from 1997 to 2017. If data center siting followed broader firm-location patterns, I would expect this employment-growth control to capture a significant amount of variation that might otherwise be reflected in γ. The second includes a fixed effect for the state of Virginia. Virginia is considered the epicenter of data center construction, accounting for 26% of all data center facilities in my data. The MSAs located in Virginia also exhibit a unique organizational structure because of their heavy reliance on independent cities.

3.2 Instrumental Variables

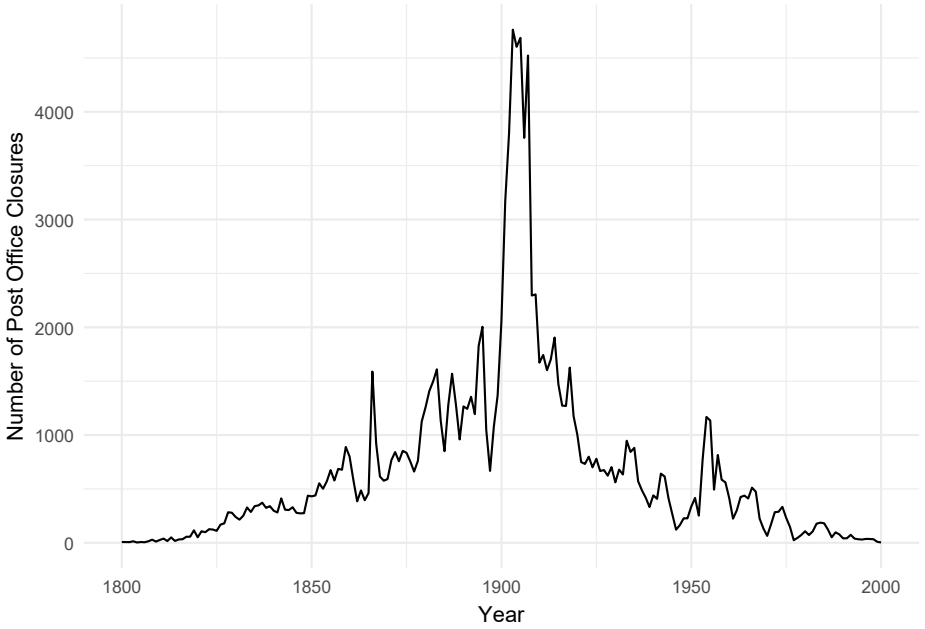
In this subsection, I introduce the Rural Free Delivery instrument, explain how I interpret its effect on modern local government centralization, and report several robustness checks on the first-stage relationship. I then defend the exclusion restriction for both metropolitan growth and the mechanism outcomes.

3.2.1 Rural Free Delivery Instrument

The previous specifications cannot be consistently estimated by OLS if the organization of local government is correlated with unobserved determinants of the outcomes. To address this issue, I employ an instrumental-variables strategy based on the Postal Service's implementation of Rural Free Delivery (RFD) in 1902. Using a complete history of U.S. post offices from Blevins and Helbock (2021), I instrument centralization with the log number of post office closures within each MSA between 1902 and 1910, the period immediately following the implementation of RFD.

Figure 5 illustrates the effect of RFD on post office closures. The vertical axis reports the number of closures, while the horizontal axis reports the year. The figure shows a large increase in closures immediately following the implementation of RFD.

Figure 5: Post Office Closures, 1800–2000



Notes: The figure plots the number of U.S. post office closures by year between 1800 and 2000. The instrument used in the analysis is the log number of post office closures within an MSA between 1902 and 1910.

I interpret the instrument as affecting modern local government organization through the municipal annexation process (Burns 1994). A greater number of post office closures proxies for a greater number of historical rural settlement nodes. A greater number of historical rural settlement nodes creates more potential annexation margins within an MSA. Because incorporation is frequently used as a legal defense against annexation by existing jurisdictions, additional potential annexation margins increase the number of defensive incorporations within the metropolitan area, producing a more decentralized local government structure.

Table 2 reports the first-stage relationship between the RFD instrument and the centralization index. Columns (1)–(4) estimate the first stage separately for each survey-year-specific centralization index, while column (5) reports the pooled specification with survey-year fixed effects. The estimated relationship is negative, large, and stable across survey years: MSAs with more post office closures following RFD are more decentralized nearly a century later.

Table 2: First-Stage Estimates

	Centralization Index				
	2017	2007	1997	1987	Pooled
	(1)	(2)	(3)	(4)	(5)
RFD: Number of Post Office Closures	−0.304***	−0.308***	−0.292***	−0.291***	−0.300***
	(0.050)	(0.051)	(0.053)	(0.053)	(0.025)
F-statistic	33.1	32.64	25.95	25.01	120.38
Observations	378	378	378	378	1,512
R^2	0.612	0.600	0.570	0.567	0.583

*p<0.1; **p<0.05; ***p<0.01

Notes: Each column reports a first-stage regression of the centralization index on the log number of post office closures between 1902 and 1910. Columns (1)–(4) use the year-specific index, and column (5) reports the pooled specification. All specifications follow the first-stage controls described in the text. Standard errors are reported in parentheses.

3.2.2 First-Stage Robustness

In this subsection, I report four robustness checks for the first stage. I first examine heterogeneity by population size, then metropolitan age, and then geography. I conclude by assessing the linearity assumption graphically.

I first examine heterogeneity by population size by interacting the instrument with indicators for quintiles of MSA population in the corresponding survey year. Table 3 reports the results, with each column repeating the exercise for a different survey-year-specific centralization index.

Table 3: First-Stage Estimates: Heterogeneity by Population Size

	Centralization Index			
	2017	2007	1997	1987
	(1)	(2)	(3)	(4)
RFD: Number of Closures × 1st Population Quintile	−0.356***	−0.341***	−0.319***	−0.346***
	(0.060)	(0.060)	(0.063)	(0.064)
RFD: Number of Closures × 2nd Population Quintile	−0.330***	−0.330***	−0.348***	−0.331***
	(0.057)	(0.059)	(0.062)	(0.062)
RFD: Number of Closures × 3rd Population Quintile	−0.294***	−0.316***	−0.286***	−0.291***
	(0.057)	(0.057)	(0.059)	(0.059)
RFD: Number of Closures × 4th Population Quintile	−0.264***	−0.269***	−0.247***	−0.212***
	(0.053)	(0.055)	(0.057)	(0.056)
RFD: Number of Closures × 5th Population Quintile	−0.186***	−0.209***	−0.179***	−0.177***
	(0.053)	(0.054)	(0.056)	(0.055)
Observations	378	378	378	378
R^2	0.602	0.592	0.563	0.562

*p<0.1; **p<0.05; ***p<0.01

Notes: Each column reports a first-stage regression that interacts the log number of post office closures between 1902 and 1910 with indicators for MSA population quintiles. The dependent variable is the year-specific centralization index. Aside from the interaction terms shown, the specifications follow the first-stage controls described in the text. Standard errors are reported in parentheses.

The table shows that the instrument predicts centralization across the MSA population distribution. The estimated effect is larger among smaller-population MSAs, although the difference is modest and the sign remains consistent across quintiles. This pattern suggests that the relationship between historical settlement dispersion and modern local government organization is not driven exclusively by either large or small metropolitan areas.

To examine heterogeneity by age, I repeat the strategy used in Table 3, replacing population quintiles with metropolitan-age quintiles. I measure metropolitan age using the difference between 1902 and the first year that a post office was established in the MSA. Table 4 reports the results and follows the same structure as Table 3.

Table 4: First-Stage Estimates: Heterogeneity by Metropolitan Age

	Centralization Index			
	2017	2007	1997	1987
	(1)	(2)	(3)	(4)
RFD: Number of Closures × 1st Age Quintile	−0.283***	−0.282***	−0.272***	−0.290***
	(0.065)	(0.067)	(0.069)	(0.069)
RFD: Number of Closures × 2nd Age Quintile	−0.321***	−0.320***	−0.317***	−0.320***
	(0.057)	(0.058)	(0.060)	(0.061)
RFD: Number of Closures × 3rd Age Quintile	−0.328***	−0.325***	−0.316***	−0.307***
	(0.054)	(0.055)	(0.057)	(0.057)
RFD: Number of Closures × 4th Age Quintile	−0.282***	−0.284***	−0.263***	−0.257***
	(0.051)	(0.052)	(0.054)	(0.054)
RFD: Number of Closures × 5th Age Quintile	−0.251***	−0.271***	−0.241***	−0.247***
	(0.051)	(0.052)	(0.055)	(0.055)
Observations	378	378	378	378
R^2	0.616	0.602	0.574	0.570

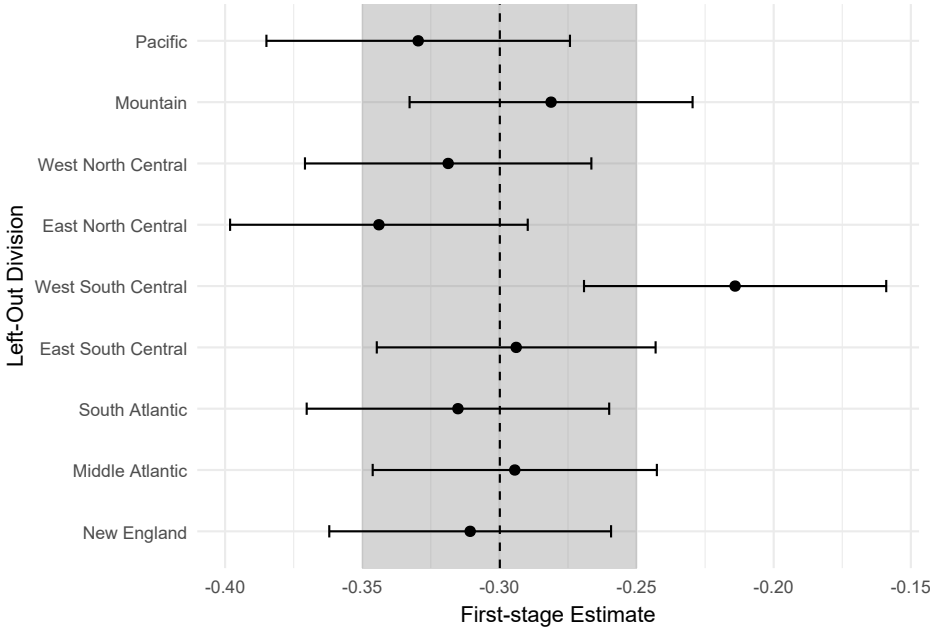
*p<0.1; **p<0.05; ***p<0.01

Notes: Each column reports a first-stage regression that interacts the log number of post office closures between 1902 and 1910 with indicators for MSA age quintiles, where age is measured by the number of years between the establishment of the first post office in the MSA and 1902. The dependent variable is the year-specific centralization index. Aside from the interaction terms shown, the specifications follow the first-stage controls described in the text. Standard errors are reported in parentheses.

The table shows that the effect of the RFD instrument does not vary significantly with metropolitan age. Across the five age quintiles, the first-stage effect of the number of post office closures on centralization is roughly constant, suggesting that the results are not driven by variation in metropolitan age.

To address heterogeneity by geography, I use a leave-one-division-out strategy. I re-estimate the pooled first-stage specification but drop one of the nine Census divisions. Figure 6 displays the results. The pooled first-stage estimate reported in Table 2 is represented by the vertical dashed line, and its 95% confidence interval is captured by the gray rectangle. Along the y-axis, each Census division is listed; the solid point reports the estimated first-stage coefficient when that division is excluded, and the horizontal line reports its 95% confidence interval.

Figure 6: First-Stage Estimates: Geographic Heterogeneity



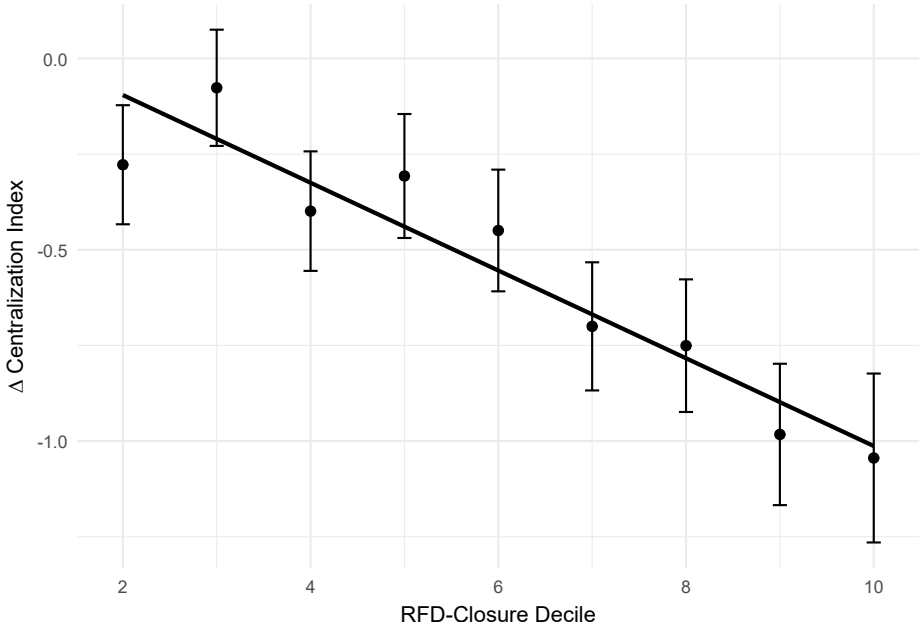
Notes: The figure shows the results of first-stage regressions estimated with individual Census divisions excluded. The x-axis reports the first-stage coefficient on the instrument based on the number of post office closures. The vertical dashed line reports the estimated coefficient for the full sample, and the gray rectangle shows its 95% confidence interval. The y-axis lists the Census division excluded from each regression. The horizontal black lines show the 95% confidence intervals, and the solid black dots show the point estimates.

Figure 6 shows that the first-stage relationship is broadly similar when individual Census divisions are excluded. Leaving any Census division out of the first stage does not significantly alter the estimated coefficient. The largest deviation from the full-sample estimate appears when the West South Central division, which includes Texas, is excluded.[24] While the point estimate is outside the full specification's confidence interval, the confidence intervals overlap. Overall, these results suggest that the relationship between historical settlement dispersion and modern local government organization is not driven by unusually large or small effects from a particular geographic region.

The final robustness check examines the assumption of linearity in two-stage least squares. Instead of modeling the number of post office closures as a continuous instrument, I divide the instrument into deciles and estimate the first stage using the resulting bins. Figure 7 plots the estimated decile effects on centralization along with 95% confidence intervals.

[24] All results remain qualitatively consistent if a Texas fixed effect is included.

Figure 7: First-Stage Estimates by Instrument Decile



Notes: The figure plots coefficients from a pooled first-stage regression that replaces the continuous instrument with indicators for deciles of the number of post office closures. Points report coefficient estimates for deciles 2 through 10 relative to the first decile; vertical bars report 95% confidence intervals; and the solid line is a linear fit through the estimated effects.

The figure shows that the relationship between the number of closures and centralization is well approximated by a linear specification. While the estimated effects are not perfectly monotone across deciles, the relationship is generally decreasing in the number of closures, with no evidence of substantial sign reversals or discontinuities. This pattern suggests that the identifying variation is not driven by a narrow subset of MSAs or by extreme values of the instrument.

3.2.3 Exclusion Restriction for Growth

The exclusion restriction for growth requires that the number of historical rural settlement nodes affects modern growth only through its effect on the organization of local government, rather than through an alternative channel. I consider two main threats to this assumption and discuss additional assumptions that would mitigate each. The first is persistent anti-growth preferences, commonly associated with NIMBYism.[25] This threat arises if the preferences that historically supported a larger number of rural settlement nodes persist over time and independently lead modern local governments to adopt anti-growth policies.

[25]Not In My Back Yard.

In that case, the instrument would capture the effect of persistent preferences in addition to the effect of local government centralization. One way to mitigate this concern is to view local government organization as a technology through which preferences are expressed. If decentralization makes it easier for residents to translate anti-growth preferences into policy, then the resulting policies and their effects on growth operate through the organization of local government rather than through a separate channel.

A second threat is persistent spatial polycentricity. Historical rural settlement nodes may reflect a more dispersed spatial organization of economic activity, which could itself persist and independently affect modern metropolitan growth. If so, the instrument would capture the effect of economic geography in addition to the effect of local government organization.

This alternative channel requires that historical spatial patterns persist into the modern period and that this persistence independently affects growth. The existing literature provides limited support for either condition. First, Kim (1995) documents substantial changes in the spatial organization of economic activity over this period. Historical dispersion would therefore need to survive extensive firm entry and exit, residential mobility, infrastructure investment, and technological change throughout the twentieth century. Second, the relationship between modern spatial organization and modern growth remains inconclusive (Li and Liu 2022). Finally, granting that dispersed economic activity has an effect on modern growth, its persistence still requires explanation. If local government boundaries impede the relocation or reorganization of economic activity, then this persistence operates through local government organization rather than through an independent channel.

3.2.4 Exclusion Restriction for Mechanisms

The exclusion restriction is more direct for the mechanism outcomes because local government employment and revenue are contemporary decisions made by governments, while data center locations respond partly to the tax and regulatory environment those governments create. A violation therefore requires historical settlement dispersion to affect these outcomes independently of the organization of local government.

For local government employment and revenue, a direct historical channel would require historical settlement patterns to continue to determine how many public workers an MSA employs and how much revenue its governments collect. The most plausible alternative, higher service costs arising from lower density and greater sprawl, predicts the opposite sign. Dispersed development should make public-service provision more labor-intensive and costly, increasing employment and revenue rather than reducing them.

Data centers impose an even greater burden on an alternative explanation. A direct historical channel must explain why settlement dispersion makes an MSA especially attractive to data centers while reducing economic activity more broadly. General differences in land availability, infrastructure, or development patterns are insufficient unless they both create this industry-specific advantage and persist independently of contemporary fiscal and regulatory policy.

The joint pattern therefore provides the strongest defense of the exclusion restriction. Historical settlement dispersion would need to independently produce lower public employment, lower total and tax revenue, greater data center investment, and lower metropolitan growth. That combination does not follow naturally from settlement dispersion alone. It follows directly from a common mechanism in which decentralization constrains public employment and revenue collection while intensifying competition for highly mobile, tax-sensitive capital.

4 Results

In this section, I present the IV estimates for the effect of centralization. I begin by estimating the effect of centralization on population and employment growth. I then examine its effects on local government employment, total revenue, and tax revenue. I conclude the empirical section by estimating the effect of centralization on data center siting decisions.

The tables report only the estimated effect of centralization on each outcome and omit other covariates from the presentation. Each regression corresponds to the specification outlined in Section 3. The pooled specifications additionally include survey-year fixed effects. Because the construction of the index removes natural units from the measure of centralization, the estimates identify only the sign and relative magnitudes of the underlying relationships. The index is standardized and oriented so that the coefficients correspond to a one-standard-deviation increase in centralization. All outcomes except the data center indicator are standardized, so the coefficients measure the standard-deviation change in the outcome caused by a one-standard-deviation increase in centralization. Standard errors are clustered at the state level. Each MSA is assigned to the state containing its largest government unit.

Each table also reports two diagnostic statistics and the OLS estimate of the effect of centralization. The first diagnostic, the Kleibergen–Paap rk Wald F-statistic, evaluates the relevance and strength of the excluded instrument while remaining robust to heteroskedasticity. Larger values indicate that the instrument more strongly predicts centralization (Kleibergen and Paap 2006).

The second diagnostic is the Anderson–Rubin p-value (Anderson and Rubin 1949; Andrews, Stock, and Sun 2019), which provides weak-instrument-robust inference by testing whether a candidate parameter value is consistent with the reduced-form relationship between the instrument and the outcome. The full procedure constructs confidence sets by testing a sequence of candidate parameter values. The p-value reported in the tables corresponds to the test in which the candidate value is zero. A small value indicates that zero would not be included in the confidence set even without relying on a strong first stage. Together, these diagnostics assess instrument strength and whether the estimated effects remain distinguishable from zero under weak-instrument-robust inference.

I first present the estimated effect of centralization on population and employment growth in Table 5. Panel A reports the effects of centralization on population growth. Columns (1)–(4) report the effect of centralization for different base years, with column (1) reporting the shortest time horizon, 2007–2017, and column (4) reporting the longest horizon, 1977–2017. Panel B of Table 5 follows the same pattern, with employment growth as the dependent variable; its longest horizon is 1978–2017.

Table 5: Effect of Centralization on Population and Employment Growth

	Standardized Growth Rate from Base Year to 2017			
	Base: 2007	Base: 1997	Base: 1987	Base: 1977
	(1)	(2)	(3)	(4)
Panel A: Population Growth				
Centralization Index	0.391**	0.499**	0.464*	0.469*
	(0.162)	(0.231)	(0.245)	(0.243)
KP rk Wald F-statistic	29.31	28.31	28.4	28.28
Anderson–Rubin p-value	0.0589	0.0189	0.0279	0.0181
Centralization OLS Estimate	0.18***	0.18***	0.21***	0.23***
Observations	378	378	378	378
Panel B: Employment Growth				
Centralization Index	0.570***	0.702***	0.391**	0.306*
	(0.208)	(0.191)	(0.170)	(0.177)
KP rk Wald F-statistic	32.73	31.37	30.89	31.05
Anderson–Rubin p-value	0.0125	9e-04	0.0522	0.1245
Centralization OLS Estimate	0.12*	0.22***	0.22***	0.23***
Observations	378	378	378	378

*p<0.1; **p<0.05; ***p<0.01

Notes: The table reports two-stage least squares estimates of the effect of centralization on metropolitan population growth (Panel A) and private employment growth (Panel B). Columns differ by the base year used to construct growth through 2017. Each regression follows the growth specification described in the text. Kleibergen–Paap rk Wald F-statistics and Anderson–Rubin p-values are reported below the coefficient estimates. Standard errors are clustered at the state level and reported in parentheses.

Table 5 reports positive coefficients for centralization in both panels. In Panel A, the population growth coefficients are positive in all four specifications. In Panel B, all four employment growth coefficients are positive and statistically significant. Across both outcomes and specifications, a one-standard-deviation increase in centralization is estimated to increase growth by approximately two-fifths of a standard deviation. The positive relationship lends more support to the coordination framework than to the competition framework. The sign of the estimated effect corresponds to the key comparative static on which these frameworks differ: models emphasizing competition between local governments predict that greater decentralization increases growth through interjurisdictional competition, while models emphasizing coordination predict that greater centralization increases growth by internalizing externalities.

I next examine local government employment as a measure of labor input use. Table 6 reports four cross-sectional specifications and one pooled specification for the effect of centralization on local government employment. Column (1) reports the estimate using the

2017 ASPEP data. Columns (2)–(4) report the corresponding estimates for 2007, 1997, and 1987, respectively. Column (5) reports the pooled specification using the centralization index constructed from the joint variation over time.

Table 6: Effect of Centralization on Local Government Employment

	Local Government Employment				
	2017	2007	1997	1987	Pooled
	(1)	(2)	(3)	(4)	(5)
Centralization Index: 2017	0.094**				
	(0.047)				
Centralization Index: 2007		0.101**			
		(0.045)			
Centralization Index: 1997			0.136**		
			(0.061)		
Centralization Index: 1987				0.125**	
				(0.050)	
Centralization Index: Pooled					0.110**
					(0.047)
KP rk Wald F-statistic	31.44	31	24.65	23.76	118.63
Anderson–Rubin p-value	0.0144	0.0053	5e-04	7e-04	0
Centralization OLS Estimate	-0.0081	-0.0068	0.0022	0.0053	-0.0021
Observations	378	378	378	378	1,512

*p<0.1; **p<0.05; ***p<0.01

Notes: The table reports two-stage least squares estimates of the effect of centralization on log local government employment. Columns (1)–(4) report year-specific cross sections, and column (5) reports the pooled specification. Each regression follows the employment specification described in the text. Kleibergen–Paap rk Wald F-statistics and Anderson–Rubin p-values are reported below the coefficient estimates. Standard errors are clustered at the state level and reported in parentheses.

Table 6 reports positive coefficients in every specification, implying that centralized MSAs employ more local government workers. All five estimates are statistically significant and similar in magnitude. They imply that a one-standard-deviation increase in centralization raises local government employment by roughly one-tenth of a standard deviation.

I next examine local government finances. Table 7 contains two panels structured identically to Table 6. Panel A reports the effect of centralization on total revenue, while Panel B reports the effect on tax revenue.

Table 7: Effect of Centralization on Total and Tax Revenue

	Dependent variable:				
	2017	2007	1997	1987	Pooled
	(1)	(2)	(3)	(4)	(5)
Panel A: Total Revenue					
Centralization Index: 2017	0.211***				
	(0.067)				
Centralization Index: 2007		0.250***			
		(0.087)			
Centralization Index: 1997			0.187**		
			(0.075)		
Centralization Index: 1987				0.130*	
				(0.068)	
Centralization Index: Pooled					0.180***
					(0.050)
KP rk Wald F-statistic	30.32	30.14	24.22	25.76	119.91
Anderson–Rubin p-value	0.0014	0	0.0041	0.0613	0
Centralization OLS Estimate	0.021	0.028	0.025	0.0041	0.016**
Observations	378	378	378	378	1,512
Panel B: Tax Revenue					
Centralization Index: 2017	0.167**				
	(0.067)				
Centralization Index: 2007		0.220***			
		(0.083)			
Centralization Index: 1997			0.163**		
			(0.076)		
Centralization Index: 1987				0.157**	
				(0.071)	
Centralization Index: Pooled					0.160***
					(0.061)
KP rk Wald F-statistic	30.32	30.14	24.22	25.76	119.91
Anderson–Rubin p-value	0.0023	0	0.0035	0.0074	0
Centralization OLS Estimate	0.021	0.054***	0.03*	0.025	0.03***
Observations	378	378	378	378	1,512

*p<0.1; **p<0.05; ***p<0.01

Notes: The table reports two-stage least squares estimates of the effect of centralization on log total revenue (Panel A) and log tax revenue (Panel B). Columns (1)–(4) report year-specific cross sections, and column (5) reports the pooled specification. Each regression follows the finance specification described in the text. Kleibergen–Paap rk Wald F-statistics and Anderson–Rubin p-values are reported below the coefficient estimates. Standard errors are clustered at the state level and reported in parentheses.

Table 7 reports positive coefficients for both total revenue and tax revenue in every specification. In both panels, the estimates are statistically distinguishable from zero at conventional levels and are similar in magnitude across specifications. The estimates suggest that a one-standard-deviation increase in centralization raises both total revenue and tax revenue by about one-sixth of a standard deviation.

Taken together, Tables 6 and 7 suggest that the positive growth effects associated with centralization are unlikely to arise from the narrow economies of size mechanism where more technically efficient centralized MSAs promote growth by passing cost-savings onto their constituents. The pattern is more readily reconcilable with a mechanism where centralization allows MSAs to internalize the externalities that underlie tax competition.

However, an extended economies of size mechanism where greater technical efficiency induces a rebound effect that increases provision is still possible. In order to partially distinguish between the tax competition and extended economies of size mechanisms, I examine location decisions made by highly mobile and tax-sensitive firms. As a concrete case, I examine where firms constructing data centers choose to locate. Column (1) of Table 8 reports the regression results of an indicator for the presence of a data center in the MSA on the centralization index; column (2) conditions on private employment growth from 1997 to 2017; and column (3) includes a fixed effect for the state of Virginia, the epicenter of data center construction.

Table 8: Effect of Centralization on Data Center Location Decisions

	Dependent variable:		
	Data Center in MSA		
	(1)	(2)	(3)
Centralization Index: 2017	−0.197**	−0.259**	−0.246**
	(0.092)	(0.114)	(0.106)
Private Employment Growth: 1997–2017		0.386**	0.363**
		(0.171)	(0.160)
Virginia Fixed Effect			0.239**
			(0.096)
KP rk Wald F-statistic	31.44	26.19	26.89
Anderson–Rubin p-value	0.0351	0.015	0.0185
Centralization OLS Estimate	0.037	0.031	0.029
Observations	378	378	378

*p<0.1; **p<0.05; ***p<0.01

Notes: The table reports two-stage least squares estimates of the effect of centralization on whether an MSA hosts at least one data center facility. The dependent variable is an indicator equal to one if the MSA contains at least one data center. The three regressions follow the data center specifications described in the text. Kleibergen–Paap rk Wald F-statistics and Anderson–Rubin p-values are reported below the coefficient estimates. Standard errors are clustered at the state level and reported in parentheses.

Table 8 reports a negative and statistically significant coefficient for centralization. The estimate shows that data centers are less likely to locate in centralized MSAs. When private employment growth is taken into account, the estimated effect increases. The estimates directly align with the prediction of tax competition models. Data centers are highly mobile and tax-sensitive, making them particularly sensitive to the coordination failures that underlie tax competition models. In contrast, the extended economies of size mechanism provides no explanation as to why data centers in particular would behave differently than the other firms that drive the employment growth results in Table 5. While not able to fully distinguish between these two mechanisms, the results suggest that externalities have a large enough effect on the behavior of local government to affect firm decision-making.

5 Conclusion

This paper studies the effect of local government organization on metropolitan growth. I measure centralization across U.S. metropolitan statistical areas using the distribution of employment across local governments and estimate its effect using an instrument based on post office closures induced by the introduction of Rural Free Delivery in 1902. The

instrument captures historical settlement dispersion, which I argue shaped modern local government organization through the processes of annexation and defensive incorporation. Using this strategy, I find that more centralized MSAs experience higher population and employment growth. I then examine local government employment, total revenue, tax revenue, and data center locations to distinguish among competing mechanisms. The results show that centralized MSAs employ more local government workers and collect more revenue, while decentralized MSAs are more likely to attract data centers, a form of highly mobile, tax-sensitive capital.

Metropolitan areas act as integrated economic units, but their political authority is often fragmented across many governments. Local government organization matters because governments do not act in isolation, and differences in organization alter the strategic environment in which local governments operate. The evidence presented in this paper suggests that an important feature of local government interactions is the potential for coordination failure. When each local government considers only the costs and benefits within its own jurisdiction, competition over the same mobile residents, firms, and fiscal base can leave cross-jurisdictional externalities unaccounted for, reducing aggregate fiscal capacity and limiting the provision of public goods whose benefits extend across jurisdictional boundaries. In this setting, centralization can improve outcomes by moving decision-making closer to the scale of the metropolitan economy and internalizing spillovers that decentralized governments do not fully take into account.

The mechanism evidence strengthens this interpretation. If centralization raised growth mainly through a narrow economies of size mechanism, we would expect centralized MSAs to use fewer public inputs conditional on population. Instead, they employ more local government labor and collect more revenue. This pattern suggests that centralization is not simply reducing the cost of a fixed bundle of public goods. It appears to expand the fiscal and administrative capacity used to provide them. At the same time, decentralized MSAs are more likely to attract data centers, even though they are less likely to attract other economic activity, as evidenced by their lower private employment growth. This pattern strengthens the case for coordination failures. Fragmented jurisdictions in decentralized MSAs may compete aggressively for mobile capital, but this competition can reduce aggregate revenue and, in turn, weaken metropolitan growth.

Centralization need not have the same effects at every scale of government. At broader geographic scales, the costs of centralization may be substantial: preferences may be more heterogeneous, local information is likely to be harder to aggregate, accountability may weaken, and spillovers across jurisdictions may be less intense. In those settings, decentralization may

improve outcomes by allowing policy to vary across places and by preserving competitive discipline among governments. Metropolitan areas present a different tradeoff. Local governments within an MSA are geographically close and economically linked through shared labor markets, infrastructure networks, the mobility of residents and firms, and overlapping tax bases. Because these interactions occur within a tightly connected spatial economy, the spillovers from local policy choices are likely to be more salient. The results of this paper suggest that, at the metropolitan scale, centralization can improve outcomes when the benefits of internalizing these spillovers exceed the costs of reducing interjurisdictional competition.

This paper's focus on metropolitan-level growth is useful because growth summarizes many dimensions of local government performance in a single outcome. However, that aggregation also limits what the estimate can show. The estimated effect of centralization on growth does not isolate a single mechanism; it reflects the combined effects of centralization across many margins, including public safety, schools, infrastructure, land use, taxation, and regulation. The aggregate estimate therefore helps identify which framework of local government behavior is most consistent with the broad pattern of outcomes, but it should not be taken to imply that coordination failures are always the driving factor or that all local public-good provision is more efficient under centralization. A natural next step is to move beyond the aggregate growth estimate and examine the particular public goods and policy margins through which local government organization shapes economic activity.

6 References

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