Coordination vs. Competition: Local Government Centralization and Metropolitan Growth

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Abstract

This paper examines how local government centralization shapes metropolitan growth. I measure centralization using the distribution of local government employment across jurisdictions within U.S. metropolitan statistical areas. To address endogeneity, I use an instrument based on post office closures caused by the introduction of Rural Free Delivery in 1902. If find that more centralized MSAs experience greater population and private employment growth. Examining mechanisms, I find that centeralized MSA 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 spillovers externalities may be a more important margin than increased competition between governments for producing growth.

1 Introduction

A central question in public economics asks how the organization of government shapes public-good provision. One way organization may matter is by altering the incentives and constraints governments face when choosing policy. A key dimension of this 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 in this debate concerns the role of competition between governments, motivating the adoption

of two distinct conceptual frameworks[1]. The decentralized position emphasizes the *competition* framework, which focuses on the information and incentives of public officials as in Tiebout (1956), Besley and Case (1995), and Hayek (1945). In contrast, the centralized position adopts the *coordination* framework and worries that externalities can turn competition between governments into a "race to the bottom." Central to the coordination framework are tax competition models developed in Wilson (1986), Zodrow and Mieszkowski (1986), and Wildasin (1988).[2]

While the debate occurs across all levels of government[3], it is particularly salient at the local level. In this paper, I measure local government centralization using the distribution of local government employment within metropolitan statistical areas (MSAs) of the United States. I then estimate the effect of centralization on metropolitan growth using cross-sectional variation across MSAs.[4] I find that centralization increases both population and employment growth across multiple time horizons using both OLS and an IV strategy based on the Postal Service's introduction of Rural Free Delivery in 1902. I then attempt to uncover the mechanisms behind centralization's positive effect on growth by examining local government employment, tax revenue, and data center siting across MSAs. Data centers provide a close empirical analog to the mobile capital found in tax competition models.

The questions surrounding local government centralization have direct policy relevance, not just academic interest. 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 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

[1]Both frameworks assume that interactions between governments are strategic in the sense that one jurisdiction's policy is in part a response to the policy chosen by another.

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

[3]Two notable arenas concern decentralizing authority from national governments to states, and centralizing national authority to a supranational institution, such as the European Union.

[4]Temporal variation in the organization of local government is sparse. I show below that measured centralization is highly stable across time. Examining the few discrete changes in centralization, as is done in the political merger literature discussed below, requires strong assumptions to rule out both selection and spillovers across unmerged units within the same MSA.

by local governments.

I confront two main empirical challenges in estimating the effect of centralization. The first is measurement, as centralization is not directly observed but is instead a latent institutional concept. Centralization could be measured using several features of local governments, including government counts or fiscal variables. Using the Annual Survey of Public Employment and Payroll, I focus on how local government employment is distributed across government units within an MSA.

Even within this narrower focus, there are many possible measures of centralization. This is closely related to the broader problem of measuring concentration. As documented by Ahern, Kong, and Yan (2024), different measurement strategies for concentration can produce substantially different conclusions.[5] I build on their insight and methodology by constructing a family of centralization statistics from a vector of local government employment shares within each MSA, similar to how they create a family of concentration statistics from industry market shares[6]. I then aggregate these statistics into an index that captures their common component.[7] However, by employing an index, I lose natural units and scale, meaning that only the sign and relative magnitudes across specifications are identified.

The second empirical challenge concerns endogeneity and its interaction with mediators of growth. Many variables that appear to be potential sources of omitted-variable bias may instead be channels through which centralization affects growth. For example, human capital likely affects growth, but centralization may directly affect the quality of public schools (Hoxby 2000) and thereby influence the accumulation of human capital. Similar arguments apply to crime if centralization affects the efficiency of policing; to infrastructure if centralization affects public investment; and to zoning if centralization affects land-use regulation. 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 limits the scope of the endogeneity problem, it does not eliminate it. My measurement strategy captures the concentration of local government employment across legally distinct governments, but it abstracts from both the taxonomy of local government types and the internal organization of individual governments. Two MSAs with the same

[5]They show that, depending on the concentration measure, industrial concentration has either increased or decreased between 1990 and 2020.

[6]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).

[7]This methodology is mainly used for parsimony and completeness. In the appendix, I show that the qualitative results do not depend on this specific construction. Using existing measures of centralization alone or altering the aggregation scheme produces similar results.

measured level of centralization may allocate public employees differently across counties, municipalities, school districts, and special districts. Similarly, an individual government unit may contain many, possibly competitive, internal departments. If these unmeasured organizational features affect the efficiency of public-good provision and are correlated with measured centralization, then they represent a remaining source of omitted-variable bias.

To address the endogeneity problem, I use an IV strategy based on the Postal Service's implementation of Rural Free Delivery (RFD) in 1902. Prior to RFD, residents of areas considered rural at the turn of the century typically collected mail at local post offices, whose viability depended on the postal activity generated by surrounding communities. With the implementation of RFD, mail was delivered free of charge to every mailbox, and the resulting loss of revenue caused a large wave of post office closures. Because RFD extended the 1863 Free City Delivery program to rural areas (United States Postal Service 2024), these closures were disproportionately concentrated outside existing large cities and therefore provide a proxy for historical settlement dispersion. Using a complete census of post offices from Blevins and Helbock (2021), I instrument centralization with the number of post office closures in each MSA between 1902 and 1910.

The RFD post office closure instrument is highly predictive of modern local government decentralization. I interpret the instrument as operating through the municipal annexation process because many jurisdictions incorporate as a legal defense against annexation attempts by established jurisdictions (Burns 1994). Because a greater number of post office closures proxies for a larger number of historically dispersed settlements, it increases the number of potential annexation margins within an MSA. As the number of annexation opportunities rises, so does the number of opportunities for defensive incorporation, producing a more decentralized local government structure. The identifying assumption is that the historical settlement dispersion proxied by post office closures affects modern outcomes only through its effect on the organization of local government.

Using this 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, and for all but the shortest population-growth horizon, greater centralization increases growth.

While the growth result lends empirical support to the centralized position in the debate, further work is required to determine why centralization would increase growth. The co-

ordination framework emphasizes the role of externalities, but it does not rule out alternative channels such as cost reductions from economies of scale in public-good provision. The remaining empirical exercises examine whether the data are more consistent with a coordination-failure mechanism or a scale-based efficiency channel.

I confront two related challenges in testing for economies of scale. First, I do not observe public-good output. Second, public goods are inherently multidimensional bundles. Even if all dimensions were observable, differences in provision could reflect tradeoffs involving unobserved preferences rather than production efficiency. For example, higher-quality roads in one MSA may reflect a stronger preference for roads over parks rather than more efficient provision. To partially address this issue, I examine local government employment as a measure of input use. I find that more centralized MSAs employ more local government workers.

Next, I find that more centralized MSAs collect both more total revenue and more tax revenue conditional on population. This pattern matches a central prediction of tax competition models. In decentralized MSAs, taxing authority is fragmented across jurisdictions, intensifying competition for a mobile tax base. Strategic complementarities in local tax reaction functions can suppress equilibrium tax rates and aggregate revenue when fiscal spillovers are not internalized. By internalizing these externalities, more centralized MSAs can sustain higher revenue collection.

The joint findings weaken the case for a pure economies-of-scale interpretation. Greater economies of scale should allow governments either to produce a given level of public goods at lower cost or to produce a greater level of public goods at constant cost. Instead, centralized MSAs appear to expand both input and output relative to decentralized MSAs of the same population size. This pattern directly contradicts a mechanism based only on labor-saving efficiency gains and is difficult to reconcile with a pure scale-based mechanism. A more natural interpretation is that decentralized governments face constraints on public-good provision, and that centralization relaxes these constraints in a way predicted by the coordination framework.[8]

In a final empirical exercise, I examine firms' data center location decisions and find that centralized MSAs are less likely to host these facilities. Given the positive effect of centralization on aggregate employment growth, the negative effect on data centers initially appears

[8]The findings also emphasize the importance of examining mechanisms for policy. Advocates of political mergers often cite reducing wasteful duplication and lowering tax burdens as a likely result of the merger. While the findings here can be viewed as supporting political mergers, the mechanisms show that increased centralization will typically increase local government employment and tax burdens by reducing tax competition.

contradictory[9]. However, relative to most other types of firms, the benefits generated by data centers are local to the jurisdiction in which the facility is located, while the costs are dispersed throughout the MSA through higher electricity prices (Alvarez et al. 2026). These features make the siting decisions of data centers especially sensitive to coordination failures among local governments.[10]

As a useful analogy, the firm's location decision resembles a seller's decision about which auction to enter. More decentralized MSAs, with a larger number of local governments, resemble auctions with more competing bidders. If decentralization lowers the effective tax or regulatory cost faced by data centers through increased tax competition, then firms have an incentive to locate disproportionately in decentralized MSAs, where governments compete more aggressively for investment. This result therefore need not conflict with the employment growth findings. Instead, it can be interpreted as evidence that coordination failures are salient enough to affect firm location decisions.

Taken together, the positive relationship between centralization and growth suggests that the coordination framework better describes local government behavior than the competition framework within this setting. The empirical evidence further suggests that centralization increases growth primarily by alleviating coordination failures rather than by exploiting economies of scale. Because interactions among local governments are inherently spatial, externalities are particularly salient within MSAs. These externalities can distort competition across decentralized jurisdictions and inefficiently suppress public-good provision. By internalizing these spillovers, centralized local governments can increase the provision of public goods that contribute to higher growth.

This work contributes to several literatures, beginning with research on political mergers. Because political mergers generally increase centralization[11], 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.

[9]The estimated effect increases when private employment growth is included in the specification.

[10]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 is to experience the negative effects, such as higher electric 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.

[11]While political mergers are likely to increase latent centralization, mergers need not increase measured centralization, especially when using a single statistic. At the time of writing, the proposal for a merger between St. Louis City and St. Louis County in Missouri entails the city being reunited within the county as a municipality, which would not reduce the number of governments. Further, shifting employment from the city, the current largest government unit, to the county would reduce the one-government concentration ratio, a step toward decentralization.

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[12] that complicate merger-based analyses.

This paper also relates to the broader literature on government centralization and institutional performance. A large macro-level literature studies how cross-country variation in fiscal and political centralization between national and subnational governments affects growth and institutional outcomes (Barro 1991; Weingast 1995; Sala-i-Martin 1997; Davoodi and Zou 1998; Treisman 2000; Rodden 2004; Iimi 2005; Baskaran and Feld 2013; Cassidy and Velayudhan 2026). In contrast to my findings, this literature often concludes that decentralization promotes economic growth (Dillinger 1994). One interpretation is that the relevant tradeoffs differ by scale. At the national level, decentralization may intensify competition across large and geographically separated jurisdictions (i.e., states), where spillovers are less salient. At the metropolitan level, however, the dominant concern may instead be the internalization of spillovers among closely linked governments, making greater centralization more effective.

Finally, 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 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.

The remainder of the paper proceeds as follows. Section 2 presents the data on local government organization, including the measurement of centralization and the outcomes I study.

[12]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 the evaluation of the merger, especially if evaluations are made after a short time horizon.

Section 3 outlines the empirical strategy and describes the instrumental-variables strategy used for identification. Section 4 reports the results and discusses their implications for understanding the mechanisms of centralization. Section 5 concludes.

2 Data

The main unit of analysis in this study is the metropolitan statistical area (MSA), as defined by the Office of Management and Budget. An MSA consists of a group of counties organized around an urban core with substantial economic integration, as measured by commuting flows. 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.

I focus on metropolitan areas rather than central cities alone because of their growing importance in contemporary urban economies. 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 the most counties: 29. Because commuting patterns and urbanization evolve over time, MSA boundaries change across decades. To maintain a consistent geographic unit, I 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 unit.

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 centralization indices discussed next.

2.1 Local Government Centralization Index

Centralization refers to the degree to which decision-making authority is delegated throughout 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 MSA. 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 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 how to map 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

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

Using this notation, ${}^{0}\lambda$ gives the number-of-governments measure, ${}^{2}\lambda$ is the inverse HHI, and $\lim_{q\to\infty} {}^{q}\lambda$ approaches the inverse one-government concentration ratio.[14] A convenient interpretation of these statistics is as "effective numbers" or "number-equivalent measures"

[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.

[14] 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 $|s|$.

(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 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[15], there is no theoretical reason for choosing one statistic over another. For that reason, I evaluate 50 statistics at uniformly spaced points[16] from the family of centralization statistics for each MSA and survey year.

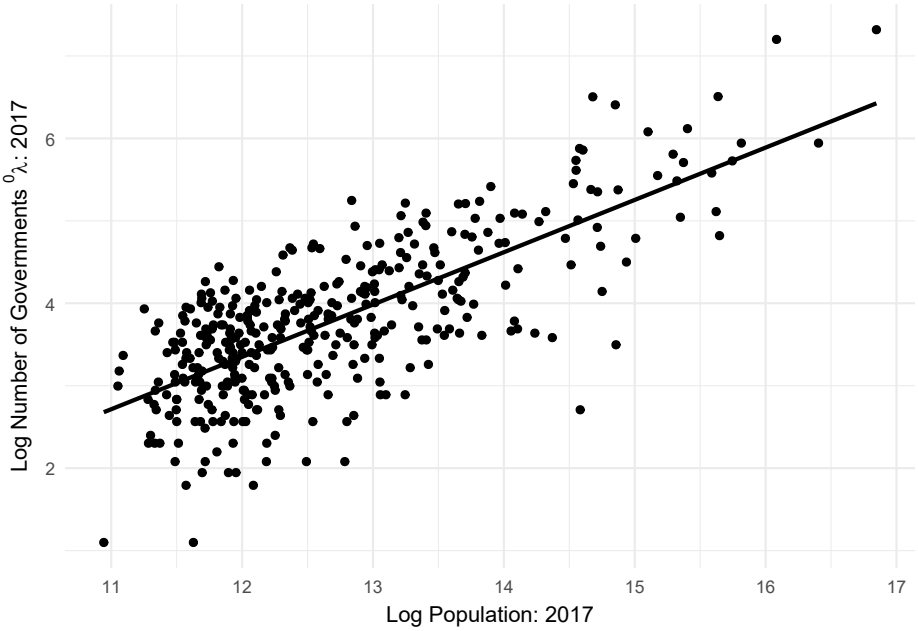
Directly comparing the sequence of effective numbers for each MSA can be misleading because it does not account for differences in MSA population or land area. For example, the number-of-governments statistic scales with population size. Due to this, a fixed number of governments can represent different levels of centralization depending on population.[17] Similarly, land area may capture geographic differences across MSAs that affect the interpretation of centralization. Figure 1 highlights this dependence. The horizontal axis plots log MSA population in 2017 and the vertical axis plots the log number of governments in 2017.

[15]Shannon entropy, ${}^1\lambda$, is unique in that it obeys Shannon (1948)'s grouping axiom. This "forces" the use of that statistic, but there is no obvious reason to enforce that axiom in this setting.

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

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

Figure 1: 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. The final step of the process is to aggregate the residualized statistics into a single index. Letting ${}^{q(i)}\tilde{\lambda}_{mt}$ represent a residualized centralization statistic i for MSA m in survey year t, 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 must select weights. I choose the weights using principal component analysis[18] because the goal is to measure the common centralization signal contained in the full curve, not to privilege any single value of q. Each statistic emphasizes a different part of the local government employment distribution: low values of q place more weight on the number of governments, while high values place more weight on the largest governments. PCA provides a disciplined way to combine these measures by extracting the dominant dimension of variation they share across MSAs. In practice, this means the index is high for MSAs that are consistently more centralized across the curve, rather than for MSAs that

[18]I show in the appendix that the results do not depend on the weighting scheme. Indices based on a simple average or a factor model have correlations greater than 0.95 with the PCA index, and both produce qualitatively similar results.

appear centralized only under one particular definition. 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.

As noted in the introduction, an alternative strategy would be to examine variation in centralization within each MSA over time. However, I find that centralization is highly persistent within MSAs. In Table 1, the 2017 centralization index is regressed on the indices from 2007, 1997, and 1987 in columns (1)–(3), respectively. Perfect persistence between the indices would correspond to an estimated coefficient of one.

Table 1: Persistence in Local Government Centralization

	Dependent variable:		
	Centralization Index: 2017		
	(1)	(2)	(3)
Centralization Index: 2007	0.988***		
	(0.008)		
Centralization Index: 1997		0.968***	
		(0.013)	
Centralization Index: 1987			0.946***
			(0.017)
Observations	378	378	378
R^2	0.976	0.936	0.895

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

Notes: Each column reports an OLS regression of the standardized 2017 centralization index on the standardized centralization index from the year listed in the row label. The sample includes 378 MSAs. Standard errors are reported in parentheses.

The results indicate a high degree of persistence. MSAs that are more decentralized in earlier periods tend to remain decentralized in later periods. This persistence motivates the cross-sectional approach used to estimate the effect of centralization.

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. I define

growth as the log difference between population in 2017 and population in a base year. 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 normalize growth rates to represent the average 10-year growth rate.

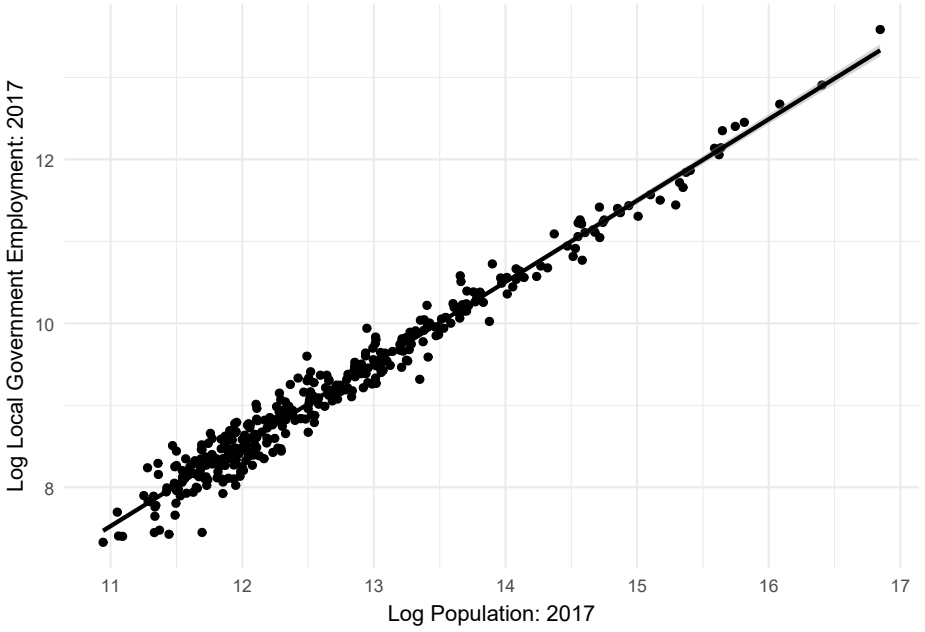
I construct employment measures using the Business Dynamics Statistics dataset. Because these data begin in 1978, I use employment levels for 1978, 1987, 1997, 2007, and 2017. Employment growth is measured analogously as the log difference between employment in 2017 and employment in a base year.

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 2 plots log local government employment against log population for 2017.

Figure 2: 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 (CoG), which complements the ASPEP by providing financial information for each government unit. I extract total revenue and tax revenue for each government unit. 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.

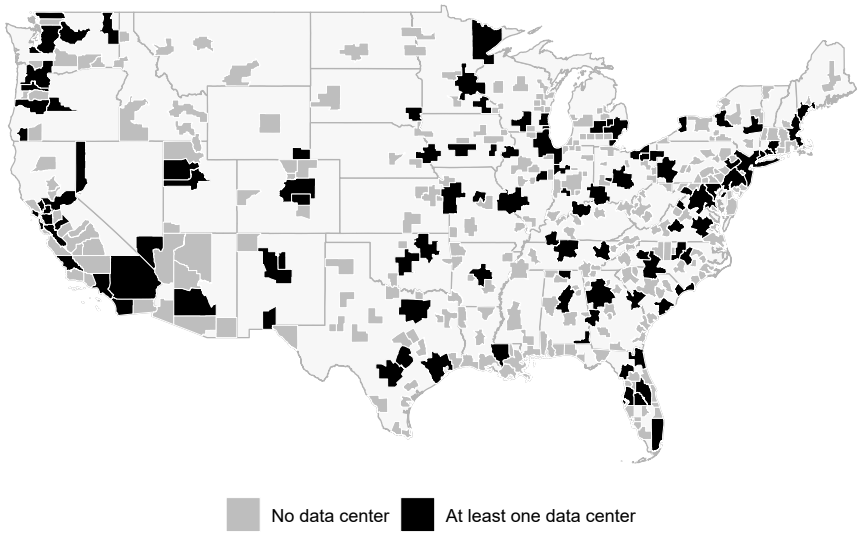
A major difference between the local government employment data and the financial data is the unit of measurement. Local government employment is measured in workers, whereas finances are measured in dollars. This creates a measurement issue because the value of a dollar varies significantly across geographic areas (Moretti 2013). To account for differences in local price levels, I use BEA income data and BDS employment data to estimate the average wage level, which helps capture differences in total and tax revenue across MSAs that are unrelated to government organization.

2.2.3 Data Centers

The final outcome is the location of data center facilities. I use data from the IM3 Open Source Data Center Atlas, which reports county locations for 1,242 data center facilities as of 2026. 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 3 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 3: 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 Time-Invariant Covariates

The baseline specifications discussed in the next section regress the outcomes above on the centralization index. In these regressions, I include a set of time-invariant covariates. These include land area and surface-water area, an indicator for the largest government type (county, city, school, or special district) in the MSA, and an indicator for the presence of a port in the MSA. I also include the number of years between the establishment of the first post office in the MSA and 1902. This variable serves a dual role by controlling for MSA age and 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 causal 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

$$Y_{m,T} = Y_{m,t}^{\beta} \cdot \exp(\epsilon_{m,T}).$$

where β is a parameter that governs how growth scales with base-year population or employment, and ϵ is an MSA-specific growth shock. Taking the logarithm and moving one unit of $Y_{m,t}$ to the left-hand side yields

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

To create the empirical specification that I take to the data, I decompose the error term into its mean, the effect of centralization, the effect of time-invariant covariates X_m, Census division fixed effects λ_m, the unobserved determinant of the outcome, and a mean-zero idiosyncratic error term as follows:

$$\epsilon_{m,T} = \alpha + \gamma C_m + \delta X_m + \lambda_m + \xi_m + u_{m,T}.$$

Substituting the error-term decomposition yields

$$\log \frac{Y_{m,T}}{Y_{mt}} = \alpha + (\beta - 1) \log Y_{m,t} + \gamma C_m + \delta X_m + \lambda_m + \xi_m + u_{m,T}.$$

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. As mentioned in the introduction, only the sign and magnitudes relative to other estimates are identified. Note that the centralization index has a standard deviation of one by construction. This is an arbitrary standardization that directly affects the estimated value of γ. If the standard deviation were increased, the estimated value would decrease, and the opposite effect would occur for a decrease. However, the sign of the effect is identified despite any arbitrary standardization. Multiplying the index by a negative value simply reorients the measure as a decentralization index on a different scale and changes only the interpretation.[19] When the same standardization is used across specifications, the estimates can be compared with one another, allowing their relative magnitudes to be identified.

3.1.2 Local Government Employment

Motivated by Figure 2, I model local government employment primarily as a function of population, with centralization explaining residual variation conditional on population. 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

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

This estimating equation is similar to the growth equations. 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 Table 1 shows that centralization is highly persistent across time, 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

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

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

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

3.1.4 Data Center Locations

The final empirical specification concerns the location of data centers across MSAs. Because data center locations are measured contemporaneously rather than in 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

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

I estimate two additional data center specifications. The first includes private employment growth from 1997 to 2017. If data centers behaved similarly to other firms, I would expect this momentum term 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 Rural Free Delivery Instrument

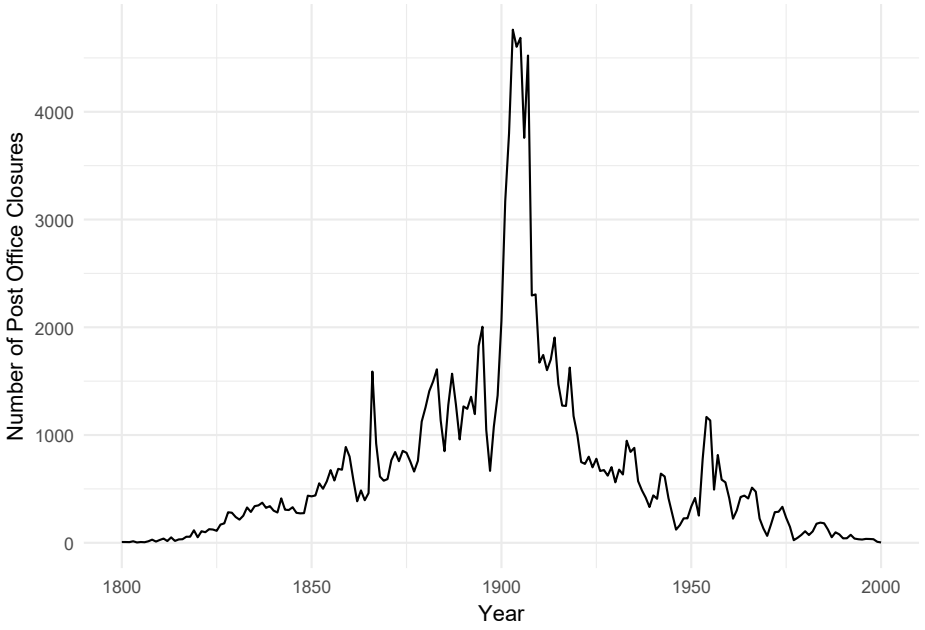
The previous specifications cannot be consistently estimated by OLS if the organization of local government is correlated with unobserved determinants of the outcomes. To address this issue, I employ an instrumental-variables strategy based on the Postal Service's implementation of Rural Free Delivery (RFD) in 1902. Using the complete history of U.S. post offices, 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.

The identifying variation exploited by the instrument comes from differences in historical settlement dispersion across MSAs, as measured by post office closures. Because RFD extended the 1863 Free City Delivery program to rural areas, the post offices that closed following its implementation were disproportionately located in communities that were rural in the early twentieth century. Prior to RFD, residents in these communities typically collected mail

at local post offices whose viability depended on surrounding postal activity. By shifting delivery away from local pickup and toward direct household service, RFD reduced the need for many small and geographically dispersed post offices. Variation in the number of closures across MSAs therefore proxies for differences in early settlement dispersion.

Figure 4 illustrates the effect of RFD on post office closures. The vertical axis reports the number of post office closures, while the horizontal axis reports the year. The figure displays a large spike in closures during the period immediately following the implementation of RFD.

Figure 4: 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. A greater number of historically dispersed settlements creates more potential annexation margins within an MSA. Because incorporation is frequently used as a legal defense against annexation, more potential margins increase the likelihood of defensive incorporation somewhere 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, which includes survey-year fixed effects.

Table 2: First-Stage Estimates

	Centralization Index				
	2017	2007	1997	1987	Pooled
	(1)	(2)	(3)	(4)	(5)
RFD: Number of Post Office Closures	−0.321***	−0.317***	−0.302***	−0.297***	−0.310***
	(0.048)	(0.048)	(0.050)	(0.051)	(0.024)
F-statistic	41.52	39.47	31.81	29.42	144.72
Observations	378	378	378	378	1,512
R^2	0.619	0.608	0.578	0.574	0.591

*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.

The RFD instrument strongly predicts later decentralization. The identifying assumption is that historical settlement dispersion, measured by post office closures, affects the modern outcomes I study only through its effect on the organization of local government. A primary threat to the exclusion restriction is that historical settlement patterns may directly affect modern outcomes through persistent private spatial organization and that this dispersed organization may independently affect growth. I present two arguments against this objection. First, private economic activity is highly dynamic, with firms adjusting through entry, exit, and relocation. Sustaining a direct channel from historical settlement dispersion to modern outcomes through private spatial organization for more than a century would therefore require substantial and persistent frictions. To the extent that such frictions matter, they are most plausibly mediated through local government organization itself, making government organization the persistence mechanism rather than an independent channel.

Second, even if historical settlement dispersion generated persistent private spatial organization, such persistence would only violate the exclusion restriction if the resulting polycentric organization exerted an independent effect on growth. However, the literature examining the relationship between polycentric metropolitan structure and economic growth remains inconclusive (Li and Liu 2022). Taken together, these considerations suggest that any long-run persistence arising from historical settlement patterns is more plausibly mediated through the durable organization of local government than through persistent private spatial organization alone.

3.2.1 First-Stage Robustness

In this subsection, I report four robustness checks for the first stage. I examine heterogeneity by population size, metropolitan age, and geography, and then assess the linearity and monotonicity assumptions. To examine heterogeneity by population size, I interact 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	−0.395***	−0.372***	−0.358***	−0.376***
	(0.057)	(0.057)	(0.060)	(0.062)
RFD: Number of Closures × 2nd Population Quintile	−0.348***	−0.334***	−0.355***	−0.352***
	(0.055)	(0.057)	(0.060)	(0.059)
RFD: Number of Closures × 3rd Population Quintile	−0.318***	−0.333***	−0.308***	−0.296***
	(0.055)	(0.054)	(0.057)	(0.057)
RFD: Number of Closures × 4th Population Quintile	−0.281***	−0.276***	−0.253***	−0.217***
	(0.052)	(0.053)	(0.055)	(0.055)
RFD: Number of Closures × 5th Population Quintile	−0.216***	−0.230***	−0.200***	−0.191***
	(0.052)	(0.052)	(0.054)	(0.053)
Observations	378	378	378	378
R^2	0.608	0.600	0.567	0.569

*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	−0.299***	−0.290***	−0.284***	−0.296***
	(0.063)	(0.064)	(0.066)	(0.066)
RFD: Number of Closures × 2nd Age Quintile	−0.334***	−0.327***	−0.328***	−0.328***
	(0.055)	(0.056)	(0.058)	(0.059)
RFD: Number of Closures × 3rd Age Quintile	−0.344***	−0.333***	−0.327***	−0.312***
	(0.051)	(0.052)	(0.054)	(0.054)
RFD: Number of Closures × 4th Age Quintile	−0.309***	−0.300***	−0.280***	−0.267***
	(0.048)	(0.048)	(0.050)	(0.051)
RFD: Number of Closures × 5th Age Quintile	−0.268***	−0.276***	−0.250***	−0.247***
	(0.047)	(0.048)	(0.050)	(0.050)
Observations	378	378	378	378
R^2	0.623	0.610	0.582	0.577

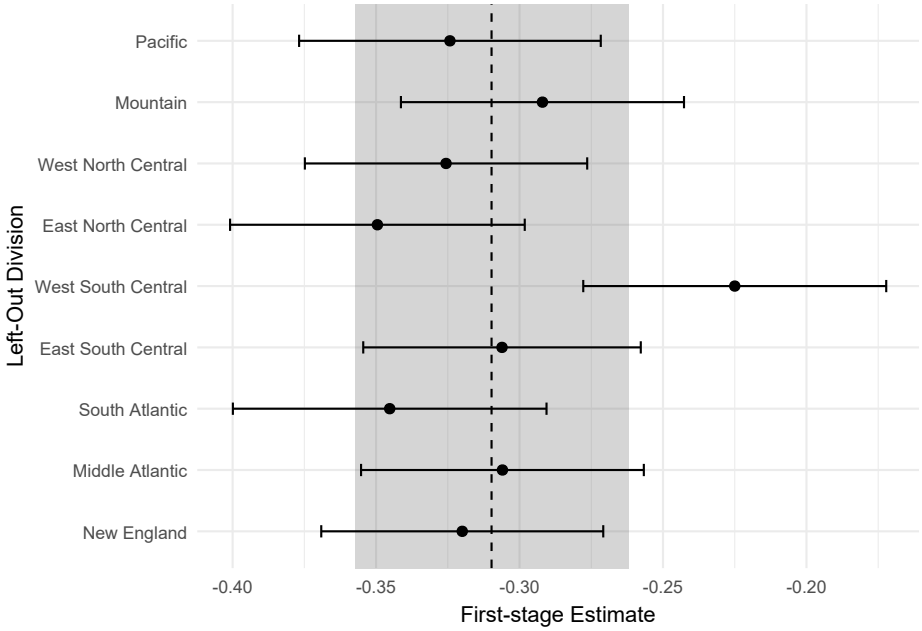
*p<0.1; **p<0.05; ***p<0.01

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 5 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 5: First-Stage Estimates: Geographic Heterogeneity



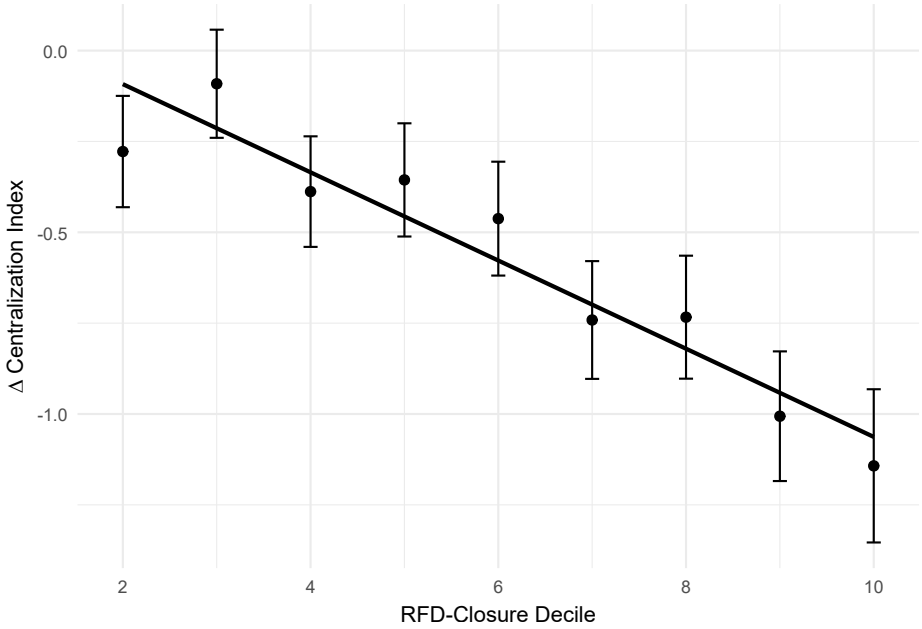
Notes: The figure shows the results of first-stage regressions estimated with individual Census divisions excluded. The x-axis reports the first-stage coefficient on the instrument based on the number of post office closures. The vertical dashed line reports the estimated coefficient for the full sample, 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 5 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[20]. 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 assumptions of linearity and monotonicity. 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 6 plots the estimated decile effects on centralization along with 95% confidence intervals.

[20] All results stay qualitatively consistent if a Texas fixed effect is included.

Figure 6: First-Stage Estimates by Instrument Decile



Notes: The figure plots coefficients from a pooled first-stage regression that replaces the continuous instrument with 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 is consistent with an approximately monotone first-stage relationship and suggests that the identifying variation is not driven by a narrow subset of MSAs or by extreme values of the instrument.

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 and includes log population, log land area, log surface-water area, time-

invariant covariates, and Census division fixed effects. 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. 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 and a longest horizon of 1978–2017.

Table 5: Effect of Centralization on Population and Employment Growth

	10-Year 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	0.018	0.036**	0.045**	0.060***
	(0.012)	(0.016)	(0.019)	(0.023)
KP rk Wald F-statistic	35.52	35	35.86	36.15
Anderson–Rubin p-value	0.1821	0.0225	0.0082	6e-04
Centralization OLS Estimate	0.011**	0.013**	0.017***	0.025***
Observations	378	378	378	378
Panel B: Employment Growth				
Centralization Index	0.041**	0.065***	0.037**	0.047***
	(0.017)	(0.014)	(0.015)	(0.017)
KP rk Wald F-statistic	39.94	39.4	39.33	40.04
Anderson–Rubin p-value	0.0245	3e-04	0.0289	0.0104
Centralization OLS Estimate	0.01	0.021***	0.021***	0.025***
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, although the 2007–2017 estimate is not statistically distinguishable from zero. In Panel B, all four employment growth coefficients are positive and statistically significant. Across both measures, the positive relationship between centralization and growth is more consistent with coordination models than with Tiebout-style models of competition. 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	0.089**				
	(0.045)				
Centralization Index: 2007		0.103**			
		(0.043)			
Centralization Index: 1997			0.116**		
			(0.056)		
Centralization Index: 1987				0.093*	
				(0.049)	
Centralization Index: Pooled					0.099**
					(0.044)
KP rk Wald F-statistic	39.66	37.7	30.38	28.09	142.81
Anderson–Rubin p-value	0.0216	0.0059	0.0032	0.0162	0
Centralization OLS Estimate	-0.0067	-0.0045	0.00099	0.0023	-0.0024
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 and imply that centralized MSA employ more local government workers. All five estimates are statistically significant, and the coefficients are similar in magnitude across columns. The positive relationship between centralization and local government employment suggests that the growth effects documented in Table 5 are unlikely to arise from a mechanism in which labor-saving economies of scale allow centralized MSAs to increase expected utility by passing cost savings on to residents.

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	0.209***				
	(0.058)				
Centralization Index: 2007		0.263***			
		(0.078)			
Centralization Index: 1997			0.181***		
			(0.059)		
Centralization Index: 1987				0.143**	
				(0.063)	
Centralization Index: Pooled					0.199***
					(0.059)
KP rk Wald F-statistic	35.74	34.58	28.49	29.52	137.87
Anderson–Rubin p-value	3e-04	0	0.0011	0.0202	0
Centralization OLS Estimate	0.025	0.02	0.022	0.0088	0.019**
Observations	378	378	378	378	1,512
Panel B: Tax Revenue					
Centralization Index: 2017	0.208***				
	(0.077)				
Centralization Index: 2007		0.295***			
		(0.105)			
Centralization Index: 1997			0.190**		
			(0.086)		
Centralization Index: 1987				0.178**	
				(0.083)	
Centralization Index: Pooled					0.218***
					(0.079)
KP rk Wald F-statistic	35.74	34.58	28.49	29.52	137.87
Anderson–Rubin p-value	0.0014	0	0.0038	0.0113	0
Centralization OLS Estimate	0.038*	0.077***	0.037*	0.029	0.046***
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 positive relationship between centralization and tax revenue is consistent with models in which decentralized governments fail to internalize the effects of taxation on neighboring jurisdictions' tax bases, leading to lower equilibrium tax collection.

Taken together, Tables 6 and 7 suggest that the positive growth effects associated with centralization are unlikely to arise from a simple mechanism based on labor-saving economies of scale. More centralized MSAs employ more local government labor and collect more revenue conditional on population. This pattern is more consistent with centralization relaxing coordination constraints on public-good provision than with reducing input use through scale economies.

I conclude by examining data center locations, which provide a close empirical analog for mobile capital in the tax competition framework. Column (1) of Table 8 reports the baseline estimate of the effect of centralization; 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	−0.201**	−0.257**	−0.244**
	(0.090)	(0.110)	(0.103)
Private Employment Growth: 1997–2017		0.372**	0.350**
		(0.170)	(0.161)
Virginia Fixed Effect			0.221**
			(0.091)
KP rk Wald F-statistic	39.66	33.15	33.65
Anderson–Rubin p-value	0.0187	0.0081	0.0106
Centralization OLS Estimate	0.025	0.02	0.018
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 and suggests that decentralized MSAs compete more aggressively for highly mobile investment, consistent with the predictions of tax competition models. When private employment growth is taken into account, the estimated effect becomes larger in magnitude and more precise. More broadly, the result suggests that coordination failures associated with fragmented local government organization are sufficiently strong to affect firm location decisions.

Taken together, the results suggest that the organization of local government plays an important role in shaping metropolitan economic outcomes. More centralized MSAs exhibit higher population and employment growth, greater local government employment, and higher revenue collection conditional on population. At the same time, decentralized MSAs appear more likely to attract highly mobile investment in the form of data centers. Viewed jointly, these patterns suggest that decentralization may intensify competition for mobile capital while limiting the ability of governments to coordinate taxation and the provision of public goods at the metropolitan level. Overall, the results are more consistent with models emphasizing coordination failures under fragmented local government organization than with models in which decentralization primarily improves outcomes by fostering competition between local governments.

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 fragmentation 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 labor 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 jurisdictions. Local government organization matters because governments do not act in isolation; organizational changes alter the strategic environment in

which they 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 labor-saving economies of scale, 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 may compete aggressively for mobile capital, but competition can result in lower aggregate revenue which leads to weaker 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, permitting, taxation, environmental regulation, and the reallocation of resources across space. 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.

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