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Tax Competition with Intermunicipal Cooperation

Abstract

We study local tax competition when municipalities can voluntarily cooperate. We compare the intensity of interjurisdictional policy interdependence between competing municipalities within the same “establishment for inter-municipal cooperation” (EIMC) and competing municipalities outside of the cooperative unit. To resolve the endogeneity of the decision to cooperate we apply the approach of Kelejian and Piras (2014). The strategic response to the average tax rate among peer members of the same EIMC is less intense than the response to the average tax rate of municipalities outside of the cooperative unit. A one percentage point decrease in the average tax rate of non-members lowers the own-jurisdiction tax rate by 0.58 percentage points, while a one unit decrease in the tax rate of towns within the EIMC lowers the own-jurisdiction rate by 0.31 percentage points. Our empirical methods can be used to study strategic interactions within other cooperative groups, including supra-national institutions such as the European Union.

JEL-Codes: C200, H200, H700, R500.

Keywords: tax competition, intermunicipal cooperation, spatial autoregressive models, endogenous weight matrix, local public finance, networks.

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Decentralized governments *compete* with each other when tax bases are mobile; however, local, state, and even national governments often *cooperate* with each other. This cooperation may limit the parasitic effects of tax competition. Cooperation can differentially affect the nature of competition with collaborators (peers) and non-collaborators. Failing to account for heterogeneous effects of the decision to cooperate may lead to an incorrect understanding of the policy effectiveness of cooperation. To answer the broad economics question of how endogenous *cooperation* influences the *competitive* interactions between various players, we focus on municipal tax competition in France.

A literature in public and urban economics uses spatial models (Brueckner and Saavedra 2001, Brueckner 2003, Buettner 2003, Revelli 2006, Revelli 2015, Parchet 2019; Brueckner 2022; Ferraresi 2022) to study competition among many local jurisdictions (Agrawal, Hoyt and Wilson 2022; Wilson 1999). This literature, however, does not take into account the fact that governments often cooperate in the provision of public services and raise taxes from pooled or shared tax bases. Jurisdictions entering a cooperative agreement often maintain substantial taxing powers, implying that tax base mobility and fiscal spillovers within and outside the cooperative will both still exist. How does the decision to cooperate affect tax competition and strategic interactions between jurisdictions?¹ Do jurisdictions within the same endogenously formed cooperative group compete differently with each other than with jurisdictions not within the cooperative group?

Although a literature theoretically studies the effect of coalition formation on policies (Burbidge et al. 1997; Konrad and Schjelderup 1999), the effect of cooperation on the competitive determination of local policies is ambiguous. The empirical impact of cooperation on strategic interactions remains unknown. On the one hand, cooperation may *reduce* strategic interactions as members of a cooperative agreement share information and have common policy goals, lessening tax competition. The fact that some public services are being provided by the cooperative group may also reduce tax competition, as municipalities need less revenues to finance the remaining services and cooperative units may also have grants that help redistribute revenues across jurisdictions. As an additional

¹We use the phrases “strategic interaction” and “tax competition” interchangeably. As discussed in Agrawal, Hoyt and Wilson (2022), tax competition is likely the dominant force.

mechanism, leakage of the tax base outside of the cooperative group shrinks not only the tax base of the municipality but also the tax base of the cooperative. Mobility of the tax base outside of the cooperative unit thus imposes “double harm” (on the municipality and on the cooperative), so pressure to respond to external jurisdictions may be higher. Leakage of the tax base to other towns within the cooperative shrinks the tax base of the municipality, but does not change the total base of the cooperative.

On the other hand, cooperation may *increase* strategic interactions. Nearby jurisdictions are more likely to compete intensely for locally mobile factors (Janeba and Osterloh 2013), and proximate jurisdictions are likely to be members of the same cooperative group. The local linkages such as commuting or segmented capital markets that prevent factors from being globally mobile may create intense tax competition among nearby peer municipalities (Wildasin 2014). Cooperation may even intensify these linkages. Finally, the formation of cooperative entities may facilitate tax collusion among municipalities, which could result in municipalities tacitly agreeing to match their nearby jurisdiction tax changes, amplifying the appearance of tax interdependence within the cooperative.

We focus on intermunicipal cooperation in France. France has historically granted both local tax autonomy and choice over whether and with whom local governments will cooperate, leading to the modern “establishment for inter-municipal cooperation” (EIMC). French municipalities can set tax rates on four main tax bases: the business, residence, developed property and undeveloped property tax bases. After joining an EIMC, some tax powers are surrendered to the cooperative unit. Yet, municipalities maintain taxation authority allowing for municipal tax competition to persist. We measure the existence of and intensity of strategic reactions using the slope of the reaction function.

Studying tax competition with cooperation poses econometric challenges. First, much of the prior literature—capturing interactions among neighboring jurisdictions using a spatial lag model with a single (weighted) average tax rate variable—does not allow for tax competition to differ among various types of competitors. In our case, we relax this by allowing for heterogeneous responses to tax changes within and outside of the endogenously formed cooperative by estimating the slope of two different tax reaction

functions in our estimating equation. Second, standard estimates of a tax competition reaction function slope assume that the spatial weight matrix used to construct the average tax rate of the neighbors of one municipality, in our case the members of the EIMC to which this municipality belongs, is exogenous.² Yet, whether a municipality is a member of the EIMC is a choice, as is who the other member municipalities are. This means that the weight matrix determining neighbors within the EIMC is endogenous. The limited set of studies that allow for heterogeneity in the intensity of tax competition rely on exogenous partitions of the spatial weight matrix using country or state borders (Cassette, DiPorto and Foremny 2012; Geys and Osterloh 2013), and are not concerned with endogenous group formation via inter-municipal cooperation.

To resolve endogeneity of the weight matrix, we apply Kelejian and Piras (2014). This approach requires having variables that predict the bilateral relationship between each pair of jurisdictions in addition to standard instruments that resolve the endogeneity of the spatial lag (tax rate) terms. In the first step, we predict the elements of the weight matrix and in the second step, we instrument for the weighted average of competitor rates. To predict the spatial weights, we construct an original archival history of syndicate membership (a less integrated form of cooperation) among municipalities. We use historical syndicate cooperation from 1900 to the 1970s to predict whether municipalities are in an EIMC today. Given the distinct nature of these syndicates, they signal an interest in places to cooperate but are not related to current day tax decisions. After predicting the elements of the weight matrix, in the second step, we use those predicted weights to construct spatial lags of the exogenous covariates as instruments for competitor tax rates.

We apply this approach on a panel of over 36,000 municipalities from 1994 to 2009. We find that tax interdependence among municipalities in France is important. Overall, tax rates in one jurisdiction are strategic complements with tax rates in competitor jurisdictions. However, when decomposing strategic interactions between municipalities

²Loosely speaking, the spatial weight matrix specifies the structure of each bilateral relationship between jurisdictions in the dataset. Each row represents a town and each column is its potential competitors. When two jurisdictions are not linked in a given way, they receive a weight of zero. More intense linkages between jurisdictions receive larger weights.

within the same EIMC and municipalities outside of the EIMC, tax interdependence for municipalities within the EIMC is much less intense. Municipalities within a cooperative unit respond less intensively when municipalities within their EIMC raise or lower tax rates. Critically, a one unit decrease in the average tax rates of non-members lowers own-jurisdiction tax rates by 0.58 percentage points, while a one unit decrease in peer tax rates lowers the own-jurisdiction rate by 0.31 percentage points. This suggests that municipalities are more concerned with non-group members “stealing” the tax base than “peer” municipalities within the cooperative.

1 Relation to Prior Literature

Policymakers have debated what the most effective policy responses are to limit tax competition. As a result, the literature has seen an increase in analyses of tax coordination and cooperation, including intermunicipal cooperation. While the literature on intermunicipal cooperation has studied its effectiveness in various countries, none of those studies have analyzed the heterogeneous effects on tax competition.³ In this section, we relate our work to the existing work on intermunicipal cooperation in France, focusing on papers that have studied the decision to cooperate or the effect of that cooperation on tax rates.

Several recent papers study intermunicipal cooperation in France. However, none of these papers study the *heterogeneous* effect of cooperation on *tax competition* among peer and non-peer municipalities. The slope of the reaction function can differ across institutional settings and various tax instruments. Even focusing on competition over a given tax rate and institutional setting, the reaction function slope can differ across jurisdictions due to many asymmetries across jurisdictions. For example, a city may respond differently to a suburb than the suburb might respond to a city due to one having an inflow of commuters and the other having an outflow. And as noted in the introduction, there are many theoretical reasons why members of an EIMC may respond differentially to each other than to towns outside the EIMC. Further, most studies assume

³For recent surveys on intermunicipal cooperation, see Hulst and van Montfort (2007), Agrawal, Poterba and Zidar (2024), Breuillé and Duran-Vigneron (2023), and Agrawal (2023).

the decision to cooperate is exogenous to the tax competition game.

Breuillé, Duran-Vigneron and Samson (2018) uses the same tax rate panel data as ours to analyze the effect of cooperation on the *level* of tax rates. The basic design is a difference-in-differences design, where treatment is determined by formation of an EIMC. The authors then show that municipal tax rates fall after joining an EIMC, while total (municipal plus EIMC) tax rates rise.⁴ While it may be tempting to argue that a decline in local tax rates alters tax competition, there is no structural link between the level of tax rates and tax competition. Intuitively, tax rates may fall simply because of economies of scale in public good provision. Nor can the paper identify the differential effect of competition with member and non-member jurisdictions.⁵

Charlot, Paty and Piguet (2015) directly study the effect of cooperation on overall tax competition, showing that tax competition is reduced with cooperation (but not showing if this is due to a change in competition among members or nonmembers of the EIMC). Using a spatial lag model, they regress the own-tax rate on a spatial lag of tax rates, allowing the spatial lag to interact with an indicator for whether the own-jurisdiction cooperates under a single business taxation regime. Unlike our paper, this does not provide a decomposition of the competitive effects between members and non-members. More generally, the decision to cooperate and the tax regime to adopt is (implicitly) treated as exogenous and not modeled as a choice. A major point of our paper is the decision to cooperate is endogenous to the tax competition game—and it will differentially affect competition with members and non-members.

Finally, in recent work, Tricaud (2021) studies municipalities that resisted joining an EIMC—and only did so after being forced to do so. These municipalities saw an increase in construction and a decrease in public services after being forced to cooperate. These effects are not apparent in places that voluntarily cooperated much earlier in time. Her study provides some evidence on why (or why not) municipalities may join an

⁴In an online appendix, the model is supplemented with a single spatial lag model, but its purpose is to make sure the main estimating equation does not suffer from omitted variable bias. Furthermore, this spatial lag is not decomposed into peer on non-peer effects.

⁵Frère, Leprince and Paty (2014) use spatial tools to study the effect on expenditures. Breuillé and Le Gallo (2017) look for spatial dependence in a cross-section, but mainly focus on non-tax related variables, again without the novelty of our decomposition.

EIMC—and that it is a strategic choice. The loss of control in urban planning suggests that holdout municipalities did not cooperate because they wanted to avoid construction occurring in the town (NIMBYism). The decline in public services suggests that smaller municipalities may have worried that cooperation would result in the shutdown of services and instead transfer them to larger towns. While this work sheds light on what might explain the decision to cooperate, the study does not have a focus on tax policy—but provides justification for the need to predict tax cooperation with exogenous factors.

The novelty of our paper relative to this literature is twofold. First, we are the first to show that strategic interactions among peer members of the same EIMC are less intense than strategic interactions with municipalities outside of the cooperative unit. Although prior studies have analyzed tax competition overall, we decompose the intensity of competition across members and nonmembers of the EIMC, documenting that cooperation can have heterogeneous effects on competition. Second, we develop a novel procedure that addresses the endogeneity of the decision to cooperate in the tax competition game.

2 Institutions

2.1 Types of Cooperation

France is a relatively decentralized country with about 36,000 municipalities with tax and spending authority. Formal municipal cooperation dates back to 1890. Initially, municipalities formed syndicates to exploit scale economies in providing a unique public good or service (clean water, sewage treatment, waste collection, etc.). These historical cooperative units could not set tax rates, but rather were financed directly by contributions from their members. The post-World War II period proved to be a very important time period for the growth of municipal cooperation in France. For example, almost 2000 municipalities created a syndicate in 1955. Figure A.1 shows the syndicates *created* from 1900 to 1955 and the syndicates *created* from 1900 to 1974. These syndicates had no taxation authority and municipalities could leave the syndicates at any time. Their duties were limited to provision of facilities, infrastructure, or networks. Although the majority

of syndicates contain between 3 and 15 municipalities each, 17% of syndicates over the period 1955-1974 are bilateral agreements between two municipalities.

Recognizing the need for more formal cooperation, the French government passed a series of laws designed to encourage what we label as current day inter-municipal cooperation with own-source tax revenue. The purpose of these reforms was to grant taxation authority and a wider range of potential functions than the historical syndicates. The law requires EIMCs to comprise a group of contiguous municipalities without any enclaves and prohibits membership in more than one EIMC. However, municipalities are free to make decisions concerning EIMC membership: they are free to leave, form their own EIMC, or join an existing EIMC.⁶

2.2 Tax Authority

In terms of taxation, there are four main local taxes in France: “taxe foncière sur les propriétés bâties” (tax on developed properties: τ^{DP}), “taxe foncière sur les propriétés non bâties” (tax on undeveloped properties: τ^{UP}), “taxe d’habitation” (resident tax: τ^R) and “taxe professionnelle” (business tax: τ^B). These taxes are generally taxes on capital and land, and described in great detail in Appendix A.1, but we next provide some details on the taxes necessary to understand competition over them.⁷

The four taxes are, in varying proportions, all based on the cadastral *rental* value of various properties, which is intended to be a proxy of the annual rents that would be collected if the property were rented out. It is calculated by multiplying a “reference local valuation rate” (i.e., an average local rate per square meter computed in 1970 and updated each year using flat-rate discounting and revaluation coefficients) with the property size weighted by the property’s features. According to a report published in 2017 by the Public Finance Department of the French Ministry of the Economy and Finance, based on a study in 5 “départements”, rental values for private residential property would increase by 151.5% if market rental values were used.

⁶Figure A.2 provides some background on the evolution and size of EIMCs.

⁷The business tax was replaced in 2010 by the Contribution Economique Territoriale (CET), which comprises a tax on the rental value of properties and a tax on value added.

The residence tax is paid by the occupant. Although numerous exemptions are based on the occupant's income, this is essentially a property tax on occupied homes or units (rented and owner-occupied). The tax on developed land and the tax on undeveloped land are applied to the cadastral rental value after abatement, respectively 50% and 20%. Unlike the residence tax, the tax on developed land is on a larger set of built-up properties (for instance, including commercial properties in addition to most occupied residential properties), is levied whether the property is occupied or not, and is paid by the owner. The tax rate on undeveloped land is payable by all owners of undeveloped land. Finally, the business tax was based on a composite tax base comprising company revenues, the property rental value of taxable assets including buildings, the rental value of movable assets and equipment, and originally, salaries. All-in-all, the combination of the taxes acts as standard taxes on capital and property, each covering different types.

If a municipality is not a member of an EIMC, it may levy all four tax rates across these different bases without restriction. Once in an EIMC, the municipal taxing powers depend on one of three “regimes” with respect to how taxes are split between the EIMC and municipality. Table 1 summarizes these possible taxation regimes where T^l is the EIMC rate and τ^l is the municipal tax rate on the tax base l . Under a “single taxation” regime, the business tax is surrendered to the EIMC, but the municipality maintains complete autonomy over the other three tax rates. Under an “additional taxation” regime, the municipality and the EIMC cohabit all four tax bases: the municipality can maintain its taxation authority on all four bases but the EIMC taxes the same base and engages in vertical tax competition (Keen 1998; Devereux, Lockwood and Redoano 2007). Finally, under a “mixed taxation” regime, the business tax is set entirely by the EIMC but the municipality and the EIMC jointly cohabit the remaining three bases. To deal with this differing choice of regimes across EIMCs, all regression specifications will include EIMC regime fixed effects (dummies for the EIMC regimes in Table 1), so that all effects are identified within a common EIMC structure.

Table 2 summarizes the tax rates. Although the taxes are on capital and land, the tax rates in the table might appear higher than property tax rates in the United

States. There are several reasons for this. First, there are very large flat-allowances of 50% for some properties. Second, property rental values are often obsolete, being based on historical values that may adjust over time at non-market rates. Third, and most importantly, all tax bases are determined based on rental prices and not values. Converting between rental prices and market values using Poterba (1992) implies that the effective tax rates on values would be substantially lower.⁸ As a result of these smaller bases, some municipalities set very high tax rates, which can even exceed 100%.

To give an idea of the magnitude of the taxes, consider the developed property tax. For example, for a 50 square meter flat with a reference rate of 14€, the yearly (taxable) cadastral rental value after abatement will be 4,200€ ($50 \times 14 \times 12 \times 0.5$) per year. For a municipal tax rate corresponding to the average rate, i.e., 12.57%, the owner will pay 527.94€. Assuming that the actual rental value is 151.5% higher (in line with the Ministry of Finance study referenced above), it would be sufficient to levy a rate of 5% to obtain the same revenue.

3 Methods

3.1 Standard Model of Horizontal Competition

In the standard horizontal tax competition model, a jurisdiction simultaneously competes over a mobile tax base. If municipalities are not atomistic or if a large share of the system of jurisdictions changes tax rates, governments react strategically in response to competitor jurisdiction tax rates. A positively sloped reaction function indicates that the own-municipality tax rate is a strategic complement with the neighboring (competing) jurisdiction tax rate, while a negatively sloped reaction function indicates that the tax rates are strategic substitutes.⁹ The larger (in absolute value) the magnitude of the reaction

⁸For example, in the United States, depending on the precise inputs to the Poterba (1992) formula, converting residential property tax rates based on values to the effective tax rate on rents may yield tax rates that are seven times larger.

⁹Strategic substitutes and complements are defined by Bulow, Geanakoplos and Klemperer (1985) as whether a more “aggressive” (lower tax rate) strategy by town i lowers (substitutes) or raises (complements) town j ’s marginal payoff from the tax competition game (what the government maximizes—revenues, welfare, property values, etc.). Loosely speaking this means whether the response to a more aggressive policy is to be more or less aggressive. The prior literature indicates the importance of this is whether tax rates converge or diverge in the long run.

function slope, the more intense the strategic response to the average of those competitor rates.¹⁰ The tax reaction function in a tax competition model has the same reduced form as with yardstick competition or fiscal spillovers, so the existence of strategic interactions per se cannot uniquely identify these models.¹¹

The usual way of capturing strategic interactions of tax policies across municipalities is to use a spatial lag model. Let i index the N municipalities in the sample and t index the T time periods. The standard approach (Brueckner 2003) to determining the existence of strategic interactions is to estimate:

$$\mathbf{y} = \rho \mathbf{W}\mathbf{y} + \mathbf{X}\beta + \mathbf{X}^E\gamma + \Gamma + \Theta + \Phi + \varepsilon \quad (1)$$

where \mathbf{y} is a $NT \times 1$ vector of tax rates, \mathbf{W} is a block-diagonal spatial weights matrix of dimensionality $NT \times NT$, Γ are time fixed effects, Θ are jurisdiction fixed effects, Φ are EIMC tax regime fixed effects, \mathbf{X} includes exogenous municipal control variables and \mathbf{X}^E includes exogenous EIMC control variables. We leave the error term in a non-parametric form given by $\varepsilon = \mathbf{R}\epsilon$ where \mathbf{R} is an unknown $NT \times NT$ matrix and ϵ is a $NT \times 1$ vector with mean zero.¹² We can interpret $\mathbf{W}\mathbf{y}$ as the spatial lag of tax rates of competitor jurisdictions. The weight matrix requires the researcher to specify an exogenous set of weights w_{ij} between municipality i (rows) and municipality j (columns) where w_{ij} demonstrates the relative connectivity between the two jurisdictions. A key to identification here is that the weights are exogenously given by the researcher.

We include a variety of control variables in \mathbf{X} and \mathbf{X}^E at the municipal and EIMC levels. We discuss the precise controls in the data section of the paper. In addition to time-varying municipality and EIMC level controls, to account for fixed EIMC charac-

¹⁰For surveys of the empirical literature, see Brueckner (2003), Revelli (2006), and Revelli (2015). Classic spatial lag studies include Brueckner and Saavedra (2001), Parchet (2019), and Lyytikäinen (2012), among many others.

¹¹Agrawal, Hoyt and Wilson (2022) define tax competition as “process by which governments set policy in the presence of competitive pressures caused by policy-induced mobility”, while yardstick competition is the process by which “voters use information about the taxes and spending levels in other jurisdictions ... to measure the competence of their own incumbent politicians.” Fiscal spillovers, studied in Case, Hines and Rosen (1993), are the public good benefits that accrue to nonresidents when a local jurisdiction enacts a policy.

¹²The spatial error term enters here in non-parametric form as advocated by Kelejian and Piras (2017).

teristics for each tax regime, we include EIMC tax regime fixed effects where the tax regimes are given by the various possible modes of cooperation in Table 1. These fixed effects account for level shifts in local tax rates when municipalities change tax regimes, possibly due to designating more or less authority to the EIMC. The fixed effects allow identification of our model to come from variation within a tax regime. However, one may be concerned that these regimes are a choice variable in the tax competition game for the smallest EIMCs, that are “communautés de communes.” For larger EIMCs, the tax regime is not endogenous because the tax regime is automatically imposed on the EIMC (Breuillé and Duran-Vigneron 2024). Thus, given the relatively small number of communautés de communes, we believe the tax regimes can be treated as exogenous. However, for “communautés de communes, similar to some of the literature on vertical tax competition, our assumption is that the tax regime is chosen by the EIMC and that no single municipality is large enough to influence this decision. Nonetheless, future research might explore the decision of which tax regime is selected by these smaller EIMCs, possibly finding exogenous predictors for the choice of regime.

We define the spatial weight matrix as binary but with a distance restriction. Define the distance from municipality i to municipality j as d_{ij} kilometers. Municipalities sufficiently far away (greater than \bar{d} kilometers) from municipality i are assumed not to be competitors and are too far away to feasibly enter into a cooperative agreement. Then one possible spatial weights matrix is a binary spatial weights matrix where $w_{ij} = 1$ if $d_{ij} < \bar{d}$ and 0 otherwise. The binary form of the weights will facilitate prediction of those weights using exogenous variation.

Although some studies have found that proximity is not a good measure of competitors at the *state* or *international* level, due to trade linkages, a distance-based threshold to determine competitors remains theoretically and empirically justified for *local* tax competition for several reasons. First, Janeba and Osterloh (2013) provide survey evidence that political leaders of municipalities, which on average are small, view nearby jurisdictions as their most likely competitors for mobile tax bases. Only the *twenty* very largest jurisdictions based on population view jurisdictions in another state or country as their

competitors. Even for these larger jurisdictions, proximity is most important. Second, local governments are likely linked by commuting patterns which, unlike trade patterns, are more geographically linked (Parchet 2019). Third, if yardstick competition is at work, for small municipalities that represent the bulk of our sample, voters likely have information on public services in nearby places and make comparisons with them. Finally, Agrawal (2015) shows that the spatial reach of local tax competition is less than an hour.

For our baseline specification, we use a threshold $\bar{d} = 45$ kilometers, as more than 99% of municipalities cooperated with neighbors located at a distance lower than 45 kilometers. This threshold is sufficiently large to capture cooperation but also sufficiently large enough to capture competition with non-members of the EIMC. We show the results are robust to using smaller thresholds, as well as adding a third weight matrix to capture competition beyond 45 kilometers.

Because jurisdictions set tax rates in a simultaneous game, they are endogenous. The researcher then estimates (1) using maximum likelihood or instrumental variables to deal with the simultaneity of tax rates. Then ρ gives the magnitude of strategic interactions. Note that in the standard model of capital tax competition with atomistic (price-taking) jurisdictions (Zodrow and Mieszkowski 1986; Wilson 1986) facing world capital markets, jurisdictions compete à la perfect competition. In this way, a reaction function with $\rho = 0$ may still indicate that policy interdependence and competitive forces are at work, but rules out *strategic* interactions.

3.2 Horizontal Competition with Cooperation

Now suppose that the researcher wishes to decompose tax competition among different sets of jurisdictions. One possible example of this might involve trying to determine if a jurisdiction competes more intensely with other nearby jurisdiction in its own state versus nearby jurisdictions in a neighboring state. Cassette, DiPorto and Foremny (2012) is one example; this paper identifies tax competition within France separately from tax competition with municipalities on the other side of the French-German border. Gérard, Jayet and Paty (2010) disentangles competition across different regions in Belgium. In

both examples, the partition is exogenous.

Continue to suppose that municipalities potentially can compete with municipalities within the exogenously given threshold of \bar{d} kilometers, but suppose that this set of nearby municipalities is partitioned into two (endogenously) determined sets: E and O. In our case, this is municipalities within the EIMC (**E**) and municipalities outside the EIMC (**O**). This partition is endogenous because it is a choice of the municipality and likely a strategic variable in the tax competition game. We partition the matrix \mathbf{W} , that has not yet been row-normalized, into the two sub-matrices for E and O. Under this partitioning, the researcher can estimate a spatial lag model with two separate weight matrices corresponding to the two groups, of the form:

$$\mathbf{y} = \rho^E \mathbf{W}^E \mathbf{y} + \rho^O \mathbf{W}^O \mathbf{y} + \mathbf{X}\beta + \mathbf{X}^E \gamma + \Gamma + \Theta + \Phi + \varepsilon \quad (2)$$

where \mathbf{W}^E is the weight matrix for the “E” set of municipalities \mathbf{W}^O is the weight matrix for the “O” set of municipalities.

The partitioning of the \mathbf{W}^E and \mathbf{W}^O matrix can easily be understood visually. Figure 1 shows the spatial weight matrices for two towns. Because the \bar{d} radius is drawn specifically for each municipality, the set of nearby jurisdictions that are not in the EIMC is different for different towns within the same EIMC. In this figure and in our baseline specification, this represents a distance of 45 kilometers. This threshold serves an additional role by allowing us to differentiate between “structural” zeros and “strategic” zeros indicating the municipalities sufficiently close that the municipality is not choosing to cooperate with. In turn, this will allow us to predict for ij pairs within \bar{d} kilometers using an exogenous source of variation.

To proceed, we make the following assumptions on the weights. Specifically, we partition the binary spatial weights matrix as $\mathbf{W} = \mathbf{W}^E + \mathbf{W}^O$. For ease of notation, let the weights in the first partition be given by $w_{ij}^E = 1$ if $d_{ij} < \bar{d}$ and i and j are in the same EIMC; $w_{ij}^E = 0$ otherwise. Given EIMCs are small, equal weight is likely similar to any other distance or characteristic-based weight. Of course, $w_{ii}^E = 0$. If a municipality does not elect to cooperate, all elements of its row in this matrix will take on a value

of zero. This then implies that $w_{ij}^O = 1$ if $d_{ij} < \bar{d}$ and i and j are *not* in the same EIMC; $w_{ij}^O = 0$ otherwise. Equal weight only among nearby places would be consistent with fixed residential locations and local labor market mobility (Wildasin 2014) and survey evidence indicating that municipalities often equally view nearby places as their competitors (Janeba and Osterloh 2013). Sensitivity to \bar{d} suggests the results are robust to the choice of the cut-off.

Each of these two partitioned matrices will then be row-normalized. Row-normalization means that each weight in a given row (town) is divided by the sum of all weights in that given row, allowing for a convenient interpretation of the spatial lag of tax rates as an average tax rate. An advantage of row-normalizing after partitioning is that the spatial coefficients are bounded, but the main reason to row-normalize after partitioning relates to our prediction equation below. Note we can make a test of equality of the coefficients on the spatial lags, however, if the null is not rejected, it does not mean that the model with one global \mathbf{W} is appropriate because the models are not nested.¹³ This specification allows the researcher to conduct a statistical test of whether $\rho^E = \rho^O$, under our row normalization.

After row normalizing, denote the row normalized matrices as $\widetilde{\mathbf{W}}^E$ and $\widetilde{\mathbf{W}}^O$. Let n_i^E denote the number of municipalities (other than yourself) in the EIMC and n_i^O as the number of competitor municipalities that are not a member of the EIMC. Then, the individual row-normalized weights are given by $\widetilde{w}_{ij}^E = 1/n_i^E$ if $w_{ij}^E = 1$ and zero otherwise.¹⁴ Similarly $\widetilde{w}_{ij}^O = 1/n_i^O$ if $w_{ij}^O = 1$ and zero otherwise. We can thus no longer rewrite row-normalized matrices as a linear combination.

Thus, as our preferred specification we estimate

$$\mathbf{y} = \rho^E \widetilde{\mathbf{W}}^E \mathbf{y} + \rho^O \widetilde{\mathbf{W}}^O \mathbf{y} + \mathbf{X}\beta + \mathbf{X}^E\gamma + \mathbf{\Gamma} + \mathbf{\Theta} + \mathbf{\Phi} + \boldsymbol{\varepsilon}, \quad (3)$$

¹³An alternative approach would be to start with a global \mathbf{W} matrix that is first row-normalized and then partitioned. An advantage of this approach is that the model with the partitioned weight matrix directly nests the model with one \mathbf{W} in case of equality of the spatial autocorrelation coefficients, so that the test of equality of spatial coefficients has a straightforward interpretation in terms of choice of specification. In this case, the coefficients with different weight matrices are not bounded. And as discussed below, more importantly, predicting the weights will become challenging.

¹⁴If a town is not a member of an EIMC, $n_i^E = 0$. Then, all elements of that row are zero.

which is a higher order spatial lag model (Elhorst, Lacombe and Piras 2012; Gupta and Robinson 2015; Lee and Liu 2010; Badinger and Egger 2010). Theoretical papers on higher order spatial lag models row normalize after partitioning (Badinger and Egger 2010; Kang and Li 2022; Li and Kang 2021).

3.3 Resolving Endogeneity of the Weight Matrix and Taxes

Given that a jurisdiction determines with whom to cooperate, the elements of \mathbf{W}^E and therefore $\mathbf{W}^O = \mathbf{W} - \mathbf{W}^E$ are endogenous even though the global weight matrix \mathbf{W} is exogenous given it is based on an exogenous distance cutoff. We might think of this endogeneity problem arising from decisions to cooperate being correlated with unobserved characteristics of the municipalities that also affect their outcome in terms of tax setting, e.g., richer municipalities have often entered later into intermunicipal cooperation or municipalities with high public services may combine with unobservable similar jurisdictions. Moreover, the decision with whom to cooperate may be simultaneously determined with tax rates as part of the tax competition game.

In order to deal with this issue, our estimation approach adapts that of Kelejian and Piras (2014) to a second order spatial autoregressive model. An alternative approach to this is proposed in Qu and Lee (2015). However, their approach implies that the set of the elements of the weights matrix are modeled as a function of a bounded function, which is appropriate for continuous weights but not for binary weights as in our problem. Other approaches use a Bayesian rather than a frequentist approach and/or base their strategy with a full parametric and microfounded specification of the formation of the network (Han 2016, Hsieh and Lee 2016, Hsieh, Lee and Boucher 2020). We keep the formation of the network unspecified and therefore rely on a reduced form estimated using exogenous variation, for which we perform extensive checks of validity and robustness.

The method of Kelejian and Piras (2014) consists of first tackling directly the endogenous nature of \mathbf{W}^E by projecting the spatial weights using an exogenous predictor variable and then using the spatial lags of the control variables computed with the projected matrix as instruments. Denoting the subset of elements of our endogenous matrix

\mathbf{W}^E that are less than \bar{d} kilometers and are off-diagonal elements in panel year t as $\bar{w}_{ij,t}^E$, the reduced form¹⁵ used to predict the weights is:

$$\bar{w}_{ij,t}^E = \beta_0 + \beta_1 s_{ij,t-39} + e_{ij}, \quad (4)$$

where s_{ij} is a variable used for predicting the weights. Endogeneity arises from e_{ij} being correlated with ε_i , the error term of the outcome equation in a given year t . The variable s_{ij} needs to be strongly correlated with current day EIMC membership, but uncorrelated with current day tax policy, i.e., $cov(s_{ij}, \varepsilon_i) = 0, \forall i, j$.¹⁶ In other words, this variable s_{ij} should impact current taxation policies only through the correlation with \mathbf{W}^E and not directly. In step one, we estimate this equation separately for each time period in our panel, which runs from 1994 to 2013.¹⁷ We follow the paper by Kelejian and Piras (2014) who derive this prediction stage without controls and instead only use the exogenous predictor of the weights.

Our baseline specification of s_{ij} , relates to historical syndicate connecting municipal pairs i and j . Specifically, we use *stocks* of syndicates formed 39 year prior to year t or earlier. In other words, to predict the weights in each year $t = 1994, \dots, 2013$, $s_{ij,t-39} = 1$ if municipalities i and j jointly were a part of the same syndicate at any time between 1900 to year $t - 39$ and 0 otherwise.¹⁸ The choice of 39 years comes from the fact that a major reform in 1955 expanded syndicate powers resulting in a shock that led to their (arguably exogenous) proliferation in the years following. For the 1994 year in our panel, we thus predict the 1994 EIMC membership with whether the jurisdictions created a common syndicate in 1955 or earlier. We predict the 1995 EIMC membership with whether the

¹⁵We use the terminology reduced form rather than first stage, because predicting the spatial weights is simply an input to constructing the instruments to the first stage. In this stage, we are not instrumenting for the endogenous weights, but rather predicting them in order to construct exogenous spatial lags as instruments that can be used in a first stage.

¹⁶Note that Kelejian and Piras (2014) do not have a structural model of link formation.

¹⁷Note that the repeated cross section approach provides the same estimates as a panel where the coefficient is interacted with a year dummy. We prefer to keep some flexibility in this first equation, rather than constrain the parameter to be constant over time, especially since we know the extent of cooperation changes over our sample period.

¹⁸If a syndicate were bilateral, this would equal one for only those two municipalities; when syndicates contain multiple municipalities, the variable is unity for all pairs of municipalities that cooperated together.

jurisdictions created a common syndicate in 1956 or earlier, and so on.

The period 1955-1974 spanning the historical syndicate membership covers important historical laws (1959, 1966, 1970, 1971) related to syndicate expansions; these laws which stimulated cooperation provide ample exogenous shocks to the proliferation of syndicates. This increase in syndicate membership helps provide some time variation in our predicted spatial weights, which helps with identification given we know EIMCs increased over the course of our estimating sample. Moreover, by exploiting syndicate creation in year $t - 39$ and before, we exploit the idiosyncratic variation in *when* (in addition to *whether*) jurisdictions began to cooperate via syndicates in the post-WWII period.¹⁹ We have verified the results are robust to starting at different years or only considering creation flows rather than the stock of places that historically cooperated together.

Syndicates are much weaker forms of cooperation that do not contain taxation authority. They are the first forms of inter-municipal cooperation, before the implementation of laws in the 1990's boosted cooperation through taxation power in EIMCs. Given the historical nature of these syndicate creations, we view syndicate membership as a signal of their desire to cooperate. If multiple municipalities were likely to cooperate on these limited public services, they are more likely to cooperate to this day, but historical creation of these syndicates is not related to EIMC membership and taxation today.

As a robustness check, we utilize a second variable, the logarithm of commuting flows in 1990, denoted c_{ij} to predict the weights. This variable is not time varying and obtaining earlier years of data is not possible. In particular, we observe the number of workers who live in municipality i and commute to work in municipality j . One may worry that the 1990 commuting flows are not far enough in the past and for this reason we only use it to verify the sensitivity of our baseline results.

We obtain our predicted weights matrix $\widehat{\mathbf{W}}^{\mathbf{E}}$ using the fitted values from the equation for cooperation, (4), and we then compute the other predicted matrix: $\widehat{\mathbf{W}}^{\mathbf{O}} = \mathbf{W} - \widehat{\mathbf{W}}^{\mathbf{E}}$. While weights matrices $\mathbf{W}^{\mathbf{E}}$ and $\mathbf{W}^{\mathbf{O}}$ are only made of elements 0 and 1, the predicted weights matrices $\widehat{\mathbf{W}}^{\mathbf{E}}$ and $\widehat{\mathbf{W}}^{\mathbf{O}}$ are not. After obtaining these predicted

¹⁹We observe all syndicates created since the 1900s and the initial membership in that syndicate, but we do not know if members left the syndicate or if the syndicate was dissolved.

matrices, we row-normalize both matrices. The need to compute of $\widehat{\mathbf{W}^{\mathbf{O}}} = \mathbf{W} - \widehat{\mathbf{W}^{\mathbf{E}}}$ necessitates we row normalize after partitioning the weight matrices rather than before. If instead, we row-normalized \mathbf{W} and then estimated $\widehat{\mathbf{W}^{\mathbf{E}}}$, the projected matrix would not necessarily on the same scale as a row-normalized \mathbf{W} .

Having resolved the endogeneity of the spatial weights matrices, we proceed by adapting the procedure set out by Kelejian and Piras (2014) to resolve the endogeneity of the spatial lag terms. Given that tax rates are selected simultaneously in a strategic game, the spatial lags of tax rates are by definition endogenous as tax changes in each jurisdiction affect taxes in other jurisdictions. Kelejian and Piras (2014) provide a procedure for a first-order spatial lag model. In what follows, we adapt their instruments to accommodate the fact that our model is a high-order model with several spatial lags. Formally, we use as instruments the set of exogenous variables \mathbf{X} , the spatial lags of the exogenous variables with respect to both matrices and with respect to the product of the two, namely we construct the instrument matrix as, $\mathbf{H} = [\mathbf{X}, \widehat{\mathbf{W}^{\mathbf{E}}}\mathbf{X}, \widehat{\mathbf{W}^{\mathbf{O}}}\mathbf{X}, \widehat{\mathbf{W}^{\mathbf{E}}}\widehat{\mathbf{W}^{\mathbf{O}}}\mathbf{X}, \widehat{\mathbf{W}^{\mathbf{O}}}\widehat{\mathbf{W}^{\mathbf{E}}}\mathbf{X}]$. The precise variables used are defined subsequently in section 4.

In order to justify the use of spatial lags—based on our predicted weight matrices—of \mathbf{X} as valid instruments, we can write down a structural model that makes clear the assumptions necessary (following Revelli 2005). Let the tax base in jurisdiction i be denoted $b_i = b(y_i, \mathbf{y}_{-i}, \mathbf{X}_i)$ where y_i is the tax rate in i , \mathbf{y}_{-i} is the vector of tax rates in other jurisdictions, and \mathbf{X}_i are the characteristics of jurisdiction i that influence the size of the tax base. Governments maximize an objective function $w_i = w(y_i, b_i, \mathbf{X}_i)$ that depends on its tax rate, tax base, and characteristics. Maximizing w_i then yields the policy reaction function $y_i = y(\mathbf{y}_{-i}, \mathbf{X}_i)$. This policy function makes it clear that under the assumption that neighboring covariates do not enter in the tax base or welfare function, then tax rates in a given jurisdiction only respond to *its* characteristics. Thus, neighboring jurisdiction tax rates also only depend on their the neighboring jurisdiction's characteristics. We follow the standard practice of assuming that *neighboring* characteristics and preferences do not enter the structural equation for the tax base and welfare function. Thus, under this assumption, they provide valid shocks to neighboring tax rates, but do not determine

other jurisdiction's tax rates directly. Then, the spatial lags of the \mathbf{X} valid instruments so long as the weight matrices are based on our prediction using syndicate membership. Of course, the exclusion restriction relies on this assumption.

We proceed with these standard instruments as our main methodological contribution is to resolve the issues relating to endogenous network formation. In particular, as shown below, some critiques of these instruments would, under the most realistic assumptions, result in estimating the *exact opposite* spatial patterns that we find below.

We cluster the standard errors at the department level (the administrative and political division that is between the regional level and the (inter-)municipal level) allowing for arbitrary correlation within the 96 departments, allowing for possible spatial correlation of errors within departments along with correlation over time. According to the Eurostat nomenclature for territorial divisions, the department is classified as a NUTS3 administrative unit.

3.4 Specification Consideration: Multiple Tax Rates

The French system allows municipalities to set four main tax rates and tax competition could potentially occur over all four of these instruments. One approach would be to estimate four separate models for each tax rate. However, this approach would require assuming that there is no correlation between tax rates. To see this, suppose that we estimated equation (1) for the business tax and the built property tax separately. Then, the available instruments for the spatial lag of the business tax would be $[\mathbf{X}, \widetilde{\mathbf{W}}\mathbf{X}]$ where $\widetilde{\mathbf{W}}$ denotes the row normalized matrix. The available instruments for the spatial lag of the built property tax rate would also be $[\mathbf{X}, \widetilde{\mathbf{W}}\mathbf{X}]$. Note that if the instruments strongly predict the spatial lag of the business tax rate and the spatial lag of the built property tax rate, they can only be valid instruments for both specifications if the spatial lag of the built property tax rate has no effect on the business tax rate. We believe this assumption is far too strong given that cross-tax base effects have been shown to be important in the literature. As an alternative, we could control for the spatial lags of the full vector of policy instruments; however, this approach would require additional instrumental variables.

To proceed, we construct a tax index and argue that municipalities compete not just on the level of taxes but on an effective tax rate. We think of such an approach as being a “tax system” approach (Slemrod and Gillitzer 2014) where the researcher tries to consider the tax system as a whole. The index simplifies a multidimensional system to a single strategic variable. This is especially justified in France where some rules limit the freedom with which municipalities can change the tax rates. Particularly relevant, municipalities cannot change the tax rates on the four local taxes independently (increasing one tax rate has an impact on the other tax rates through complex linkages). We construct the tax index, which in our preferred specification is a weighted average—with weights ν^l on tax rate $l = B, R, DP, UP$ —of the four main taxes municipalities have access to, $(\nu^{DP}\tau^{DP} + \nu^{UP}\tau^{UP} + \nu^B\tau^B + \nu^R\tau^R)$. In particular, we utilize only the tax rate set by the municipality and not the EIMC; if a tax rate is surrendered entirely to the EIMC—as in the case of the business tax under the mixed taxation or single taxation regime—it is zero in our index. Recall that all models include EIMC regime fixed effects to account for any level shifts arising from the surrender of a tax rate to the cooperative. Our focus is on municipal tax competition and thus we exclude all EIMC tax rates. While it might be tempting to study competition in the total tax rate (the sum of the town and EIMC tax indexes), presumably, the EIMC tax rate is chosen in some sort of cooperative fashion by its component towns. As a result, it is not clear that looking for town-level interactions in the choice of the total rate is theoretically grounded, given that a component of the total index is chosen cooperatively. In some specifications, we will explore municipal competition with EIMCs that may arise via vertical tax competition, but leave the study of competition among EIMCs to future work.

To construct the weighted average, for each tax rate $l = B, R, DP, UP$, we weight by the fraction ν^l of (aggregated at the national level) municipal tax revenue from each tax instrument using the first (1994) or last (2009) year of our sample. We hold the weights constant across municipalities and years so that any changes are driven by actual changes in tax rates. In robustness checks, we verify the sensitivity of the empirical analysis to various weights: time-varying weights where ν^l are calculated each year, fixed weights

based on municipal and EIMC revenue combined, and an unweighted index that is the simple average of the taxes, $(\tau^{DP} + \tau^{UP} + \tau^B + \tau^R)/4$. The latter of these is not preferred because it gives disproportionate weight to the least important taxes.

The construction of the tax index deals with a well-known theoretical problem, but previously unresolved empirical concern, that is, that governments set multiple tax rates simultaneously. The prior empirical literature often focuses on a single tax instrument when studying tax competition, which implicitly assumes that the cross-base reaction function has zero slope. Such an assumption is extremely strong, so researchers often rely on exogenous controls as proxies for other tax changes. Needless to say, such an approach is unsatisfying. Our index approach proposes a novel solution to this issue by arguing that governments compete over multiple tax rates, for which we can identify a (weighted) average effect. And it is reasonable to consider an effective tax rate as all four of the taxes are on different types of capital and land.

3.5 A Note on a Spatial Approach vs. Difference-in-differences

We have elected to follow the conventional spatial econometrics approach rather than implement a difference-in-differences (DiD) design. Our paper requires thinking about formation of an EIMC and then tax changes—both of which are (endogenous) events—not just a single program evaluation event study. We wish to isolate the covariance of own-jurisdiction tax rates with competitor tax rates, with an emphasis on the heterogeneity between competitors inside the EIMC and competitors outside of the EIMC. Then thinking about formulating an event study design, there are two possible “events” we could consider: (1) the formation of the EIMC and (2) municipal tax rate changes. Critically, a staggered DiD design that uses EIMC formation as an event would find the effect of joining an EIMC on the level of tax rates. Such a specification captures the effect of EIMC membership on tax rates via many mechanisms other than competition including economies of scale in public services. There is also no apparent way to isolate the effects of the event into within EIMC effects and external effects, absent constructing spatial lags (as we do) and using them as dependent variables in a DiD analysis. Further, such a design would ignore the fact that any EIMC formations are not exogenous or random, but

rather—as is the point of our paper—occur endogenously in the tax competition game.

Given our paper is interested in the strategic responses as taxes change, we could alternatively define competitor tax changes as the event. One way to do this would be to separately define an event as a tax change of municipalities within the EIMC and separately define another event as tax changes of municipalities nearby the EIMC (both defined based on changes in spatial lags of tax rates). One could then conduct a staggered DiD using these events, with the own-jurisdiction municipal tax rate as the outcome of interest. One could then see if the DiD coefficient for the peer EIMC member events is different from that of the neighbors outside the EIMC. However, defining an event study this way has many practical challenges. First, defining events based on spatial lags again ignores the endogeneity of those events. Second, the spatial tax index is an average of four tax rates. In addition, it is a spatial average across many towns. As a result, the spatial lag changes very frequently and sometimes by small amounts and in close succession. Thus, there is no “clean” way to define an event. While it is possible to focus on only large tax changes, this raises selection concerns and the fact that taxes change frequently makes it unlikely there will be many large clean events necessary to verify pre-trends. Finally, estimating one event study without controlling for the other event would lead to an omitted variable bias, which would necessitate entering the other spatial lag as a covariate—and requiring us to use an IV procedure.

4 Data

We use a similar panel dataset as Breuillé, Duran-Vigneron and Samson (2018) on the four main municipal and inter-municipal tax rates in France, but we augment it with comprehensive data on municipal cooperation (both current day and historical). These data are collected by the “Direction Générale des Finances Publiques” (Ministry of Finance) and cover the period 1994-2013, except for the business tax, which covers the period 1994-2009. Indeed, the business tax was replaced in 2010 by the “Contribution Economique Territoriale” (CET), which comprises a tax on the rental value of proper-

ties and a tax on value added.²⁰ For the residence tax, the property tax on developed land and the property tax on undeveloped land, we use the statutory tax rate. For the business tax, we calculate an average tax rate as a special regime called “régime fiscal de la taxe professionnelle de zone” coexists with the standard business tax rate. EIMCs having created an area called “zone d’activités économiques” have the possibility to levy an additional business tax rate for this area.

In addition to this tax data, the originality in our dataset comes from the collection of substantial historical data on syndicate membership from 1900 to today from the Ministry of the Interior, which has never been exploited before. These data tell us which municipalities formed a syndicate with each other and in which year. These data do not include later departures from or abolishing of the syndicates. In addition, the territorial changes of both municipalities (mergers, creation) and EIMCs (creation, enlargement, mergers), as well as the adoption of new legal status by some EIMCs required us to harmonize the database. After this comprehensive data assembly we construct one of the most complete panel datasets concerning municipal public finance cooperation in France.

Time-varying control variables at the municipal level are limited. At the municipal level yearly frequency, we include in \mathbf{X} several variables: percent of single-parent families, percent of executives (for the population between 25 to 54 years), population, number of jobs held by people between 25 and 54 years in the municipality, and the unemployment rate for the same segment of the population. Some enter in spline form to allow for non-linear effects of characteristics such as population. We also have access to political variables, but we only have access to these in election years, so we set these variables to remain at their prior level in all non-election years. We also control, in \mathbf{X}^E , for several characteristics at the EIMC level to adjust for any EIMC level factors that may affect municipal tax rates. These include the number of towns in the EIMC, the number of years in the EIMC, a dummy for whether the EIMC was created prior to our sample, and dummy variables for the EIMC type (“communautés d’agglomération”, “communautés de communes”, “communautés urbaines”, and other older statuses such as “communautés de

²⁰Given this reform was major, we focus on the period until 2009. For an analysis of the business tax reform, see Ly and Paty (2020).

ville” and “districts”, see Appendix A.1).²¹ In addition, to capture EIMC demographic and political factors, we aggregate municipal controls to the EIMC level. The set of controls also includes EIMC tax regime fixed effects designed to capture any level effects of the choice of regime (Table 1) following formation of an EIMC.

Before proceeding, we describe some of the descriptive statistics for the variables in our regression. Given our interest is in studying tax competition, we wish to highlight the amount of variation we have in taxes. First, Figure 2 visually shows the number of tax changes to our tax index along with the growth rate of the tax index.²² As can be seen, the average municipality changes at least one of its tax rates several times over the course of our sample, resulting in significant changes in the municipality’s tax index. In particular, the average municipality experiences more than seven tax changes over the course of our sample. This suggests that we have ample variation to identify tax competition.

5 Results

We present a series of results before presenting our preferred specification. First, we present OLS results where we do not resolve any endogeneity concerns. Then, we present results currently common in the literature where we treat neighboring jurisdiction tax rates as endogenous (and appropriately IV using spatial lags of the exogenous covariates) but where the spatial weight matrix is treated as exogenous. Finally, we present our preferred results where we resolve both endogeneity of the spatial weights matrix and the tax rates. The successive presentation of results in this manner allows us to discuss the various sources of bias at each stage and to highlight how accounting for the endogeneity of cooperation is critical. Further robustness checks are presented later.

5.1 Baseline Results

The first column of Table 3 shows the estimate of ρ from a row-normalized variant of (1), ignoring all endogeneity concerns. The coefficient on $\widetilde{\mathbf{W}}\mathbf{y}$ is the slope on the strategic re-

²¹The current EIMC types are often determined by population and place certain restrictions on the EIMC, in some cases relating to the tax regime and the competencies that can be selected.

²²One noticeable feature is that our tax index variable is falling for many municipalities. These trends in tax rates are given in figure A.3.

action function regardless of EIMC membership. This coefficient masks any heterogeneity in tax competition across municipalities within the EIMC and outside of the EIMC. A one percentage point increase in the tax index of competitor municipalities significantly raises the own jurisdiction’s tax index by less than one half of a percentage point. Like much of the prior literature (for example, Brueckner 2003), tax rates at the municipal level are strategic complements. The results are similar regardless of whether the tax index is constructed using the weights $\nu^B, \nu^R, \nu^{DP}, \nu^{UP}$ in the first year of our sample (“1994 Index”) or the last year of our sample (“2009 Index”).

In the second column of the table, we partition the weight matrix into the E and O components in (3). Again, in this specification, we ignore all endogeneity concerns. The slope of the reaction function to other municipal tax rates in the same EIMC is slightly smaller than with respect to tax rates set by towns outside of the EIMC. We marginally cannot reject that the coefficients ρ^E and ρ^O are the same. However, these results should be interpreted with caution: the spatial lags of the tax rates are endogenous both because of simultaneous setting of tax rates and endogenous weights from the decision with whom to cooperate.

One might be tempted to compare the coefficient on the single weight matrix $\widetilde{\mathbf{W}}\mathbf{y}$ with the coefficients on the partitioned weight matrices $\widetilde{\mathbf{W}}^E$ and $\widetilde{\mathbf{W}}^O$. However, row-normalization introduces a non-linear transformation, so there is no direct relationship between the partitioned and the single weight matrix. While there are papers that propose other standardizations of the weight matrix—see for instance Kelejian and Prucha (2010) who propose a normalization by the minimum of the maximal values of the row sums and column sums—these alternatives involve non-linear transformations.

Next, we instrument for the spatial lags assuming the weight matrices are exogenous. Following the standard approach (Gibbons, Overman and Patacchini 2015), table 4 presents the results. In the case of exogenous weights, we use instrument set, $\mathbf{H}^* = [\mathbf{X}, \widetilde{\mathbf{W}}^E\mathbf{X}, \widetilde{\mathbf{W}}^O\mathbf{X}, \widetilde{\mathbf{W}}^E\widetilde{\mathbf{W}}^O\mathbf{X}, \widetilde{\mathbf{W}}^O\widetilde{\mathbf{W}}^E\mathbf{X}]$, which relies on the original $\widetilde{\mathbf{W}}^E$ and $\widetilde{\mathbf{W}}^O$ weight matrices and not the projected ones. Although such an IV procedure reduces any simultaneity bias, it ignores any potential endogeneity from the decision to cooperate.

Relative to the OLS estimates, this approach raises the coefficient on the spatial lag of taxes for municipalities outside of the EIMC, but substantially lowers the coefficient on the spatial lag of taxes for municipalities inside the EIMC, making it close to zero. In our setting, the partition of $\widetilde{\mathbf{W}}^{\mathbf{E}}$ and $\widetilde{\mathbf{W}}^{\mathbf{O}}$ is determined by municipal choices to cooperate, which is clearly endogenous and these results may reflect that. Thus, like the OLS results in the prior table, this table should be interpreted with caution. Nonetheless, they provide a useful benchmark to help us determine the bias of ignoring the endogeneity of spatial weights and whether or not that endogeneity is important.

5.2 Results Resolving the Endogeneity of the Weight Matrix

Now we turn to our preferred results using the Kelejian and Piras (2014) approach to resolve the endogeneity of the spatial weights matrix. Recall, in the first step, we use historical syndicate membership creation to predict current day EIMC cooperation. We do this for every year in the data set. In table 5 we present the estimation results for (4).

Given the weight matrix has a dimensionality of approximately $36,000 \times 36,000$, even focusing on towns within 45 kilometers yields a very large number of pairs (over 17 million in each year) we need to predict. Given the large number of observations, it is expected that the coefficients pertaining to the syndicate membership are (statistically) significant. For this reason, we turn to the signs and magnitudes of the coefficients to judge their relevance. All signs are as expected: historically forming a syndicate with another town increases the probability of being in the same (current-day) EIMC. In particular, for 2008, syndicate membership creation almost in 1969 and earlier increases the probability of being in the EIMC with that same town by 8 percentage points. Furthermore, the instrument coefficients increase as time goes on, which is consistent with the proliferation of EIMC membership and syndicates over our sample.

The prediction stage, (4), is a linear probability model, making it possible that the elements of our predicted weight matrix $\widehat{\mathbf{W}}^{\mathbf{E}}$ could lie outside the $[0, 1]$ interval. Given the coefficients in table 5 are between $[0, 1]$ and because the estimated intercepts are close to zero, all predicted weights lie on the $[0, 1]$ interval. The predicted weights within our

threshold \bar{d} range between 0.006 and 0.097 with a mean of 0.025.

With respect to the validity of the prediction variable, as discussed above, syndicate membership is a much weaker form of cooperation than EIMC because syndicates have no taxation authority. In addition, the duties and services that could be performed are quite limited, often confined to collection and elimination of household waste, water treatment/distribution and public transport. Although some syndicates persist today, recent reforms have favored their dissolution into large EIMCs. Thus, exogeneity rests on the belief that these decisions were made long into the past and do not have persistent effects on tax setting authority. Furthermore, although they both signal a desire to cooperate, the mapping into EIMC cooperation and syndicates is not one-to-one. Overall, the differing functions and taxing authority strengthen the case for the syndicates being an exogenous predictor of cooperation.

In order to provide empirical support for the validity of using syndicate membership, we show that two neighboring municipalities that were part of the same syndicate long ago are not closer in terms of economic characteristics today, compared to two neighboring municipalities that were not part of the same syndicate. For each town i , we identify all of the competitor jurisdictions j within \bar{d} kilometers of the municipality. We then regress the observable characteristics of municipalities j in 1994 on an indicator for whether municipalities i and j formed a syndicate, a constant, and municipality i fixed effects. Table A.1 shows the results. Only one variable indicates a statistical difference between municipalities in the same syndicate and municipalities not part of the same syndicate. Moreover, the only economically significant variable relates to the size of municipalities, indicating that municipalities are more likely to cooperate with smaller neighbors who benefit from economies of scale. We conclude that neighboring municipalities that historically formed a syndicate are no closer in terms of observables than neighboring municipalities that were not part of the same syndicate.

We next show the results of applying Kelejian and Piras (2014) and using the instrument set $\mathbf{H} = [\mathbf{X}, \widehat{\mathbf{W}^E \mathbf{X}}, \widehat{\mathbf{W}^O \mathbf{X}}, \widehat{\mathbf{W}^E \mathbf{W}^O \mathbf{X}}, \widehat{\mathbf{W}^O \mathbf{W}^E \mathbf{X}}]$ that relies on the projected matrices. Critically, the syndicates are not directly in the instrument set, except for their

use in constructing the projected matrices. Table 6 presents the results of the Kelejian and Piras (2014) procedure. This is our preferred specification because it resolves *both* endogeneity of the weight matrix and the simultaneity of tax rates. The F statistic for the strength of instruments is well above its critical value. Further, the p-value on the Hansen J statistic is high, suggesting that if one instrument is valid, other instruments are also valid instruments. The Hansen J statistic tests the overidentifying restrictions with respect to instruments, which are constructed using the fitted weight matrices. All specifications also satisfy the stability condition for multiple weight matrices (Elhorst, Lacombe and Piras 2012).

As discussed above, the structural model that justifies these instruments relies on the assumption that neighbors' covariates do not enter into the tax base or welfare function of the own-jurisdiction. In addition, in our setting, the weights used to construct the instruments must not be endogenous to the tax competition game—something achieved by predicting the weights with syndicate membership.

A one percentage point increase in the average tax index of other EIMC members, increases the own-jurisdiction tax index by 0.307 percentage points. On the other hand, a one percentage point increase in the average tax index of towns outside the EIMC, increases the own-jurisdiction tax index by 0.580 percentage points.²³ A formal test of the statistical difference in the coefficients rejects their equality at the one percent level. Thus, towns respond less aggressively to the spatial average of EIMC tax rates than they do to the spatial average of town tax rates outside the EIMC.

In terms of identification, all specifications include EIMC regime fixed effects to account for possible differentiated effects of the tax regimes listed in Table 1 (additional taxation, single taxation, mixed taxation, no cooperation). As shown in Table 6, the results are robust to the inclusion of EIMC demographic and political controls.

Critically, the results are also robust to how we construct the index and to the choice of the threshold \bar{d} . Comparing the prior results (column 2) to a threshold where \bar{d} is 20 kilometers (column 2 Table 7), raises the coefficients slightly but any difference across

²³This result is robust to starting the panel in 1999, the year the Chevènement law was passed, which was a turning point in the expansion of EIMCs.

these specifications is statistically insignificant. Although the choice of \bar{d} was grounded in the existing literature on the spatial reach of tax competition, it does not qualitatively affect the results or the interpretation of our coefficients. Table A.2 also shows a threshold of 26 kilometers, which adds an additional 50 towns to the set of towns outside the EIMC on average. Despite this increase in the number of towns, the coefficients again change only slightly. These sensitivity checks also validate our choice of spatial weights, suggesting the responses to the averages are relatively similar within these areas.

With respect to the validity of the predictor used to estimate (4), we predict cooperation using more recent, but still historical, commuting flows as an additional instrument. As can also be seen in table A.2, the results are very similar to the results in table 6. This shows that our results are robust even if we have a just identified model (only syndicates) or an overidentified model (syndicates and commuting flows). Our results are therefore not sensitive to the choice of predictors.

How do the results resolving the endogeneity of the spatial weights matrix (Table 6) differ from the ones that do not (Table 4)? Comparing the two tables indicates that not accounting for endogenous weights biases ρ^E downward but that ρ^O is minimally impacted. Intuitively, the (endogenous) decision of whether to cooperate and with whom to cooperate with has the largest bias on strategic interactions with one's peers, i.e., municipalities within the same EIMC. A key point of our analysis is to show that failing to account for the endogenous formation of cooperative units yields substantially different estimates of the extent of tax competition than using an econometric design that accounts for this endogeneity. Simply put, the decision to cooperate and with whom to cooperate is a simultaneous choice in the tax competition game. In our setting this choice biases the coefficient on tax rates for EIMC almost to zero, even when some strategic interactions persist. Critically, tax cooperation does not entirely eliminate strategic tax competition.

5.3 Limitations

Before proceeding, we wish to address certain critiques of the spatial lag model. Various papers (McMillen 2010; Gibbons and Overman 2012) have criticized the spatial lag

model and the use of the spatial lags of the explanatory variables as instruments for lacking quasi-experimental variation. Under this critique, the spatial lag model may be useful at identifying correlations, but not causal effects because using neighboring \mathbf{X} 's as instruments may not satisfy the exclusion restriction. We acknowledge this critique, which other scholars have recently attempted to address by exploiting institutional shocks that “force” municipalities to change their tax rates (Lyytikäinen 2012; Parchet 2019). While ideally, we would exploit such instruments, they are not available in our setting.

As with most instruments, there are reasons why the exclusion restriction may not hold. One possibility is that the true model is a spatial Durbin model where the neighboring jurisdiction \mathbf{X} 's enter directly into (3)—something that would be true if they entered into the tax base and welfare functions at the end of section 3.3. Another possibility is that the neighboring jurisdiction \mathbf{X} 's do not enter (3) directly, however, they are correlated with unobservable (omitted) own-jurisdiction variables that influence local tax rates. To fix ideas, suppose that there is an omitted variable $z_{i,t}$ that has a correlation with the outcome variable $y_{i,t}$ and with a correlation with an element of \mathbf{X} . Then if $cov(z_{i,t}, z_{j,t}) \neq 0$ for $j \neq i$ so that jurisdiction i and j have unobservable common shocks, the exclusion restriction will fail because the neighboring \mathbf{X} 's will have a direct effect on own-jurisdiction taxes via the correlation with the omitted variable.

If this is the case, unobservable variables in jurisdiction i that also influence neighboring jurisdiction j , this would bias us against finding statistically different effects with EIMC and non-EIMC municipalities. Or if the bias from this critique were large enough, it may bias us to find the opposite pattern we identify $\rho^E > \rho^O$. Intuitively, unobserved common shocks to municipalities i and j should be positively correlated: if one municipality experiences an economic shock or a cost shock to providing public services, nearby jurisdictions also experience a similar shock. This, in turn would raise ρ^E relative to the true strategic interaction. If further away municipalities are less likely to experience the same shocks, then such unobservables would induce a smaller bias to the reaction function slope for municipalities outside of the EIMC. Given that we find $\rho^E < \rho^O$, criticisms of the spatial lag model suggest that the true reaction function slope ρ^E would be even

more different from ρ^O than we estimate.

Thus, were the failure of the exclusion restriction a result of omitted variables, any statistical test of differences in these coefficients would be more likely to reject our null hypothesis than with our current instruments. As a result, if these critiques are valid, we conclude the difference $\rho^E - \rho^O$ is biased toward zero. We find a statistically significant difference $\rho^E - \rho^O$, so our results are likely a lower bound on the difference. And our setting is one where, as discussed in the introduction, the sign of the difference is theoretically ambiguous so finding a significant difference (even if it is too small) is interesting in its own right. Moreover, given our emphasis is less on the instruments in \mathbf{H} and more on the instruments for cooperation, we do not believe this critique threatens the important new insights of our model that highlight simply using a projected weight matrix versus the endogenous weight matrix yields very different coefficients.

5.4 Robustness to the Weight Matrices and Normalization

The precise form of the spatial weights is unknown to researchers. We have made two assumptions: (1) all jurisdictions are equally weighted and (2) municipalities do not compete with municipalities further than 45 kilometers. We focus our robustness checks on the latter of these assumptions, as predicting non-binary spatial weights with syndicate membership would necessitate further modifications to our empirical design.²⁴ As discussed previously, Table 7 shows that our results are robust to changing the threshold \bar{d} . In other words, shrinking the number of towns outside of the EIMC yields similar reaction function slopes, indicative of similar competition among all towns in the market area. Finally, in Table A.3, by adding a third weight matrix, we show that the results are robust to controlling for tax changes of competitors that are further away. But, these towns further than 45 kilometers do not generate further strategic interactions, suggesting our threshold was appropriate for defining competitors.

The robustness to the distance threshold \bar{d} is critical because we have elected to first partition the weight matrix \mathbf{W} into its two sub-parts and then to separately row-normalize

²⁴Predicting continuous weights would likely require having a predictors that was not dichotomous, moving us away from syndicate membership.

both partitioned matrices. A concern is that our test of equality of ρ^E and ρ^O is sensitive to normalizing after—rather than before—partitioning the weight matrix. Suppose that each municipality j within 45 kilometers of i affects municipality i equally, but the number of municipalities in the same EIMC as i is smaller than the number outside the EIMC. Then, by row normalizing after partitioning, the estimate for ρ^E may be different than ρ^O even if true bilateral reaction functions would be identical. In practice, \mathbf{W}^E contains fewer non-zero elements than \mathbf{W}^O for our baseline value of \bar{d} (see Figure 3).

Given the test of our null hypothesis may be sensitive to the choice of when to normalize, we next select a value \bar{d} such that (on average) a municipality has the same number of non-zero observations in \mathbf{W}^E and \mathbf{W}^O . Then, row normalizing implies that on average each row in both matrices is divided by similar numbers. Using a threshold $\bar{d} = 7.25$ yields such a distribution, but comes with the cost that it may not include all members in the EIMC if the EIMC is large. Again, Table 7 indicates the results are similar in this case. This indicates that when individual jurisdiction have approximately similar weights in both matrices, competition remains less intense within the EIMC. Thus, our result do not appear to be driven by the fact that there are more non-zero elements in the “O” matrix rather than the “E” matrix.

5.5 Mechanisms via Vertical Competition

We find that strategic tax competition among municipalities in France is less intense for towns *within* the same cooperative unit than *across* cooperative units. Can theory shed light on why this result might arise? Appendix A.4 presents a theoretical model that rationalizes this finding. Linking our empirical results to theory requires some specification of the government objective function. If governments are Leviathan—revenue maximizers using the language of Brennan and Buchanan (1980)—our results imply that tax rates of neighboring municipalities within the EIMC have a smaller impact on marginal revenue than taxes outside of the EIMC, perhaps because of local linkages such as attachment to home or connections via commuting. If governments maximize a quasilinear welfare function, then under mild assumptions, our result is consistent with the provision of the

EIMC public services dampening tax competition within the EIMC. Given this, a plausible channel may arise via vertical fiscal competition (Besley and Rosen 1998; Keen 1998; Agrawal, Brueckner and Brülhart 2024).

In order to isolate this possible mechanism, we analyze the effect of EIMC tax rates. In Table 8, we use the municipal tax index as the dependent variable \mathbf{y} , but then study how it responds to $\widetilde{\mathbf{W}}^{\mathbf{E}}\mathbf{y}^{\text{EIMC}}$ and $\widetilde{\mathbf{W}}^{\mathbf{O}}\mathbf{y}^{\text{EIMC}}$ where \mathbf{y}^{EIMC} is the EIMC tax index. Thus, the coefficient on the own-EIMC tax index $\widetilde{\mathbf{W}}^{\mathbf{E}}\mathbf{y}^{\text{EIMC}}$ estimates a vertical tax reaction function in the presence of cooperation, which differs from the prior literature on vertical tax competition because the municipality formally cooperates with the higher tier government as part of the tax competition game. Thus, the higher level government’s tax rate is partially chosen with input from each individual municipality within the cooperative. As in Brülhart and Jametti (2019), given multiple “federal” governments (EIMCs), we control for neighboring tax rates in those EIMCs $\widetilde{\mathbf{W}}^{\mathbf{O}}\mathbf{y}^{\text{EIMC}}$ to account for diagonal tax competition (Agrawal 2016). Municipalities might respond to these EIMC tax rates differently than municipal tax rates if municipalities use other municipalities as a yardstick for their tax rates, but do not compare themselves to other EIMCs.

Focusing on the first column using all observations, ρ^E is negative, suggesting that municipalities offset the tax changes of their own EIMC. In part, this may be the result of a transfer of some public services to the EIMC, reducing the tax revenue needs of the municipality. The magnitude of the effect suggests municipalities offset EIMC tax changes by less than one-for-one, consistent with other negatively sloped vertical reaction functions in the literature. On the other hand, ρ^O is positive albeit close to zero: when neighboring EIMCs lower their tax rate, nearby municipalities also lower their tax rate to remain competitive, just as they would respond to other municipalities. However, this effect resulting from diagonal tax competition is much smaller than the response to nearby municipal rates and is statistically insignificant. In this manner, municipalities compete more intensely with other municipalities outside their EIMC than with other EIMCs. This may be because municipalities often have access to a broader set of taxing instruments than EIMCs (except under the additional taxation regime) or because municipalities use

other municipalities (not EIMCs) as a yardstick.

Using all municipalities to estimate vertical reactions functions includes towns in EIMCs with the single taxation regime. In this regime, the EIMC does not cohabit the same tax base as the municipalities as is usually the case with vertical tax competition. Further, in the mixed taxation regime, municipalities are no longer able to choose the business tax rate. Thus, in column (2) we drop towns in the single tax regime, while in column (3) we drop towns in the single tax or mixed taxation regime. The qualitative effects are similar, although the magnitudes and significance of the effects change slightly. There might be reasons to prefer using all municipalities, however. First, although EIMCs do not cohabit the same tax bases as towns in the single taxation regime, EIMC tax changes on the business tax rate will affect the municipal tax rates on other bases if the tax bases are related—in other words, if the cross-price elasticity of the municipal tax bases with respect to the business tax rate are non-zero. Indeed because the business tax covers some of the same properties as the developed land tax, there are persuasive reasons why this may be the case. At the same time, this complicates the interpretation as whether the bases are complements or substitutes likely matters (Hoyt 2017). Second, although the mixed tax regime forces municipal rates on one base to be zero, this choice may be the outcome of vertical tax competition.

The prior table is useful to help shed light on the mechanism, but it also raises concerns that our baseline model omits controls for vertical tax competition (Besley and Rosen 1998). To address this concern, we supplement (3) by including the EIMC tax rates above in addition to the municipal tax rates already in the specification. After including the EIMC tax rate, the coefficient on municipal tax rates within the EIMC is still significantly lower than the coefficient for municipalities outside of the EIMC. And consistent with the results above, the slope of the vertical reaction function is negative.

5.6 Robustness

Given one of the novelties of our paper is the creation of a tax index to study tax competition with multiple instruments, we wish to further verify the sensitivity of the

results to the index creation methodology. In all tables, we have already shown the results are robust to utilizing weights ν_l at the start or end of our sample. Table A.4 shows that the results are also robust to using time varying weights from the national distribution of tax revenue. However, when using an unweighted index, so that each taxing instrument receives equal weight, ρ^E is much smaller than ρ^O , though our relative conclusions remain the same. Given some taxes represent less than 5% of municipal revenue, we think that it is unlikely that municipalities strategically utilize all four rates equally. Finally, in the second panel of this table, we construct the weights using the national distribution of total (EIMC plus municipal) revenue, rather than the national distribution of only municipal revenue. Again, results are similar.

Finally, Appendix A.9 outlines an alternative identification strategy, with the results presented in Table A.5. Intuitively, we show the results are robust to resolving the endogeneity of the spatial weights matrix using an alternative IV approach that does not rely on us predicting the spatial weights but instead using exogenous proxies for the weight matrices directly.

6 Conclusion

Voluntary tax cooperation dampens tax competition. Municipalities strategically respond less intensely to the average tax rate of their peer group and than with the average tax rate of municipalities that they do not have a relationship with. In the process, we have made methodological advances to the tax competition literature involving endogenous network formation. To do so, we have assembled comprehensive panel data on municipal tax rates and historical data on other forms of municipal cooperation. Assembling and working with such large spatial weights matrices involves substantial computation intensity, but also allows the researcher to answer questions not possible previously. The methods we have used to overcome substantial threats to identification and the methods used to study multiple simultaneously determined policy levers are novel and they significantly advance the state of the applied tax competition literature. They can be applied to study the effect of voluntary cooperation among small jurisdictions in most federal systems.

Like with all empirical designs, our empirical method comes with assumptions. We show that intermunicipal cooperation is important to local outcomes and that it is endogenous to the tax competition game. Failure to account for the fact that the decision to cooperate and with whom to cooperate is a choice variable yields biased outcomes.

Our methods can be extended to study firm, worker, and trade groups. In the case of firms, collaboration may dampen competition. Recent advances in patent data or shared distribution networks may facilitate how bilateral links between firms affect their strategic interactions with firms not linked on the same network. Our methods could be useful to the IO estimation of strategic reaction functions (Brueckner and Luo 2014) in the presence of such distribution networks. With respect to international trade and international relations, the establishment of regional “free trade” organizations that foster policy coordination, may result in trade creation within the cooperative, but whether they encourage trade diversion from outside of the organization is an empirical question. In addition, there is great debate over whether the formation of the European Union has dampened tax competition. Our paper is also relevant to studying tax competition at the international setting and the influence of supra-national institutions on such competition.

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Table 1: Tax Regimes: Total Tax Rate in a Municipality

	No Cooperation	Additional Taxation	Single Taxation	Mixed Taxation
Business Tax	τ^B	$\tau^B + T^B$	T^B	T^B
Residence Tax	τ^R	$\tau^R + T^R$	τ^R	$\tau^R + T^R$
Developed Property Tax	τ^{DP}	$\tau^{DP} + T^{DP}$	τ^{DP}	$\tau^{DP} + T^{DP}$
Undeveloped Property Tax	τ^{UP}	$\tau^{UP} + T^{UP}$	τ^{UP}	$\tau^{UP} + T^{UP}$
Share over the period 1994-2009	34%	44.3%	18.6%	3.1%

This table shows tax rates that are levied depending on the tax regime in force among the four existing tax regimes, i.e., the absence of cooperation and the three tax regimes that EIMCs can elect to setup. τ^l is the municipal tax rate on the tax base l and T^l is the EIMC rate for respective municipalities, with $l = B, R, DP, UP$. Source: Breuillé, Duran-Vigneron and Samson (2018).

Table 2: Summary Statistics on Tax Rates

	Observation	Mean	SD	Min	Max
Residence Tax	582,722	8.65	4.02	0	45.06
Developed Property Tax	582,722	12.57	6.01	0	61.95
Undeveloped Property Tax	582,722	40.86	26.03	0	342.23
Business Tax	582,722	7.73	6.37	0	31.74
(1994) weighted index	582,722	10.27	4.32	0	37.93
(2009) weighted index	582,722	10.26	4.17	0	38.18
raw average index	582,722	17.45	8.21	0	98.88

This is the mean, minimum, maximum, and standard deviation of tax rate across the years 1994-2009 for all municipalities. Note the tax rate can be very large given the tax base is on rents and not values, and those rents are determined using historical valuations.

Table 3: Baseline OLS Results

	1994 Index		2009 Index	
	(1)	(2)	(3)	(4)
$\widetilde{\mathbf{W}}\mathbf{y}$	0.459*** (0.048)		0.524*** (0.054)	
$\widetilde{\mathbf{W}}^E\mathbf{y}$		0.338*** (0.042)		0.342*** (0.043)
$\widetilde{\mathbf{W}}^O\mathbf{y}$		0.425*** (0.041)		0.461*** (0.044)
Controls	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y
EIMC Controls	Y	Y	Y	Y
P-Value: $\rho^E = \rho^O$	n/a	0.148	n/a	0.057
Observations	582,722	582,722	582,722	582,722
Clusters	96	96	96	96

This table presents a baseline estimate of the single and partitioned weight matrix models. Columns (1)-(4) without a prime ignore all endogeneity concerns. Column (1) and (3) estimate equation (1) after row normalizing a single weight matrix while columns (2) and (4) estimate (3). The tax rate index is constructed using the weights at the top of the table. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.

Table 4: Classic IV Assuming Exogenous Weights

	1994 Index		2009 Index	
	(1)	(2)	(3)	(4)
$\widetilde{\mathbf{W}}_{\mathbf{y}}$	0.537*** (0.072)		0.610*** (0.077)	
$\widetilde{\mathbf{W}}^{\mathbf{E}}_{\mathbf{y}}$		0.021 (0.045)		0.037 (0.043)
$\widetilde{\mathbf{W}}^{\mathbf{O}}_{\mathbf{y}}$		0.571*** (0.074)		0.635*** (0.079)
Controls	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y
EIMC Controls	Y	Y	Y	Y
P-Value: $\rho^E = \rho^O$	n/a	0.000	n/a	0.000
Observations	582,722	582,722	582,722	582,722
Clusters	96	96	96	96

This table presents a baseline estimate of the single and partitioned weight matrix models. Columns (1)-(4) estimate the models with spatial IV, ignoring endogeneity of the spatial weight matrix. Column (1) and (3) estimate equation (1) after row normalizing the weight matrix while columns (2) and (4) estimate (3). The tax rate index is constructed using the weights at the top of the table. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. Standard errors are clustered at the department level.

Significance: *90%, ** 95%, ***99%.

Table 5: Predicting the Weight Matrix Elements

	(1) 1994	(2) 1995	(3) 1996	(4) 1997	(5) 1998	(6) 1999	(7) 2000	(8) 2001
Syndicate Membership ($t - 39$ or earlier) F Statistics	0.0142*** (0.00005)	0.0185*** (0.00006)	0.0219*** (0.00007)	0.0261*** (0.00007)	0.0296*** (0.00008)	0.0317*** (0.00008)	0.0371*** (0.00009)	0.0419*** (0.00009)
	66,670	91,050	103,100	121,500	143,600	153,600	187,100	217,600
	(9) 2002	(10) 2003	(11) 2004	(12) 2005	(13) 2006	(14) 2007	(15) 2008	(16) 2009
Syndicate Membership ($t - 39$ or earlier) F Statistics	0.0506*** (0.0001)	0.0576*** (0.0001)	0.0668*** (0.0001)	0.0729*** (0.0001)	0.0751*** (0.0001)	0.0757*** (0.0001)	0.0782*** (0.0001)	0.0804*** (0.0001)
	275,300	316,600	435,000	509,400	536,000	549,400	578,400	602,100

This table presents the results of the regression given by equation (4). To do so, we regress the decision to cooperate in the given year for a pair of municipalities less than \bar{d} away, on whether that pair formed a syndicate 39 years before period t or earlier. We estimate this regression for separately for each year t in our data. Significance: *90%, ** 95%, ***99%.

Table 6: Kelejian and Piras (2014) Results: Allowing for Endogenous EIMC Formation

	45 km Threshold			
	1994 Index		2009 Index	
	(1)	(2)	(3)	(4)
$\widetilde{\mathbf{W}}^E \mathbf{y}$	0.244*** (0.070)	0.307*** (0.074)	0.234*** (0.068)	0.290*** (0.071)
$\widetilde{\mathbf{W}}^O \mathbf{y}$	0.649*** (0.073)	0.580*** (0.071)	0.676*** (0.077)	0.595*** (0.076)
Controls	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.000	0.006	0.000	0.006
Observations	582,722	582,722	582,722	582,722
Clusters	96	96	96	96
F statistic	65.05	42.99	76.35	89.02
Hansen J statistic	0.266	0.343	0.313	0.366

This table presents estimates using the Kelejian and Piras (2014) approach, which resolves endogeneity of the spatial weight matrix and the spatial lag. All columns estimate (3) following Kelejian and Piras (2014). The tax rate index is constructed using the weights at the top of the table. Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic. The Hansen J statistic reported, which takes account of the fitted weight matrices, is the p-value. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.

Table 7: KP Results for Different Distance Thresholds

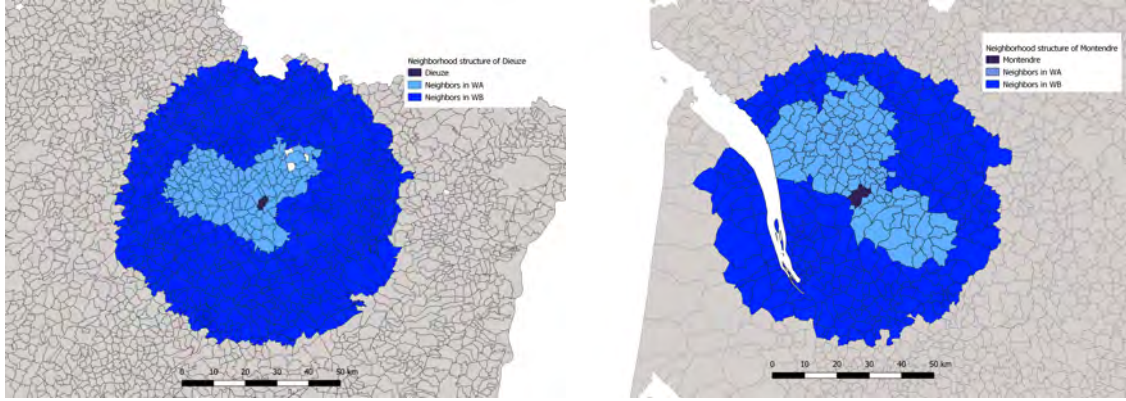
	20 km Threshold				7.25 km Threshold			
	1994 Index (1)	(2)	2009 Index (3)	(4)	1994 Index (1')	(2')	2009 Index (3')	(4')
\widetilde{W}^{E_y}	0.305*** (0.070)	0.362*** (0.076)	0.287*** (0.067)	0.345*** (0.079)	0.319*** (0.063)	0.299*** (0.058)	0.300*** (0.058)	0.301*** (0.062)
\widetilde{W}^{O_y}	0.660*** (0.087)	0.598*** (0.094)	0.686*** (0.092)	0.601*** (0.105)	0.558*** (0.118)	0.591*** (0.115)	0.550*** (0.124)	0.535*** (0.121)
Controls	Y	Y	Y	Y	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.002	0.095	0.002	0.116	0.072	0.057	0.096	0.163
Observations	582,722	582,722	582,722	582,722	582,722	582,722	582,722	582,722
Clusters	96	96	96	96	96	96	96	96
F statistic	156	120.4	165.1	76.97	63.63	96.61	153.6	111.8
Hansen J statistic	0.461	0.155	0.315	0.141	0.297	0.325	0.414	0.349
This table presents estimates using the Kelejian and Piras (2014) approach for different values of \bar{d} . Columns without a prime use a distance threshold of 20 km, while columns with a prime use a distance threshold of 7.25 km. In the case of this last threshold, we select the distance threshold such that both weight matrices have the same number of municipalities in them on average. All columns estimate (3) following Kelejian and Piras (2014). The tax rate index is constructed using the weights at the top of the table. Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic. The Hansen J statistic reported, which takes account of the fitted weight matrices, is the p-value. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.								

Table 8: Vertical Tax Competition and EIMC Tax Rates

Dependent Variable	Municipal Rate					
	1994 Index		2009 Index		2009 Index	
Explanatory Variable	EIMC Rate (1)	EIMC Rate (2)	EIMC Rate (3)	EIMC Rate (1')	EIMC Rate (2')	EIMC Rate (3')
$\widetilde{W}E_y^{EIMC}$	-0.771*** (0.160)	-0.666*** (0.136)	-0.787*** (0.125)	-0.784*** (0.185)	-0.656*** (0.144)	-0.789*** (0.128)
$\widetilde{W}O_y^{EIMC}$	0.084 (0.123)	0.143 (0.101)	0.242*** (0.087)	0.111 (0.149)	0.140 (0.118)	0.252** (0.102)
Controls	Y	Y	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y	Y	Y
EIMC Controls	Y	Y	Y	Y	Y	Y
P-Value: $\rho^E = \rho^O$	0.001	0.000	0.000	0.001	0.001	0.000
Observations	582,722	474,217	456,294	582,722	474,217	456,294
Clusters	96	96	96	96	96	96
F statistic	27.10	27.11	31.89	46.78	22.57	34.23
Hansen J statistic	0.272	0.450	0.450	0.232	0.433	0.423

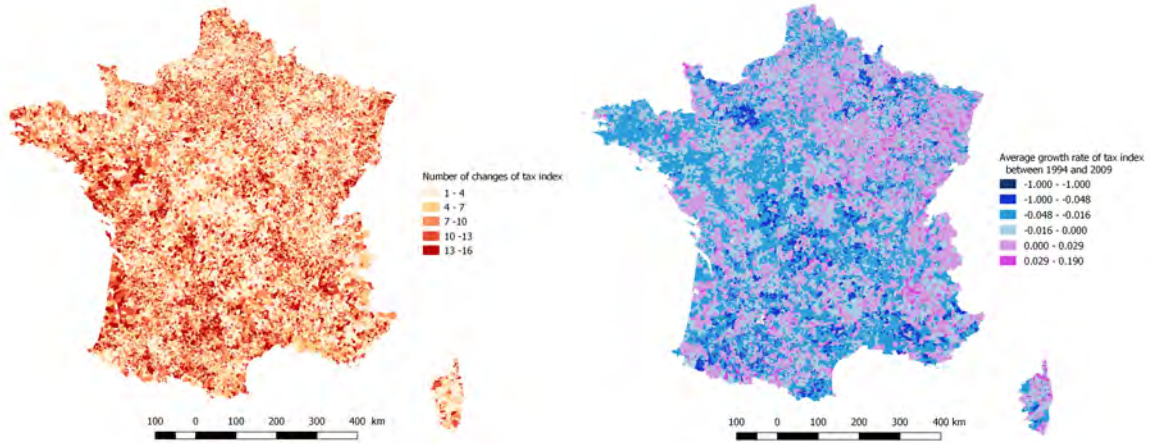
Relative to the prior table, the independent variables are the tax index for EIMC alone. The dependent variable is still the municipal tax rate. All columns estimate (3) following Kelejian and Piras (2014). Column (1) includes all municipalities, column (2) drops municipalities in the single taxation regime, and column (3) drops municipalities in the mixed taxation regime. The tax rate index is constructed using the weights at the top of the table. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic and the Hansen J statistic is the p-value. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.

Figure 1: Examples of Weight Matrices Partition for Towns



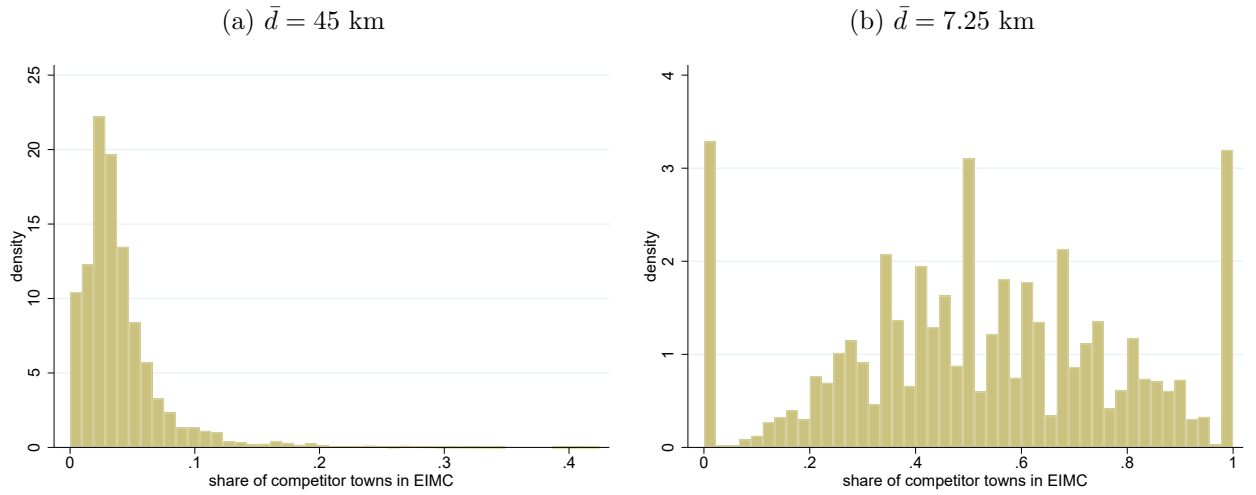
These two figures show the \mathbf{W}^E (light blue) and \mathbf{W}^O (dark blue) matrix for two towns in France. The light colored towns are within the same EIMC while the dark colored towns are outside of the EIMC, but still in close proximity to the given municipality. Note that for each town in a given EIMC, the dark blue ring extends in forty-five kilometers so that it includes a different set of town for different municipalities. The regions not shaded in blue are sufficiently far away that they are assumed to not be linked with the municipality.

Figure 2: Spatial Variation in Tax Rate Changes



These two figures show the number of tax changes and the average growth rate of tax changes at the municipality level. Both maps are constructed using our tax index variable, which holds fixed the weights on each of the four tax instruments such that all changes are driven by statutory tax rate changes and not changes in the weights used to construct the index.

Figure 3: Distribution of Share of Municipalities within \bar{d} that Are in the EIMC



These two figures show the distribution of the spatial weights by depicting the share of municipalities within the distance threshold \bar{d} that are in the EIMC. Thus, for example, 0.3 indicates that 30% of towns within \bar{d} kilometers are in the EIMC. The distributions change over time, and these figures are for 2009.

Online Appendix for “Tax Competition with Intermunicipal Cooperation”

David R. Agrawal, Marie-Laure Breuillé, and Julie Le Gallo

A Appendices

A.1 Additional Institutional Details

In this section, we discuss the formal laws and types of cooperation that arise in France.

Municipal cooperation was codified and formally introduced in the 22nd March 1890 Law, although informal cooperation of municipalities predates the law. Under the 1890 law, municipalities were allowed to form “syndicat intercommunal à vocation unique,” (SIVU) which were a group of municipalities which jointly provide a unique public good or service (clean water, sewage treatment, waste collection, curricular and co-curricular activities, etc.). Prior to World War II, these syndicates were not widespread. However, laws passed in 1955 and 1959 helped to encourage their formation. In particular, in 1955 “syndicat mixte” (SM) were allowed to be formed. These syndicates allowed municipalities to cooperate with other municipalities while also cooperating with other tiers of government or public bodies to provide one or more public services. Since 1959, municipalities were able to form “syndicat intercommunal à vocation multiple” (SIVOM) which are groups of municipalities that are able to perform several public services. It is important to note that these cooperative units cannot vote on tax rates, but rather are financed directly by contributions from their members. These contributions are either “contributions budgétaires”, i.e., an expenditure item in the budget of each member municipality, or much more rarely “contributions fiscalisées”, i.e., tax revenues voted and collected by each member municipality on at least one tax base for the direct benefit of these syndicates that only choose the amount of expected “contributions fiscalisées”. This last form of contribution, similar to additional taxation, is more transparent for the taxpayer and gives increased autonomy to the cooperative unit, as the funding does not go through municipal budgets. It should be noted that the tax rates associated to “contributions fiscalisées” are isolated from municipal budgets and therefore excluded from our analysis.

The post-World War II period proved to be a very important time period for the growth of municipal cooperation in France. For example, almost 2000 municipalities created a syndicate in 1955, 1720 municipalities created a syndicate in 1960, 2122 in 1965, 1763 in 1970 and 2298 in 1974. Figure A.1 shows the new syndicates created prior to the year 1955 and the new syndicates created prior to 1974. These syndicates had no taxation

authority, their duties were limited to issues of facilities, infrastructure, or networks and municipalities could leave the syndicates at any time. Although these three syndicate structures persist to today, they have been declining in popularity. Our predictor variable will exploit the idiosyncratic timing of syndicates created in period $t - 39$, while also accounting at the same time for the stock of syndicates formed earlier.

Recognizing the need for more formal cooperation, the French government passed a series of laws designed to encourage what we label as current day inter-municipal cooperation with own-source tax revenue. The law of 31 December 1966 made possible the creation of “communautés urbaines” (CU) in agglomerations of more than 50,000 inhabitants. The law of 31 December 1970 facilitated inter-municipal cooperation, in particular by extending to rural municipalities the possibility of setting up districts. The “law of 6 February 1992 on the territorial administration of the Republic” established “communautés de communes” (CC), which must contain less than 50,000 inhabitants and “communautés de villes” for urban areas. The so-called Chevènement law of 1999 on the strengthening and simplification of inter-municipal cooperation deeply reformed the modalities of inter-municipal cooperation by suppressing some forms of EIMCs (“districts”, “communautés de villes”), by improving the functioning of others (“communautés urbaines” with a new population threshold of 500,000) and by creating new EIMCs (“communautés d’agglomération” (CA) which must have a population between 50,000 and 500,000 inhabitants). The purpose of these reforms was to grant taxation authority and a wider range of potential functions than the historical syndicates. For example, CC must construct local development plans, waste collection, and spatial planning. In addition, it can carry out additional functions over environmental protection, housing, road maintenance, cultural and sports facility construction, sanitation, and water provision. More integrated levels of the EIMC have an even larger set of services that can be provided. The Chevènement law requires EIMCs to comprise a group of contiguous municipalities without any enclaves and prohibits membership in more than one EIMC. However, municipalities are free to make decisions concerning EIMC membership: they are free to leave, form their own EIMC, or join an existing EIMC.

Finally, we describe the calculation basis for the four main taxes. The tax on developed properties is paid by owners on built real property. The owner is statutorily responsible for the tax. Rural buildings assigned permanently for agricultural use are exempt and localities may create other exemptions. The taxable value is determined each year and is equal to half the cadastral *rental* value. The property tax on undeveloped land is statutorily paid by the owners of undeveloped land. The tax base includes land assigned to farms, quarries and mines, bodies of water, rural buildings, and railway land. In general, the base is equal to the cadastral rental value reduced in part to reflect the costs of agricultural properties. The residency tax applies regardless of ownership to rental or free occupancy buildings in a municipality. The tax applies to all furnished premises for

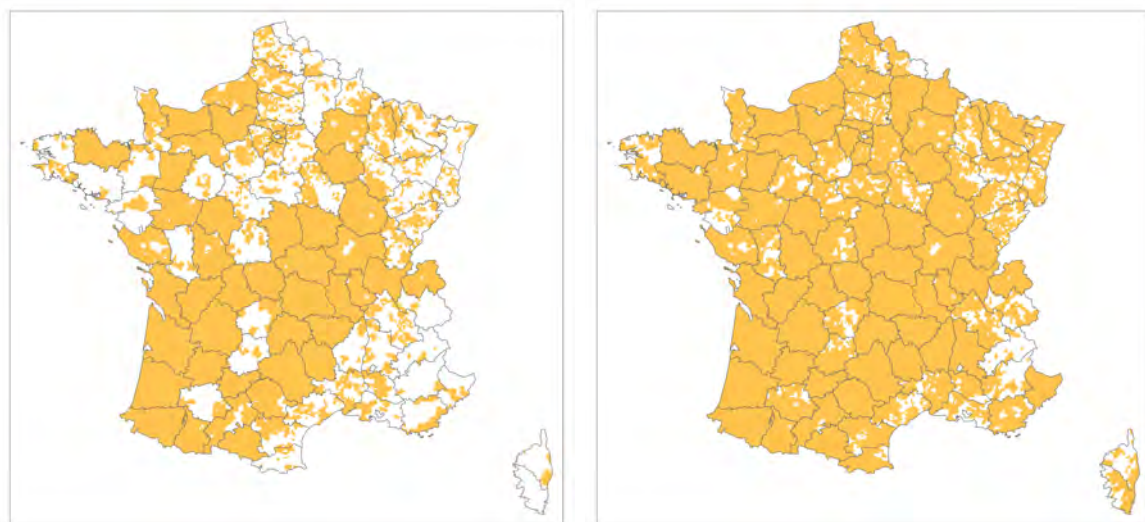
residence (primary or secondary) and any dependent structures. This tax is statutorily levied on the person occupying the unit. Some exemptions exist. Again, the value of the property is determined by the cadastral rental value, but it decreases with allowances. The business tax generally concerns company properties and certain exemptions are granted based on turnover. Until 2002, the tax was levied on the rental value of buildings, tangible assets and on 18% of the salaries paid by the business. The last of these three bases was phased out and eliminated following 2002. Additionally, the business tax covers assets such as computers and machinery. Structures owned by businesses are thus subject to the property tax and business tax. As of 2010, the business tax was entirely eliminated.

A.2 Maps and Descriptive Stats

Figure A.1 shows syndicates created over the periods 1900-1955 and 1900-1974 while Figure A.2 shows the evolution of EIMCs in France from 1994 to 2009. As can be seen, both syndicates and EIMCs became more common over time, justifying the need to have some time variation when predicting EIMC membership. Figure A.3 shows tax trends for each specific taxing instrument.

Figure A.1: Syndicates Created, 1900-1955 and 1900-1974

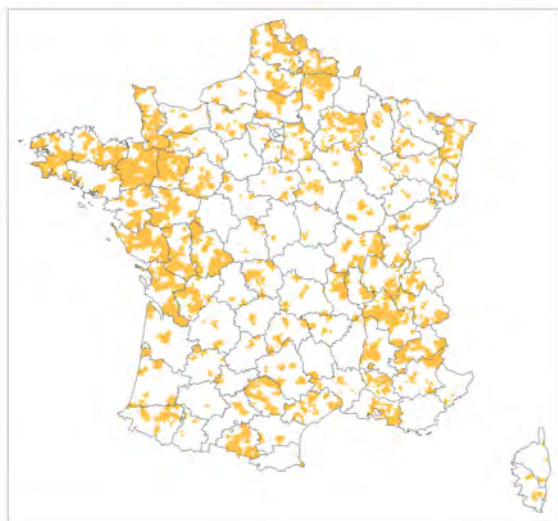
(a) Syndicates Created between 1900 and 1955 (b) Syndicates Created between 1900 and 1974



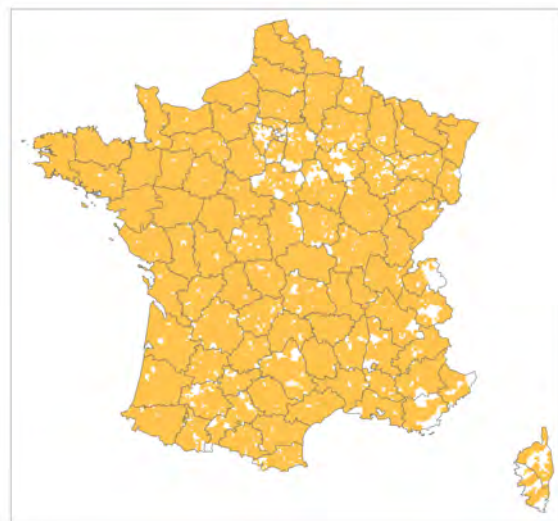
This figure shows syndicates created before 1955 and 1974.

Figure A.2: EIMC Evolution in France, 1994 to 2009

(a) EIMCs in 1994

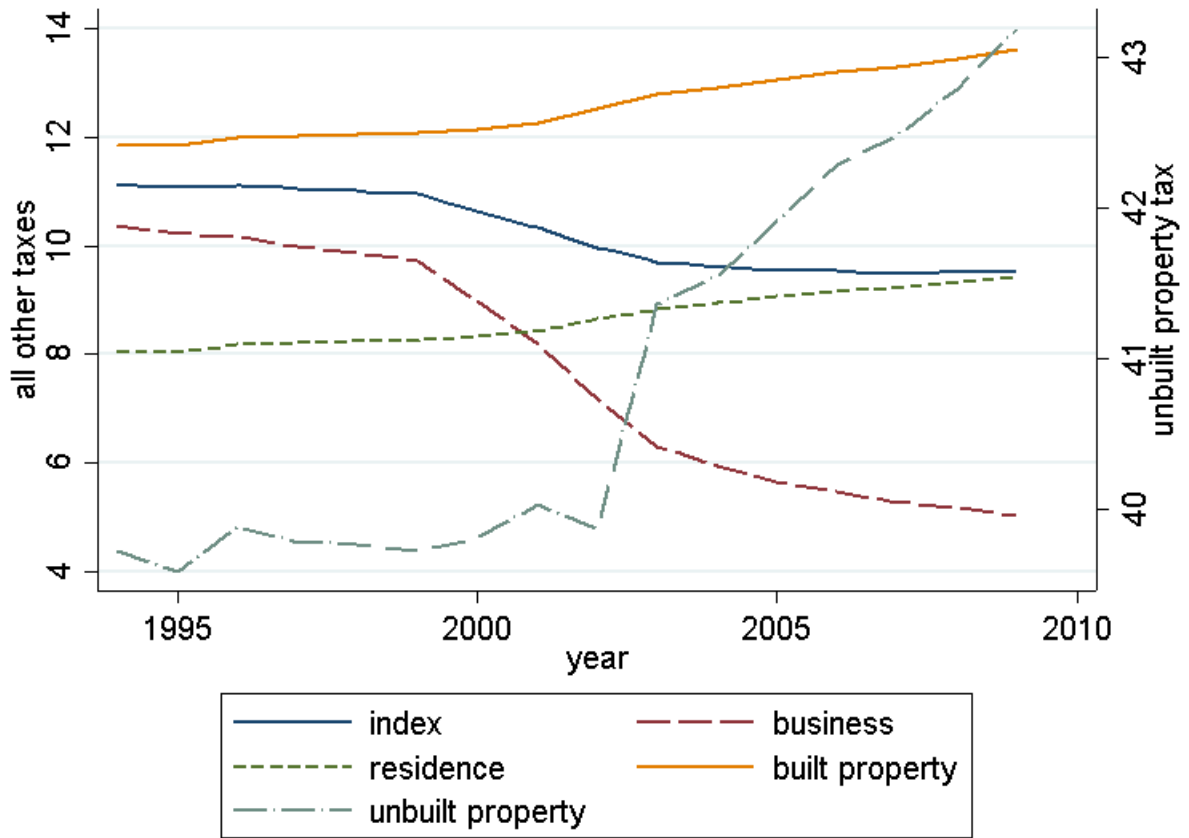


(b) EIMCs in 2009



This figure shows EIMCs that existed in 1994 and 2009.

Figure A.3: Tax Trends



These two figures show the trends in the individual taxes and the tax index over time. We present the mean value for each year. Note that the unbuilt property tax uses a different scale for the vertical axis (right axis) as all other taxes (left axis).

A.3 Further Details on Taxes and Weights

The business tax base consists of a portion of the amount of any wages and fees paid (until the abolition of this part in 2003) and the rental value of any tangible fixed assets—e.g., land, buildings, physical equipment, etc. As such, the business tax is comparable to a tax on the value of capital (and historically, labor) used by the firm during the production process. The residential tax base (“taxe d’habitation”, whose end is planned for 2023, except for secondary residences and vacant properties) is an housing tax for primary and second houses paid by dwelling occupants (tenants or owners); numerous tax exemptions or deductions are applied for poor and dependent taxpayers. The developed property tax base (“taxe sur le foncier bâti”) is paid by the owners and users of buildings (housing, car parks, outbuildings, shops, business and industrial premises, ...), while the undeveloped property tax (“taxe sur le foncier non bâti”) is paid by the owners and usufructuaries of undeveloped land.

The weights are the fraction of aggregated municipal revenue from each tax source

(in the first year of our sample). In order to do this, we calculate the total amount of revenue raised by each tax instrument across all municipalities in France and divide it by the total revenue raised from all four tax instruments. In this way, no one municipality can unduly influence the weights in the index. We hold the weights constant across municipalities and years so that any changes are driven by actual changes in tax rates. In particular the weighted tax index is calculated as $.471459\tau^B + .0308044\tau^{UP} + .2702719\tau^{DP} + .2274647\tau^R$. In the paper, we verify alternative specifications of different weights. One concern is that there may have been broad secular change in the importance of each tax instrument over our panel (as evidenced in figure A.3). To verify the robustness of this, we calculate the weights using the last year of our sample and re-estimate all models.

In addition, given the various EIMC tax regimes, we also verify the results are robust to calculating the tax index using fixed weights based on the EIMC and municipal revenue sources. We also show the results using time varying weights, where we calculate the fraction of revenue from each tax base each year. Finally, we construct a naive index that simply places equal weight on all tax rates.

A.4 Theoretical Model

To interpret the results of our empirical approach, we provide a general model of tax interactions among municipalities that cooperate via the formation of EIMCs. Let i index a finite number of municipalities. Municipality i 's government chooses the tax policy that maximizes a general welfare function $\Psi^i = \Psi(t_i, T_E, b(t_i, \mathbf{t}_{-iE}, \mathbf{t}_{-iO}, \mathbf{T}), B(t_i, \mathbf{t}_{-iE}, \mathbf{t}_{-iO}, \mathbf{T}))$, where t_i is municipality i 's tax rate, \mathbf{t}_{-iE} is the vector of tax rates of municipalities other than i which belong to the same EIMC, \mathbf{t}_{-iO} is the vector of tax rates of municipalities which are not in the EIMC of municipality i , T_E is the tax rate of the EIMC to which municipality i belongs to and \mathbf{T} is the vector of all EIMCs' tax rates. The welfare function depends on the level of the tax rates and on the respective tax bases where $b(\cdot)$ is the municipality's tax base and $B(\cdot)$ is the EIMC's tax base.

This general welfare function has several advantages, and generalizes much of the prior tax competition literature both in its functional form (Wilson 1999) and how vertical and horizontal externalities interact (Brühlhart and Jametti 2006; Keen and Kotsogiannis 2004). First, by using a general function $\Psi(\cdot)$ we do not impose that the governments maximize tax revenue or social welfare, which allows us to take no stance on the political economy motives of government. Second, by having $\Psi(\cdot)$ be a function of the tax rates and tax bases, we do not need to model the role of private consumption, which also depends on the tax rates. To be clear, our functional form allows for the presence of private consumption but does not place any assumptions on the elasticity of substitution between public and private goods.

Municipal governments are linked together through their tax policies, inside the same EIMC and outside the EIMC, via the functions $b(\cdot)$ and $B(\cdot)$. This linkage can be commuting patterns from the residence to the workplace (Wildasin 2014), commercial transactions (Kanbur and Keen 1993) or upstream/downstream industrial linkages (Wildasin 1993). Critically, the tax base is negatively related to the own-jurisdiction tax rate and positively related to the competitor jurisdiction tax rates. The aggregate tax base at the country level can be either exogenous or endogeneous. Each municipality i optimizes with respect to its tax rate t_i , taking other municipal and inter-municipal tax rates as given. We assume the functions $\Psi(\cdot)$, $b(\cdot)$ and $B(\cdot)$ satisfy the necessary conditions for an interior maximum and a positive maximizing tax rate. At the Nash equilibrium, municipality i selects t_i such that: $\frac{\partial \Psi^i}{\partial t_i} \equiv \Psi_{t_i} + \Psi_b b_{t_i} + \Psi_B B_{t_i} = 0$, where subscripts on $\Psi(\cdot)$, $b(\cdot)$ and $B(\cdot)$ denote partial derivatives. The simultaneous solution of the municipal and inter-municipal optimality conditions yields the Nash non-cooperative equilibrium $(t_i^*, \mathbf{t}_{-i\mathbf{E}}^*, \mathbf{t}_{-i\mathbf{O}}^*, \mathbf{T}^*)$, assuming it exists. As our interest is in the slope of the strategic interaction functions, we apply the implicit function theorem on the first-order condition and differentiate with respect to an element t_{kE} of $\mathbf{t}_{-i\mathbf{E}}$ or t_{kO} of $\mathbf{t}_{-i\mathbf{O}}$, and evaluate at $(t_i^*, \mathbf{t}_{-i\mathbf{E}}^*, \mathbf{t}_{-i\mathbf{O}}^*, \mathbf{T}^*)$. Assuming $\Psi_{bB} = \Psi_{Bb} = 0$, we obtain:

$$\begin{aligned} \frac{dt_i}{dt_{kE}} &= -\frac{\partial^2 \Psi^i / \partial t_i \partial t_{kE}}{\partial^2 \Psi^i / \partial t_i \partial t_i} = -\frac{\partial \Psi_{t_i} / \partial t_{kE} + \partial \Psi_b b_{t_i} / \partial t_{kE} + \partial \Psi_B B_{t_i} / \partial t_{kE}}{\Psi_{t_i b} b_{t_i E} + \Psi_{t_i B} B_{t_i E} + \Psi_{bb} b_{t_i E} b_{t_i} + \Psi_b b_{t_i t_{kE}} + \Psi_{BB} B_{t_i E} B_{t_i} + \Psi_B B_{t_i t_{kE}}} \\ &= -\frac{\partial \Psi_{t_i} / \partial t_{kE} + \partial \Psi_b b_{t_i} / \partial t_{kE} + \partial \Psi_B B_{t_i} / \partial t_{kE}}{SOC}, \end{aligned} \quad (\text{A.1})$$

$$\begin{aligned} \frac{dt_i}{dt_{kO}} &= -\frac{\partial^2 \Psi^i / \partial t_i \partial t_{kO}}{\partial^2 \Psi^i / \partial t_i \partial t_i} = -\frac{\partial \Psi_{t_i} / \partial t_{kO} + \partial \Psi_b b_{t_i} / \partial t_{kO} + \partial \Psi_B B_{t_i} / \partial t_{kO}}{\Psi_{t_i b} b_{t_i O} + \Psi_{t_i B} B_{t_i O} + \Psi_{bb} b_{t_i O} b_{t_i} + \Psi_b b_{t_i t_{kO}} + \Psi_{BB} B_{t_i O} B_{t_i} + \Psi_B B_{t_i t_{kO}}} \\ &= -\frac{\partial \Psi_{t_i} / \partial t_{kO} + \partial \Psi_b b_{t_i} / \partial t_{kO} + \partial \Psi_B B_{t_i} / \partial t_{kO}}{SOC}. \end{aligned} \quad (\text{A.2})$$

The second-order condition guarantees that the denominator is negative. Let k index a municipality other than i . Therefore, the signs of the tax reaction functions $\frac{dt_i}{dt_{kE}}$ and $\frac{dt_i}{dt_{kO}}$ are, respectively, determined by the cross-derivatives $\frac{\partial^2 \Psi^i}{\partial t_i \partial t_{kE}}$ and $\frac{\partial^2 \Psi^i}{\partial t_i \partial t_{kO}}$, which are ambiguous in the absence of additional assumptions. Critically, notice that if the change in the base or the marginal welfare for a given competitor's (E or O) tax rate is different within or outside the EIMC, the slope of the reaction functions will be different. *Only if all changes and rates of change are the same within and external to the EIMC, will tax competition be equally intense.*

Obviously we wish to start with the most general model, but to highlight intuition, we place more structure on the functional form to shed light on the mechanisms that might underlie our empirical analysis.

Consider first a simple case with (local government) revenue maximization $\Psi^i = t_i b(t_i, \mathbf{t}_{-i\mathbf{E}}, \mathbf{t}_{-i\mathbf{O}}, \mathbf{T})$ implying $\frac{\partial \Psi^i}{\partial t_i} = b(t_i, \mathbf{t}_{-i\mathbf{E}}, \mathbf{t}_{-i\mathbf{O}}, \mathbf{T}) + t_i b_{t_i} = (1 + \epsilon_{b, t_i}) b = 0$, where $\epsilon_{b, t_i} \equiv b_{t_i} \frac{t_i}{b}$ is the elasticity of the tax base with respect to t_i . We have $\frac{dt_i}{dt_{kE}} SOC = -(b_{t_{kE}} + t_i b_{t_i t_{kE}})$. Then, as tax rates are strategic complements in this example, the reaction function with respect to other towns in the EIMC will be less steeply sloped than

with respect to other towns outside of the EIMC if $b_{t_{kE}} + t_i b_{t_i t_{kE}} < b_{t_{kO}} + t_i b_{t_i t_{kO}}$. If the second order terms are small or equal, this will hold true if the change of the tax base with respect to t_{kE} is smaller than the derivative of the tax base with respect to t_{kO} . Such a prediction is possible, but is inconsistent with theoretical models where jurisdictions are linked by commuting (Wildasin 2014). Our empirical results shed light on the welfare function, suggesting that it is likely governments maximize more than revenue.

Therefore, consider now the maximization of a standard quasi-linear function $\Psi^i = c^i + \lambda \log(g^i) + \Gamma \log(G^E)$, where λ and Γ are strictly positive parameters, c^i is the private consumption, g^i a public good financed by the tax revenues $t_i b^i$ of the municipality i and G^E an inter-municipal public good financed by the tax revenues $T_E B^E$ of the EIMC it belongs to. Suppose that the inter-municipal tax base is the sum of member municipalities' tax bases, i.e., $B^E(t_i, \mathbf{t}_{-iE}, \mathbf{t}_{-iO}, \mathbf{T}) = b^i + \sum_{j \neq i} b^{jE}$. A main difference in the choice of neighboring tax rates is that municipalities other than i which belong to the same EIMC are likely to internalize more tax externalities (through \mathbf{t}_{-iE}) on $B(\cdot)$ than municipalities other than i which belong to others EIMCs (through \mathbf{t}_{-iO}) as their citizens enjoy the consumption of the inter-municipal public good. Analytically, at the Nash equilibrium, municipality i selects t_i such that: $\frac{\partial \Psi^i}{\partial t_i} \equiv c_{t_i}^i + \frac{\lambda}{g_i}(b^i + t_i b_{t_i}^i) + \frac{\Gamma}{G_E} T_E (b_{t_i}^i + \sum_{j \neq i} b_{t_i}^{jE}) = 0$. Applying the implicit function theorem on the first-order condition and differentiating with respect to an element k of \mathbf{t}_{-iE} or \mathbf{t}_{-iO} , we obtain at $(t_i^*, \mathbf{t}_{-iE}^*, \mathbf{t}_{-iO}^*, \mathbf{T}^*)$:

$$\frac{dt_i}{dt_{kE}} = - \frac{\partial^2 \Psi^i / \partial t_i \partial t_{kE}}{\partial^2 \Psi^i / \partial t_i \partial t_i} = - \frac{c_{t_i t_{kE}}^i - \frac{\lambda}{g_i^2} g_{t_i}^i g_{t_{kE}}^i + \frac{\lambda}{g_i} g_{t_i t_{kE}}^i - \frac{\Gamma}{G_E^2} G_{t_i}^E G_{t_{kE}}^E + \frac{\Gamma}{G_E} G_{t_i t_{kE}}^E}{SOC}, \quad (\text{A.3})$$

$$\frac{dt_i}{dt_{kO}} = - \frac{\partial^2 \Psi^i / \partial t_i \partial t_{kO}}{\partial^2 \Psi^i / \partial t_i \partial t_i} = - \frac{c_{t_i t_{kO}}^i - \frac{\lambda}{g_i^2} g_{t_i}^i g_{t_{kO}}^i + \frac{\lambda}{g_i} g_{t_i t_{kO}}^i - \frac{\Gamma}{G_E^2} G_{t_i}^E G_{t_{kO}}^E + \frac{\Gamma}{G_E} G_{t_i t_{kO}}^E}{SOC}. \quad (\text{A.4})$$

Plugging in the the derivatives and assuming that tax externalities on i 's marginal utilities for private consumption or public good provision are identical whether they are generated by a change in t_{kE} or t_{kO} ,²⁵ the difference between the two reaction functions²⁶ boils down to:

$$\frac{dt_i}{dt_{kE}} - \frac{dt_i}{dt_{kO}} = - \frac{-\frac{\Gamma}{G_E^2} T_E^2 \left(b_{t_i}^i + \sum_{j \neq i} b_{t_i}^{jE} \right) (b_{t_{kE}}^{kE} - b_{t_{kO}}^{kE}) + \frac{\Gamma}{G_E} T_E (b_{t_i t_{kE}}^{kE} - b_{t_i t_{kO}}^{kE})}{SOC}. \quad (\text{A.5})$$

The difference between the two reaction functions slopes, $\frac{dt_i}{dt_{kE}} - \frac{dt_i}{dt_{kO}}$, will depend

²⁵The assumption of tax externalities on i 's marginal utilities for private consumption or public good provision being equal means that $c_{t_i t_{kE}}^i = c_{t_i t_{kO}}^i, b_{t_{kE}}^i = b_{t_{kO}}^i, b_{t_i t_{kE}}^i = b_{t_i t_{kO}}^i, b_{t_{kE}}^{jE} = b_{t_{kO}}^{jE}, b_{t_i t_{kE}}^{jE} = b_{t_i t_{kO}}^{jE} \forall j \neq i, k$.

²⁶The detailed expressions of the tax reaction functions are $\frac{dt_i}{dt_{kE}} = \frac{c_{t_i t_{kE}}^i - \frac{\lambda}{g_i^2} (b^i + t_i b_{t_i}^i) t_i b_{t_{kE}}^i + \frac{\lambda}{g_i} (b_{t_{kE}}^i + t_i b_{t_i t_{kE}}^i) - \frac{\Gamma}{G_E^2} T_E^2 \left(b_{t_i}^i + \sum_{j \neq i} b_{t_i}^{jE} \right) (b_{t_{kE}}^i + \sum_{j \neq i, k} b_{t_{kE}}^{jE} + b_{t_{kE}}^{kE}) + \frac{\Gamma}{G_E} T_E (b_{t_i t_{kE}}^i + \sum_{j \neq i, k} b_{t_i t_{kE}}^{jE} + b_{t_i t_{kE}}^{kE})}{-SOC}$ and $\frac{dt_i}{dt_{kO}} = \frac{c_{t_i t_{kO}}^i - \frac{\lambda}{g_i^2} (b^i + t_i b_{t_i}^i) t_i b_{t_{kO}}^i + \frac{\lambda}{g_i} (b_{t_{kO}}^i + t_i b_{t_i t_{kO}}^i) - \frac{\Gamma}{G_E^2} T_E^2 \left(b_{t_i}^i + \sum_{j \neq i} b_{t_i}^{jE} \right) (b_{t_{kO}}^i + \sum_{j \neq i} b_{t_{kO}}^{jE}) + \frac{\Gamma}{G_E} T_E (b_{t_i t_{kO}}^i + \sum_{j \neq i} b_{t_i t_{kO}}^{jE})}{-SOC}$.

on the sign of $(b_{t_{kE}}^{kE} - b_{t_{kO}}^{kE})$, provided that the difference in the second-order effects are sufficiently small ($b_{t_i t_{kE}}^{kE} - b_{t_i t_{kO}}^{kE} \approx 0$). To see this, note that the second-order condition and the effect of a change in t_i on the inter-municipal tax base, $b_{t_i}^i + \sum_{j \neq i} b_{t_i}^{jE}$, are both negative while $\frac{\Gamma}{G^{E2}} T_E^2$ is positive. Thus the sign of (A.5) depends on the sign of the term $(b_{t_{kE}}^{kE} - b_{t_{kO}}^{kE})$, which captures the differential effect on the inter-municipal public good provision of the taxation of municipality kE in the EIMC and municipality kO outside the EIMC. This is likely to be negative. A rise in t_{kO} increases b^{kE} and therefore G^E due to tax base shifting, whereas a rise in t_{kE} has the opposite and stronger (in absolute terms) effect, leading to $\frac{dt_i}{dt_{kE}} - \frac{dt_i}{dt_{kO}} < 0$.

Jurisdiction i accounts for the fact that a rise in tax rate t_{kE} reduces the inter-municipal public goods provided to citizens of municipalities i and k that are members of the EIMC. This follows a reduction in the inter-municipal tax base induced by a leakage of k 's tax base ($b_{t_{kE}}^{kE} < 0$), and therefore raises i 's tax rate less than it would in response to an increase in the tax rate of an outside jurisdiction kO that does not belong to the same EIMC.

Inter-municipal cooperation, through the joint public good provision, is thus expected to induce less intense reactions among municipalities in the same EIMC to limit the reduction in the inter-municipal public good provision, G^E , by protecting the inter-municipal tax base common to i and k .

Empirically, we find tax competition among municipalities in France is less intense for towns within the same cooperative unit than across cooperative units. The model above rationalizes our empirical results. If governments are Leviathan, our results imply that tax rates of neighboring municipalities within the EIMC have a smaller impact on the tax base than taxes outside of the EIMC. If governments maximize a quasilinear welfare function, then under mild assumptions, our result is consistent with the provision of the EIMC public services dampening tax competition within the EIMC. The latter story seems consistent with other theories of multi-tiered taxation and public good provisions (Hoyt 2001; Revelli 2003). Our findings suggest that intermunicipal cooperation may reduce the parasitic effects of strategic interdependence. However, given the slope of the reaction function cannot be linked to whether taxes are too high or too low, we cannot make welfare implications.

A.5 Syndicate Correlation with Observables

Following the procedure in the text, Table A.1 tests whether towns that enter the same syndicate are similar (or different) from towns that are nearby but not a part of the same syndicate.

Table A.1: Syndicate Validity

	Intercept	Syndicate Indicator
Right Vote Share	54.168*** (0.133)	-0.051 (0.265)
Population	1,844.112*** (130.800)	-398.087 (261.599)
Share of Population, 0-14	18.761*** (0.032)	-0.022 (0.063)
Share of Population, ≥ 60	23.274*** (0.075)	0.111 (0.150)
Share of Population, ≥ 75	7.669*** (0.031)	0.044 (0.061)
Single Parent Families	7.747*** (0.029)	0.049 (0.058)
Density	153.682*** (8.348)	-22.839 (16.697)
Area	14.555*** (0.109)	0.285 (0.219)
Unemployment Rate	7.943*** (0.045)	0.029 (0.090)
Share Executives	6.291*** (0.051)	0.009 (0.102)
Number of Jobs	169.996*** (8.869)	-30.794* (17.738)
Tax Index	11.163*** (0.054)	0.018 (0.108)
Residence Tax	8.105*** (0.049)	0.027 (0.097)
Built Property Tax	12.029*** (0.093)	-0.043 (0.187)
Unbuilt Property Tax	39.458*** (0.276)	0.090 (0.552)
Business Tax	10.292*** (0.047)	0.044 (0.094)
Municipal FE	Y	Y

For each town i 's neighbors within 45 km, we calculate the average of each of the characteristics above, separately for its neighbors that it partnered as a syndicate with and neighbors that it did not participate in a syndicate with. This table regresses each characteristic above on a dummy for whether the neighbors were in a syndicate or not, a constant, and town fixed effects. Standard errors are clustered at the department level.

Significance: *90%, ** 95%, ***99%.

A.6 Additional Robustness Checks

Tables A.2 present additional results in the text. This table show results using different distance thresholds and also syndicate membership and commuting patterns to predict current day cooperation.

Table A.2: KP Results : Other Robustness Checks

	26 km Threshold				Past Commuting Flows as an Instrument			
	1994 Weights (1)	2009 Weights (2)	2009 Weights (3)	2009 Weights (4)	1994 Weights (1')	2009 Weights (2')	2009 Weights (3')	2009 Weights (4')
$\widetilde{W^E y}$	0.281*** (0.069)	0.351*** (0.076)	0.267*** (0.068)	0.334*** (0.077)	0.257*** (0.062)	0.341*** (0.068)	0.256*** (0.058)	0.329*** (0.064)
$\widetilde{W^O y}$	0.657*** (0.074)	0.587*** (0.078)	0.688*** (0.079)	0.601*** (0.088)	0.639*** (0.061)	0.578*** (0.060)	0.674*** (0.067)	0.600*** (0.067)
Controls	Y	Y	Y	Y	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.000	0.058	0.000	0.059	0.000	0.011	0.000	0.007
Observations	582,722	582,722	582,722	582,722	582,722	582,722	582,722	582,722
Clusters	96	96	96	96	96	96	96	96
F statistic	293.4	84.41	217.6	53	108.6	66.49	65.93	42.60
Hansen J statistic	0.612	0.277	0.476	0.228	0.232	0.339	0.261	0.298

This table presents estimates using the Kelejian and Piras (2014) approach, which resolves endogeneity of the spatial weight matrix and the spatial lag. Unlike the main results, columns (1)-(4) without a prime use a distance threshold of 26 km, which guarantees full coverage of EIMCs. In columns (1')-(4'), we also use historical commuting patterns an instrument for the spatial weights in addition to syndicate membership. All columns estimate (3) following Kelejian and Piras (2014). The tax rate index is constructed using the weights at the top of the table. Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the

Kleibergen-Paap rk Wald F statistic and the Hansen J statistic is the p-value. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.

A.7 Far Away Competitors

Keeping in mind our research compares tax competition within an EIMC with tax competition outside of the EIMC, we believe that estimating equation (3) captures critical localized competition. In particular, this specification compares tax competition among municipalities that share *economic linkages* such as commuting and segmented capital markets that are likely to generate strategic interactions. Nearby towns are the most likely to react *strategically* to each other. Far away towns not linked by commuting or segmented capital markets are atomistic (price-takers) with respect to world capital markets will therefore not interact strategically because they compete à la perfect competition (Wildasin 2014). However, failing to account for tax competition with towns that are far away could bias the coefficient estimates if \bar{d} is misspecified. Here we take an alternative approach to verify robustness to the threshold. Let F superscript the set of municipalities that are far away. We estimate

$$\mathbf{y} = \rho^E \widetilde{\mathbf{W}}^E \mathbf{y} + \rho^O \widetilde{\mathbf{W}}^O \mathbf{y} + \rho^F \widetilde{\mathbf{W}}^F \mathbf{y} + \mathbf{X}\beta + \mathbf{X}^E \gamma + \Gamma + \Theta + \Phi + \varepsilon, \quad (\text{A.6})$$

using the approach of Kelejian and Piras (2014). Define $\widetilde{w}_{ij}^F = 1/n_i^F$ if $d_{ij} \geq \bar{d}$ and $d_{ij} \leq \bar{\bar{d}}$ and zero otherwise, where $\bar{\bar{d}} > \bar{d}$ is a second distance threshold creating another ring of competitors and n_i^F is the number of competitor jurisdictions in the area between \bar{d} and $\bar{\bar{d}}$. We select $\bar{\bar{d}}$ such that a similar number of municipalities have non-zero weights in the $\widetilde{\mathbf{W}}^F$ and $\widetilde{\mathbf{W}}^O$ matrices on average. Critically, the elements of this third weight matrix $\widetilde{\mathbf{W}}^F$ are exogenous given that the municipalities cannot not cooperate with these jurisdictions because they are sufficiently far away.

To instrument, we continue to use the set $\mathbf{H} = [\mathbf{X}, \widehat{\mathbf{W}}^E \mathbf{X}, \widehat{\mathbf{W}}^O \mathbf{X}, \widehat{\mathbf{W}}^E \widehat{\mathbf{W}}^O \mathbf{X}, \widehat{\mathbf{W}}^O \widehat{\mathbf{W}}^E \mathbf{X}]$ but now supplement it with instruments $\mathbf{H}^F = [\widetilde{\mathbf{W}}^F \mathbf{X}, \widetilde{\mathbf{W}}^E \widetilde{\mathbf{W}}^F \mathbf{X}, \widetilde{\mathbf{W}}^O \widetilde{\mathbf{W}}^F \mathbf{X}, \widetilde{\mathbf{W}}^E \widetilde{\mathbf{W}}^O \widetilde{\mathbf{W}}^F \mathbf{X}]$.²⁷

Table A.3 presents the results. First, relative to the results presented in Table 6, the coefficients on the first two spatial lags are similar, although the test of differences between the coefficients is marginally insignificant in some cases. It is expected that standard errors increase when instrumenting for more variables. Thus, the decline in significance is expected given the added endogenous variable in the regressions. This suggests that our results are not influenced by municipalities that are far away. Second, the coefficient on the third spatial lag is often quite small and insignificant. This is consistent with new theoretical models of tax competition (Wildasin 2014), which highlight that spatial interdependence among atomistic jurisdictions in world capital markets should only arise if jurisdictions are linked by labor flows or other economic linkages that lead to market power. As towns in the \mathbf{W}^F matrix are more than 45 kilometers away, it is likely that

²⁷We do not include all possible permutations between $\widehat{\mathbf{W}}^E$, $\widehat{\mathbf{W}}^O$, and $\widetilde{\mathbf{W}}^F$ in the instrument set because computationally, it would result in possible multicollinearity between the instruments.

any labor linkages are relatively small. This null result on the third weight matrix helps justify our distance restriction.

Table A.3: Three Weight Matrices

	1994 Weights		2009 Weights	
	(1)	(2)	(3)	(4)
$\widetilde{\mathbf{W}}^{\mathbf{E}}\mathbf{y}$	0.287*** (0.066)	0.302*** (0.069)	0.274*** (0.064)	0.289*** (0.067)
$\widetilde{\mathbf{W}}^{\mathbf{O}}\mathbf{y}$	0.511*** (0.141)	0.472*** (0.137)	0.626*** (0.159)	0.594*** (0.156)
$\widetilde{\mathbf{W}}^{\mathbf{F}}\mathbf{y}$	0.073 (0.139)	0.113 (0.136)	-0.033 (0.157)	-0.007 (0.155)
Controls	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.161	0.287	0.0597	0.106
Observations	582,722	582,722	582,722	582,722
Clusters	96	96	96	96
F statistic	19563	2840	3600	8374

Unlike table (6), in columns (1)-(4), we estimate (A.6) with three weight matrices. These models use the approach of Kelejian and Piras (2014). Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic and the Hansen J statistic is the p-value. Standard errors are clustered at the department level.

Significance: *90%, ** 95%, ***99%.

A.8 Robustness to Index Creation

Table A.4 shows robustness to various specifications of the weights used to construct the tax index.

Table A.4: KP Results : Robustness to Index Construction

	Town Level				Weights Using EIMC + Town Revenue			
	Unweighted Index (1)	(2)	(3)	(4)	1994 Weights (1')	2009 Weights (2')	2009 Weights (3')	2009 Weights (4')
$\widetilde{\mathbf{W}}^{\mathbf{E}}\mathbf{y}$	0.024 (0.035)	0.065 (0.041)	0.204*** (0.066)	0.261*** (0.069)	0.250*** (0.072)	0.315*** (0.075)	0.230*** (0.067)	0.287*** (0.071)
$\widetilde{\mathbf{W}}^{\mathbf{O}}\mathbf{y}$	0.821*** (0.072)	0.790*** (0.072)	0.759*** (0.061)	0.684*** (0.063)	0.632*** (0.071)	0.565*** (0.069)	0.685*** (0.077)	0.604*** (0.077)
Controls	Y	Y	Y	Y	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.000	0.000	0.000	0.000	0.000	0.010	0.000	0.004
Observations	582,722	582,722	582,722	582,722	582,722	582,722	582,722	582,722
Clusters	96	96	96	96	96	96	96	96
F statistic	51.58	37.52	62.94	45.58	61.46	37.34	83.75	97.33
Hansen J statistic	0.192	0.236	0.371	0.503	0.257	0.334	0.317	0.359

This table presents estimates using the Kelejian and Piras (2014) approach, which resolves endogeneity of the spatial weight matrix and the spatial lag. Columns without a prime use a distance threshold of 45 km, while columns with a prime use a distance threshold of 25 km. All columns estimate (3) following Kelejian and Piras (2014). The tax rate index is constructed using the weights at the top of the table. Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic and the Hansen J statistic is the p-value. Standard errors are clustered at the department level. Significance: *90%, **95%, ***99%.

A.9 As Alternative Estimation Strategy

Finally, while the Kelejian and Piras (2014) approach is appealing for its ability to transparently predict cooperation, we can also implement a (more) standard IV approach. In particular, we only need to follow Kelejian and Piras (2014) if in the second step we want to use instruments $[\widehat{\mathbf{W}^E \mathbf{X}}, \widehat{\mathbf{W}^O \mathbf{X}}, \widehat{\mathbf{W}^E \mathbf{W}^O \mathbf{X}}]$. In particular, Kelejian and Piras (2014) need to instrument for the elements of the weight matrix because they do not have any other exogenous matrices. However, we have an exogenous weight matrix (\mathbf{W}) that does not depend on cooperation. Recall that this matrix is filled with ones if towns are within \bar{d} kilometers and zero otherwise. Thus, by standard assumptions that geographic proximity is not endogenous to tax rates, \mathbf{W} is exogenous. Thus, we use spatially lagged instruments that are related to \mathbf{W} rather than instruments that use the E and O predicted matrices.

The goal is to find instruments that can predict the weight matrices and the tax rates in one step rather than two. Let \mathbf{X} be of dimensionality $NT \times K$ and contain the K exogenous control variables. Let \mathbf{S} be the $NT \times NT$ matrix consisting of the exogenous elements s_{ij} for all pairs of ij as described in Section 3.3, i.e., syndicates created in 1955 onward and earlier, each stacked over each other. These are all exogenous relationships between towns i and j that are correlated with EIMC membership. Then we immediately proceed by instrumenting in the context of our estimating equation rather than predicting the weight matrix. We first instrument $\widetilde{\mathbf{W}}\mathbf{y}$ in estimating equation (1) by the spatially lagged exogenous variables $\widetilde{\mathbf{W}}\mathbf{X}$. Then, the set of instruments for the estimating equation (3) is now $\Xi = [\mathbf{X}, \widetilde{\mathbf{W}}\mathbf{X}, \widetilde{\mathbf{S}}\mathbf{X}]$ where the tildes denote the matrix is row-normalized.

The second panel in Table A.3 presents the results of this simple IV approach. The results are similar in magnitude to table 6. However, as expected due to this procedure not exploiting information about the decision to cooperate, the strength of the instruments becomes weaker although still within accepted values. Given the similarity of the coefficients, but the weaker instruments, this validates our decision to use the Kelejian and Piras (2014) approach. Again, a one percentage point increase in the tax rate of municipalities within the EIMC raises a town's tax rate by 0.322 percentage point; a one percentage point increase in the tax rate of municipalities outside of the EIMC raises that same town's tax rate by 0.597 percentage points. Thus, we still conclude that cooperation dampens the strategic interactions between municipalities.

Table A.5: An Alternative IV Strategy

	1994 Weights		2009 Weights	
	(1)	(2)	(3)	(4)
$\widetilde{\mathbf{W}}^E \mathbf{y}$	0.249*** (0.079)	0.322*** (0.086)	0.230*** (0.068)	0.298*** (0.075)
$\widetilde{\mathbf{W}}^O \mathbf{y}$	0.676*** (0.073)	0.597*** (0.070)	0.707*** (0.075)	0.619*** (0.074)
Controls	Y	Y	Y	Y
Municipal FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
EIMC Regime FE	Y	Y	Y	Y
EIMC Controls	N	Y	N	Y
P-Value: $\rho^E = \rho^O$	0.000	0.006	0.000	0.002
Observations	582,722	582,722	582,722	582,722
Clusters	96	96	96	96
F statistic	6.602	6.054	8.803	5.312
Hansen J statistic	0.0264	0.0210	0.0320	0.0262

These models do not use the approach of Kelejian and Piras (2014) and instead rely on an alternative approach to resolving the endogeneity of cooperation. Even columns control for EIMC characteristics while odd columns do not. To test the equality of $\rho^E = \rho^O$ we present the p-value from a test of the coefficients. The F-statistic is the Kleibergen-Paap rk Wald F statistic and the Hansen J statistic is the p-value. Standard errors are clustered at the department level. Significance: *90%, ** 95%, ***99%.