Consumption Tax Competition Among Governments: Evidence from the United States

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September 2007 (Revised)

Abstract

The paper contributes to a small but growing literature that estimates tax reaction functions of governments competing with other governments. We analyze consumption tax competition between US states, employing a panel of state-level data for 1977–2003. More specifically, we study the impact of a state's spatial characteristics (i.e., its size, geographic position, and border length) on the strategic interaction with its neighbors. For this purpose, we calculate for each state an average effective consumption tax rate, which covers both sales and excise taxes. In addition, we pay attention to dynamics by including lagged dependent variables in the tax reaction function. We find overwhelming evidence for strategic interaction among state governments, but only partial support for the effect of spatial characteristics on tax setting. Tax competition seems to have lessened in the 1990s compared to the early 1980s.

JEL codes: H73, H87, H20, H70, C33

Keywords: Tax competition; tax reaction function; consumption taxation; economic geography

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1 Introduction

US states have the legal power to set their own sales and excise taxes on goods and services. Consequently, sales tax rates and bases differ by state. In 2002, for example, Mississippi levied the highest sales tax rate (7 percent) of all US states. In contrast, Delaware, Montana, New Hampshire, and Oregon did not impose a sales tax at all. Similarly, excise tax rates and bases vary substantially by state. In 2002, New York levied a cigarette excise of US$ 1.50 per pack, whereas Kentucky imposed a rate of only US$ 0.03 per pack. All states levied an excise tax on cigarettes but 19 states did not charge excises on wine. Because commodity tax bases (i.e., the goods and services purchased by individuals) are mobile, states will seek to steal tax base from one another by undercutting their neighbors' consumption tax rates. This may unleash a tax competition game in which states repeatedly interact with each other. Our paper tries to empirically assess whether such strategic interaction exists between US states.

We analyze consumption tax competition among US states, employing a panel data set of state-level consumption taxes (i.e., retail sales taxes on goods and services and excise taxes) for 1977–2003 covering 48 states.[1] To this end, we estimate (reduced-form) tax reaction functions of state governments. A tax reaction function relates the tax rate of the home state to the tax rates of neighboring states and various characteristics of the home state.[2] The slope of the tax reaction function indicates to what degree state government compete with each other.

Consumption tax competition has predominantly been studied from a theoretical point of view.[3] Recently, researchers' attention has shifted from theoretical to empirical work.

[1]We do not cover sales and excise taxes at the local (i.e., county and municipal) level. Federal excises on transportation, communication, energy, alcohol, and tobacco are excluded as well because the focus of our analysis is on horizontal tax competition (i.e., between states) only. See Besley and Rosen (1998) and Devereux et al. (2007) for a empirical model incorporating both horizontal tax competition and vertical tax competition (i.e., between states and the federal level).

[2]See Breuckner (2003) for an overview of the literature on tax reaction functions.

[3]Key contributions are those of Mintz and Tulkens (1986), Kanbur and Keen (1993), Lockwood (1993),

Prior contributions are small in number and focus primarily on the United States.[4] All studies employ the concept of a linear tax reaction function. Estimated slopes of the tax reaction function vary substantially. Some studies find counterintuitive negative slopes for sales taxes (cf. Rork, 2003), whereas others find values close to 0.9 for excises (cf. Egger et al. 2005b). The latter suggests a substantial degree of interaction in tax setting, almost one for one. On average, across all studies, the tax reaction coefficient is 0.5.

Our paper contributes to the literature in three ways. First, our study employs an average "effective" tax rate (AETR) as measure of the tax burden.[5] The AETR on consumption is defined as the ratio of the sum of sales tax and excise tax revenues to total consumption. Such a measure reflects the overall effective tax burden on consumption and should therefore be preferred over studies based on nominal (or statutory) sales tax rates only. Studies on commodity tax competition use either statutory sales tax rates (e.g., Rork, 2003; and Luna, 2004) or statutory (specific) excise tax rates (e.g., Nelson, 2002; Egger et al., 2005b; and Devereux et al., 2007).[6] The study by Egger et al. (2005a), using data for OECD countries, is a notable exception because they are the only ones analyzing AETRs. In the context of the United States, studies have not employed AETRs yet, reflecting the absence of official statistics on consumption at the state level. In this paper, we approximate state consumption on goods and services by non-durable retail sales by state—taken from the *Survey of Buying Power*—and an estimate for durable consumption.

A second contribution is that we explore the effect of a state's spatial characteristics (i.e., its size, geographic position, and border length) on tax setting. Spatial effects are

Trandel (1994), Haufler (1996), Ohsawa (1999), Wang (1999), Nielsen (2001, 2002), Ohsawa (2003, 2004), and Ohsawa and Koshizuka (2003). Wilson (1999) provides an overview of the tax competition literature.

[4]Empirical studies on consumption tax competition in the United States are: Nelson (2002), Rork (2003), Luna (2004), Egger et al. (2005b), and Devereux et al. (2007). Evers et al. (2004) focus on diesel excise competition in Europe. Egger et al. (2005a) deal with tax competition among OECD countries.

[5]The AETR is thus an implicit consumption tax. See Mendoza et al. (1994) for a further exposition on the concept of AETRs.

[6]Devereux et al. (2007) correct statutory excises (defined in specific form) for inflation to arrive at a *real* tax rate. Note that the definition of an AETR implies that we do not have to worry about inflation correction.

taken into account in the regression equation in two ways. We employ three different weighting schemes in characterizing the weighted average of AETRs of competing jurisdictions. We expect our estimate of the tax reaction coefficient (i.e., the slope of the tax reaction function) to be sensitive to the ex ante imposed spatial structure. In addition, we explicitly model (as separate variables in the equation) both time-variant and time-invariant spatial characteristics, which may affect the intercept of the tax reaction function.

Our third contribution is the explicit acknowledgement of the possibility of dynamics in the tax reaction function. If states react to each others' tax setting, the weighted average of competitors tax rates (which we use as an explanatory variable) is endogenous. The literature addresses this by employing an instrumental variable (IV) approach, typically also including state-specific fixed effects and time-specific fixed effects. We show that results obtained in this framework suffer from serial correlation in the disturbances. It cannot be dealt with by including an instrumented lagged dependent variable in the "levels" specification (as proposed by Devereux et al., 2007) because of the correlation between the error term and the lagged dependent variable caused by the presence of state-specific fixed effects. To address this problem, we apply the Arellano-Bond (1991) Dynamic Panel Data (DPD) estimator to the tax reaction function written in "first differences."

We find overwhelming evidence of strategic interaction among state governments. The tax interaction coefficient in the static or "levels" specification (which does not correct for autocorrelation) is sensitive to the type of weighting scheme chosen. It yields tax interaction coefficients in the range [0.49, 0.65], where the upper bound is obtained if competitors tax rates are weighted by contiguity and the lower bound results if population density weights are employed. By applying the DPD estimator to the dynamic (or first differenced) specification, we find tax reaction coefficients in the range [0.38, 0.41], which are much smaller than those for the static model. Any time-invariant spatial characteristics are dropped from the dynamic equation. The static model yields mixed evidence on the effect of state size (as measured by population) on tax setting, whereas state size is not significant in the dynamic specification. Finally, our results indicate that strategic interaction has lessened in the 1990s compared to the early 1980s, suggesting an absence of a "race to the bottom" in AETRs on consumption.

The paper is organized as follows. Section 2 provides a theoretical background to consumption tax competition. Section 3 sets out the methodological framework and discusses identification issues. Section 4 presents data on tax rate changes. Section 5 discusses the empirical results and performs a simple sensitivity analysis. Finally, Section 6 concludes.

2 Hypotheses

Our analysis builds on the theoretical tax competition literature, in which the strategic interaction among governments in tax setting is analyzed. The classic reference in the analysis of origin-based commodity tax competition is Kanbur and Keen (1993), who employ a simple cross-border shopping model, featuring two jurisdictions of fixed areal size. Kanbur and Keen consider a uniformly distributed population, which differs in size across jurisdictions. Households buy one unit of a commodity, which has a fixed producer price (assumed to be the same in both jurisdictions). A commodity's retail price in jurisdiction i consists of the sum of a specific consumption tax, τ_i, and the producer price. The representative household faces fixed transaction costs per unit of traveled distance if it purchases goods across the border. No travel costs are incurred if the consumer purchases goods locally. It follows that the consumer's decision to cross-border shop depends on a comparison between the transactions costs incurred in purchasing the goods in the other jurisdiction and the consumption taxes saved in doing so.

Both governments are assumed to set their consumption tax rates to maximize revenue, while taking as given the tax rate set by the other jurisdiction. This yields a tax reaction function of the general form: $\tau_i = f(\tau_j; \mathbf{V}_i)$, where $\mathbf{V}_i$ is a vector of characteristics of state i (e.g., state size) and f is a linear function (with $f' > 0$).[7] The tax reaction functions for the two jurisdictions can be solved to yield closed-form solutions for the optimal (Nash) tax rates. Equilibrium tax rates are shown to be below the social optimum—reflecting the effect of tax competition—and to be asymmetric (see below).

Ohsawa (1999) extends Kanbur and Keen's model to a multi-jurisdictional setting in which countries differ in areal size and consumers are uniformly distributed across markets.[8] He verifies the robustness of Kanbur and Keen's results to a larger number of jurisdictions. In turn, Ohsawa and Koshizuka (2003) investigate commodity tax competition between two jurisdictions in a two-dimensional setting, that is, including jurisdictional size and jurisdictional shape (e.g., border curvature and border length). In addition to showing that spatial characteristics matter, Ohsawa and Koshizuka (2003) demonstrate that the results obtained by Kanbur and Keen (1993) and Ohsawa (1999) are still valid. The above mentioned papers lead to the following three hypotheses, which we will employ in our empirical analysis.[9]

Kanbur and Keen (1993) show that strategic interaction in tax rate setting results in upward-sloping tax reaction functions (Hypothesis 1). Obviously, the "knife-edge" case of a zero slope is of little practical interest because it implies that interaction between (local) governments is absent.

Hypothesis 1 (Kanbur and Keen, 1993) *A jurisdiction's consumption tax rate is positively related to that of its neighbors.*

[7] In fact, Kanbur and Keen (1993) employ specific functional forms to show that the tax reaction functions are piecewise linear and upward sloping (featuring a slope between zero and unity). Many tax competition models based on general functional forms (see Brueckner, 2003) do not yield sign restrictions.

[8] In Ohsawa's model population density is constant across countries, whereas in Kanbur and Keen's world countries differ in population density.

[9] In view of the well developed existing theoretical frameworks, we have chosen not to develop our own analytical model.

Jurisdictional size plays a key role in consumption tax rate setting. Relatively small jurisdictions have a smaller intercept of the tax reaction function than large jurisdictions [Hypothesis 2(a)]. By undercutting the tax of its large neighbor, a small jurisdiction attracts cross-border shoppers (and thus generates extra revenue at a given tax rate), which exceeds the revenue loss from a lower tax rate applied to the consumption by its residents. For a large jurisdiction, however, the revenue loss on the domestic tax base exceeds the revenue gain from cross-border shoppers. Ohsawa (1999) hypothesizes that the tax rate of the home state rises with the size of the neighboring jurisdictions. Nielsen (2001) shows that this relationship is not clear-cut; when the size of the neighboring state grows, the size of the fiscal externality of a tax rate change rises, causing the home country to decrease its tax rate [Hypothesis 2(b)].

Hypothesis 2 (Kanbur and Keen, 1993; and Nielsen, 2001) *(a) Small home jurisdictions tend to set lower equilibrium consumption tax rates than large jurisdictions; and (b) The consumption tax rate of the home jurisdiction is strictly decreasing in the jurisdictional size of its competitors.*

Spatial characteristics of jurisdictions affect tax setting as is demonstrated by Ohsawa and Koshizuka (2003). Peripheral jurisdictions—of which (part of) their border is not exposed to cross-border shopping—set higher tax rates [Hypothesis 3(a)]. For example, Florida features a large unexposed border on the Atlantic Ocean and the Mexican gulf and is therefore expected to set higher tax rates on consumption. For a given jurisdiction size, a more curved border or an increase in border length means a larger area exposed to cross-border shopping, giving rise to a higher competitive pressure from neighboring jurisdictions. Consequently, exposed jurisdictions set lower tax rates [Hypothesis 3(b-c)].

Hypothesis 3 (Ohsawa and Koshizuka, 2003) *(a) For equally sized jurisdictions, consumption tax rates in peripheral jurisdictions are significantly higher than those in juris-*

dictions situated in the center of a federal country; (b) The consumption tax rate of a jurisdiction decreases if its border becomes more curved; and (c) The consumption tax rate of a jurisdiction decreases if its border length increases.

3 Methodology

This section estimates tax reaction functions specified in reduced form. To measure empirically strategic interaction among local governments, we need to address the issue of identification. In other words, do our results point to strategic interaction or is there some other cause (e.g., common shocks to a state's tax policy)? After a brief discussion of identification, this section describes the econometric specification of the tax reaction function, presents various weighting matrices, and discusses some econometric issues.

3.1 Identification in the Endogenous Interactions Model

Manski (1993) shows that the parameters in models of social/spatial interaction, the class to which tax competition belongs, are only identified under some strict assumptions. He defines three types of interaction: (i) contextual effects (related to exogenous characteristics of the group); (ii) endogenous effects (i.e., the interaction between the units in the group); and (iii) correlated effects (i.e., characteristics that the units have in common, making them behave similarly). The challenge is to disentangle these three effects econometrically in a single equation.

To formally illustrate this, consider the following general cross-sectional model for a given time period:

$$Y_i = \alpha + \delta \mathrm{E}(Y_i|\mathbf{X}_i) + \mathbf{X}_i'\beta + \mathrm{E}(\mathbf{X}_i|\mathbf{Z}_i)'\kappa + u_i, \qquad i = 1, ..., N, \tag{1}$$

where Y_i is the dependent variable (in our case the tax rate), $\mathbf{Z}_i$ is a vector of exogenous characteristics of the group (where boldface characters denote vectors), $\mathbf{X}_i$ are the observed characteristics of the units, E is the expectations operator, and N denotes the number of cross-sectional units. The parameters to be estimated are α, δ, β, and κ. The unobserved characteristics of individuals are included in u_i and are assumed to be correlated across the individuals in the group, that is, $\mathrm{E}(u|\mathbf{X}_i, \mathbf{Z}_i) = \mathbf{Z}_i'\eta$. This implies that the expected value of Y_i given the observed variables $\mathbf{X}_i$ and $\mathbf{Z}_i$ is given by:

$$\mathrm{E}(Y_i|\mathbf{X}_i, \mathbf{Z}_i) = \alpha + \delta\mathrm{E}(Y_i|\mathbf{X}_i) + \mathbf{X}_i'\beta + \mathrm{E}(\mathbf{X}_i|\mathbf{Z}_i)'\kappa + \mathbf{Z}_i'\eta. \tag{2}$$

In this equation, the endogenous effect is measured by the parameter δ, the contextual effect by κ, and the correlated effect by η. The reduced form of this model:

$$\mathrm{E}(Y_i|\mathbf{X}_i, \mathbf{Z}_i) = \alpha/(1-\delta) + \mathrm{E}(\mathbf{X}_i|\mathbf{Z}_i)'(\kappa+\beta)/(1-\delta) + \mathbf{Z}_i'\eta/(1-\delta), \qquad \delta \neq 1, \tag{3}$$

shows that the different social effects cannot be identified separately without imposing further restrictions.

As a first step in solving the specified identification problem, we can consider some of the practical restrictions imposed by the tax competition literature.[10] In general, the literature ignores the interaction effect between the observed group characteristics and the observed individual characteristics and thus assumes implicitly that $\kappa = 0$. This leaves use with the identification of the endogenous effect, δ, and the correlated effect, η, which is infeasible because both the conditional mean, $\mathrm{E}(Y_i|\mathbf{X}_i)$, and the exogenous group characteristics, $\mathbf{Z}_i'$, are constant over the cross-sectional units. The spatial econometrics literature address this issue by replacing $\mathrm{E}(Y_i|\mathbf{X}_i)$ with $\mathbf{W}Y_i$, where $\mathbf{W}$ is a $N \times N$ matrix of exogenously given spatial weights; $\mathbf{W}Y_i$ is thus a weighted average of the dependent variable in other (neighboring) jurisdictions. The identification problem is solved because

[10] As Revelli (2005) points out there is also a second identification issue that plagues the empirical tax competition literature more generally. Based on a reduced-form equation such as (3), we are not able to discriminate between alternative theories of local government interaction (e.g., tax competition, yardstick competition, and expenditure spillovers). We will not address this in the paper because it requires estimating a structural model.

the weighted average of neighbors introduces some cross-sectional variation in $\mathbf{W}Y_i$, as not all jurisdictions in the sample are treated identically, while $\mathbf{Z}_i$ remains constant.

3.2 Econometric Specification

The econometric specification of the theoretical tax reaction function explicitly takes into account the spatial pattern of tax competition. We employ a panel data set so that we can control for unobserved heterogeneity and study the dynamics of tax competition.

Tax Reaction Function

The AETR of state $i = 1, ..., N$ at time $t = 1, ..., T$ is denoted by $\bar{\tau}_{it}$, where N denotes the number of states and T represents the number of time periods. Now using the two assumptions introduced in Section 3.1, that is, assuming $\kappa = 0$ and replacing the conditional mean with the weighted average of the dependent variable in other (neighboring) jurisdictions, the tax reaction function of state i can be written as:

$$\bar{\tau}_{it} = \alpha_i + \eta_t + \delta \sum_{j=1}^{N} w_{ij} \bar{\tau}_{jt} + \mathbf{Q}'_{it}\gamma + \mathbf{X}'_{it}\beta + \varepsilon_{it}, \tag{4}$$

where α_i is a state-specific fixed effect, η_t denotes the year-specific fixed effect, δ is the slope parameter, $\mathbf{Q}_{it}$ and $\mathbf{X}_{it}$ denote vectors of variables representing spatial and demographic characteristics of states and various control variables, respectively, with γ's and β's as parameters. Notice that the correlated effect from the social interactions model of Section 3.1 implies a fixed time effect in a panel data model, which is measured by η_t. An error term, ε_{it}, completes the function. The tax rate of state i is a function of tax setting by its competitors j, which is represented by the "spatial lag" term, $\sum_{j=1}^{N} w_{ij} \bar{\tau}_{jt}$, where w_{ij} is an element of a prespecified $N \times N$ matrix of spatial weights (denoted by $\mathbf{W}_k$, where $w_{ij} = 0$ for $i = j$, see below). Because the AETR is by definition in the range $[0, 1]$, and thus a bounded outcome score, we take a logistic transformation $\bar{\tau}_{it} \equiv \ln \frac{\tau_{it}}{1-\tau_{it}}$, where τ_{it} is the AETR.[11] The logistic transformation is applied to the AETR variable on both sides of equation (4).

Based on Hypothesis 1, we expect positively sloped reaction functions.[12] To test Hypothesis 2(a), we include the population size of state i (i.e., the home jurisdiction) and expect to find $\gamma_1 > 0$. Given Hypothesis 2(b), we expect the weighted population size of neighboring states (i.e., those other than the home jurisdiction) to yield $\gamma_2 < 0$. Sea-bordered states—for which the dummy variable takes on the value one—are expected to set higher tax rates, that is, $\gamma_3 > 0$ [Hypothesis 3(a)]. Border curvature—defined as the ratio of border length and state size—depresses home tax rates and thus $\gamma_4 < 0$ [Hypothesis 3(b)]. Border exposure, which is measured by the population density along the border region of states i and j, has a depressing effect on home tax rates (i.e., $\gamma_5 < 0$).

Our static specification includes year-specific fixed effects and state-specific fixed effects. We include time effects to capture shocks that affect all states simultaneously, for example, a rise in the world oil price. The time effect also picks up changes in federal excise taxes, which we have not explicitly modeled. State-specific fixed effects—which are time invariant—are incorporated to control for unobserved heterogeneity across states as well as observed historical differences. Intuitively, some states (e.g., Delaware, Montana, New Hampshire, and Oregon) oppose any sales taxation.

Weight Matrices

The weighting matrix reflects the degree to which other states influence a given state's tax setting behavior. Defining a weighting matrix is a standard practice in the spatial econometrics literature not only for identification purposes (see Section 3.1), but also for

[11]The logistic transformation was originally suggested by Johnson (1949) to analyze bounded outcome scores.

[12]Kanbur and Keen's (1993) analytical model yields $0 < \delta < 1$. In addition, the empirical literature also puts bounds on δ. Stationarity in the spatial lag model requires that $1/\omega_L < \delta < 1/\omega_U$, where ω_L (ω_U) denotes the smallest (largest) characteristic root of $\mathbf{W}_k$. Note that the largest characteristic root is indeed one if the spatial weights are row-normalized, that is, the rows add up to unity.

reducing the large number of parameters that otherwise need to be estimated. The literature does not give much formal guidance on the choice of appropriate weight matrix. Most often (fixed) geographic criteria are used, which yield purely exogenous weights. We apply three different specifications of weight matrices all of which relate to neighboring states. The first matrix—which has been used before by Egger et al. (2005a)—is constructed using the contiguity of states, that is, whether they share a common border. The elements of the neighboring states matrix, $\mathbf{W}_C$, are:

$$w_{ij} \equiv \begin{cases} b_{ij}/\sum_{j=1}^{N} b_{ij} > 0 & \text{for } i \neq j \\ 0 & \text{for } i = j \end{cases}, \tag{5}$$

where b_{ij} is a border dummy which equals one when states i and $j = 1, ..., N$ share a common border and zero otherwise. Diagonal elements are by definition zero. Because rows are normalized, the spatial lag represents a weighted average of tax rates.[13]

The previous weight matrix treats neighboring states with long borders—and thus providing more opportunities for cross-border shopping—in the same manner as states with short borders. Therefore, we also experiment with a second weighting scheme, which takes into account the length of the border between states i and j. The typical element of the border length matrix, $\mathbf{W}_B$, is:

$$w_{ij} \equiv \begin{cases} l_{ij}/\sum_{j=1}^{N} l_{ij} > 0 & \text{for } i \neq j \\ 0 & \text{for } i = j \end{cases}, \tag{6}$$

where l is the length (in miles) of the common border between states i and j. States with long borders, however, are not necessarily those featuring the largest number of cross-border shoppers. The incidence of cross-border shopping also depends on the population

[13]To reflect a gravity type of approach, Egger et al. (2005a) employ the inverse of the squared distance between two states as a weighting matrix that multiplies the tax rates of neighboring jurisdictions. In contrast to weight matrices based on neighboring states, the distance scheme captures tax competition among *all* states. The elements of a typical distance matrix, $\mathbf{W}_D$, are $w_{ij} = (1/d_{ij}^2)\sum_{j=1}^{N} 1/d_{ij}^2 > 0$ for $i \neq j$ and $w_{ij} = 0$ for $i = j$, where d_{ij} reflects the geographical distance between the largest cities of states i and j. Weighting all states gives rise to tax reaction coefficients close to unity, which are unrealistically high and close to the stationarity bound mentioned in footnote 12. We therefore do not pursue this approach further.

density along the state border, which the final weighting scheme intends to capture. We calculate the population along the border as $s_{ij} \equiv P_{ij} + P_{ji}$, where P_{ij} is the population in all counties in state i adjacent to the common border of states i and j and P_{ji} denotes the population in all counties in state j adjacent to the common border of states i and j. The elements of the population density matrix, $\mathbf{W}_P$, are:

$$w_{ij} \equiv \begin{cases} s_{ij}/\sum_{j=1}^{N} s_{ij} > 0 & \text{for } i \neq j \\ 0 & \text{for } i = j \end{cases} . \tag{7}$$

We take population data at the county level for the year 2000 and assume that the weights remain constant over time.

Control Variables

The control variables can be classified into three broad categories: fiscal, political, and business cycle variables. The first category measures the effect of differences in fiscal policies across states. Two measures are used. The first is per capita public expenditure, lagged one period. Intuitively, as public expenditure rises, the state needs more revenue to balance its budget, providing an incentive to raise consumption tax rates.[14] Second, we use the lagged tax structure, which is defined as the ratio of direct tax revenue to indirect tax revenue. States with a higher tax ratio are expected to levy lower consumption taxes.

In keeping with Egger et al. (2005a) and Devereux et al. (2007), we include a variable representing a state's political orientation, which gets the value one in a year the governor of a state is a Democrat and a zero otherwise. We hypothesize that Republican states prefer a smaller size of the public sector—and therefore are less likely to set high tax rates—than Democratic states (cf. Reed, 2006). The unemployment rate is used to measure the impact

[14] The majority of states are required to balance their budget at the end of the fiscal year (28 in our sample) and some (seven in our sample) require a balanced budget over a two-year cycle. In addition, 36 states have debt restrictions of which 14 require a popular vote to issue any debt. See Table 3 of Poterba and Rueben (2001).

of the business cycle on tax setting behavior of governments. It picks up two opposing effects. On the one hand, in an economic downturn state governments are less inclined to raise tax rates, which suggests a negative effect on tax rates. On the other hand, the unemployment rate captures the effect of automatic stabilizers.[15] A higher unemployment rate leads to more social security outlays, which suggests a positive effect on tax rates. It is not a priori clear which force dominates; the unemployment rate can therefore have either sign.

Econometric Issues

Equation (4) shows that the consumption tax rates of competitors enter contemporaneously (i.e., $\bar{\tau}_i$ depends on $\bar{\tau}_j$ in the same time period), implying that we have to control for endogeneity. In that case, ordinary least squares (OLS) estimation will be inconsistent, reflecting correlation between $\bar{\tau}_{it}$ and ε_{it}. We therefore resort to the IV approach, which yields consistent estimates even in the case of spatial error dependence.[16] Following Kelejian and Prucha (1998) and Kelejian and Robinson (1993), a mix of explanatory variables and weighted explanatory variables is used as instruments. More specifically, the weighted AETRs of neighboring states are instrumented with the weighted unemployment rate (lagged one period) and the weighted per capita public expenditure (also lagged one period). The matrix $\mathbf{W}_k$ defines the weights. All the other (unweighted) predetermined explanatory variables are also included in the instrument matrix.

[15] Note that we find a small correlation coefficient (i.e., -0.37) between the unemployment rate and per capita public expenditure.

[16] Spatial error dependence implies that the error components of jurisdiction i are correlated with those of jurisdiction j. The Moran I test statistic (which is not reported) provides evidence of spatial correlation for all three weighting schemes. Ignoring spatial error dependency may give rise to false evidence of strategic interaction. Kapoor et al. (2007) have developed a three-step procedure to estimate a "spatial error" (also known as spatial autocorrelation) panel data model. This procedure puts additional structure on the unobserved spatial component in the error term. See the Appendix for a description.

3.3 Dynamics

Typically, dynamics are neglected in the estimation of tax reaction functions. A notable exception is Devereux et al. (2007), who deal with serial correlation in the error term by including a lagged dependent variable in their model.[17] Because the lagged dependent variable correlates with the state fixed effect, they instrument it by including the second lag of the dependent variable. This instrument, however, still correlates with the error term (including the fixed effects) and thus invalidates the results. An ideal instrument would have been the state deficit-to-GDP ratio if it were not subject to legal and political restrictions (see footnote 14). We cannot think of any other candidate instruments and therefore adopt an alternative approach.

We include a lagged dependent variable in the tax reaction function of equation (4):

$$\bar{\tau}_{it} = \alpha_i + \lambda\bar{\tau}_{i,t-1} + \delta\sum_{j=1}^{N} w_{ij}\bar{\tau}_{it} + \gamma'\mathbf{Q}_{it} + \beta'\mathbf{X}_{it} + \varepsilon_{it}, \tag{8}$$

where λ is the coefficient of the lagged dependent variable, which captures dynamics. Subsequently, we use the Arellano-Bond (1991) DPD estimator, which is a General Method of Moments (GMM) estimator correcting for endogeneity by including lags of the dependent and explanatory variables (see below). The model is first differenced, implying that any (unobserved) state fixed effects as well as (observed) time-invariant variables are excluded. By applying the first differencing operation to (8), we obtain:

$$\tilde{\bar{\tau}}_{it} = \lambda\tilde{\bar{\tau}}_{i,t-1} + \delta\sum_{j=1}^{N} w_{ij}\tilde{\bar{\tau}}_{it} + \tilde{\mathbf{Q}}_{it}'\gamma + \tilde{\mathbf{X}}_{it}'\beta + \tilde{\varepsilon}_{it}, \tag{9}$$

where $\tilde{r}_{it} \equiv r_{it} - r_{i,t-1}$ for $r \in \{\bar{\tau}, \mathbf{Q}, \mathbf{X}, \varepsilon\}$. It is important to recognize that the coefficients λ, δ, γ, and β are still identified in the first differenced model and have the same interpretation as in the levels model. When estimating this model, the use of the DPD solves the endogeneity problem by instrumenting both the time-lag of the dependent variable and the

[17] The presence of heteroscedasticity can be easily dealt with by employing White standard errors, which does not require a modification of the empirical framework.

weighted tax rates of neighboring states. For instrumenting the time-lag of the dependent variable, we use the dynamic instruments suggested by Arellano and Bond (1991), that is, higher-order lags (starting at $t-2$) of the dependent variable in levels.[18] As instruments for the weighted AETRs of neighboring states, we choose per capita public expenditure and the unemployment rate (appropriately weighted by the respective $\mathbf{W}_k$ matrix). It is important to recognize that the GMM method is robust against the distribution of the dependent variable.

Finally, the proposed instruments used in the GMM estimator must be valid, meaning that they are independent of unobserved heterogeneity and the error term. When the number of instruments is greater than the number of included endogenous variables, the validity of the selected instruments can be tested via an overidentifying restrictions test. We employ a Sargan overidentification test,[19] which indicates that our instruments are valid (see Tables 3–5 below).

4 Data

Our (balanced) panel data set covers 48 states over the period 1977–2003. Table A.1 in the Appendix presents the data definitions and sources. We do not include Alaska and Hawaii in our panel because these two states do not share borders with any other states in the United States. In addition, the District of Columbia (DC) is excluded, because of its special characteristics. DC is extremely small in size (68.3 square miles) and is mainly a working district.[20]

[18]See Baltagi (2005, Section 8.2) for details.

[19]The null hypothesis of the Sargan test states that the overidentifying restrictions are valid. The Sargan statistic is χ^2_{n-l} distributed, where n denotes the rank of the instrument matrix and l is the number of estimated coefficients.

[20]People living in DC spend their money in the surrounding states (i.e., Maryland and Virginia), where the majority of shopping malls is located.

4.1 Estimating Average Effective Tax Rates

The AETR is defined as the ratio of consumption tax revenue to (before-tax) consumption expenditures. Official statistics on consumption expenditures by state are not available. Following Ostergaard et al. (2002), we approximate private nondurable consumption expenditures at the state level by state-level data on retail sales of nondurable goods, which are reported in the *Survey of Buying Power* (published in *Sales and Marketing Management*). State-level private spending on durable consumption goods is estimated (see the Appendix). We prefer using AETRs instead of *statutory* sales tax rates as indicator of the tax burden for three compelling reasons. First and foremost, consumers base their consumption decision upon the *total* consumption tax burden on goods. More specifically, the consumer compares the difference in the tax burden between the neighboring state j and that of the own state i with the transaction (i.e., transport and communication) costs of purchasing in state j.[21] Suppose a consumer purchases one unit of a consumption good subject to both an *ad valorem* sales tax, τ_s (measured as a percentage of value), and a *specific* excise tax, τ_e (measured in US dollars per unit). Given that the sales tax on goods and services is paid on an excise-tax inclusive base, we get tax payments (excluding any federal excises) of:

$$T \equiv (p + \tau_e)(1 + \tau_s) - p = \tau_e + p\tau_s + \tau_e\tau_s, \tag{10}$$

where p denotes the sales price exclusive of tax. On excisable commodities (i.e., beer, cigarettes, distilled spirits, gasoline, and wine) the consumer pays *both* excises (the first term on the right-hand side of (10)) and sales tax (the second term), which none of the previous studies takes into account. Various commodities that are typically purchased across borders are subject to excises. The share of excises in total US consumption tax revenue in the year 2002 amounts on the order of 40 percent. To study tax competition,

[21]Federal excises do not play a role in this comparison, but county level sales taxes on goods and services could be important. Unfortunately, we do not have data on the latter.

one can thus not solely focus on one part of the consumption tax category. Equation (10) also shows that the consumer pays "tax-on-tax" (the last term), which is not picked up by measures based on the sum of statutory tax rates. Although small in many cases, the *tax interaction effect* makes a difference for items such as distilled spirits. For example, in the state New Mexico the sales tax rate amounts to 5 percent and the excise on distilled spirits is US$ 6.06 per gallon, yielding a tax-interaction effect of US$ 0.30 per gallon. Second, AETRs include all relevant components of a tax law (such as exemptions) and take into account the degree of tax enforcement, allowing us to compare states with very distinct tax structures and tax enforcement cultures. For example, Montana does not have a sales tax but generates a significant amount of consumption tax revenue (23.6 percent of total revenue in 2001), reflecting excise tax revenue. Third, AETRs change annually, whereas statutory tax rates change less frequently, which is particularly the case for sales tax rates.

4.2 Descriptive Statistics on Tax Rate Setting

The top panel of Table 1 presents statistics describing the number of tax rate changes across states and over time. Not surprisingly, state governments tinker the most with gasoline excises. Indeed, gasoline sales in border regions are known to react strongly to price differentials between states. Excises on cigarettes feature the second highest mean number of changes. The normalized standard deviation[22] of tax rate changes for these two products is the smallest, suggesting that the majority of states cluster around the mean and thus compete heavily. Nebraska adjusts its gasoline excises the most frequent, that is, every other 16 months. New York is the leader in changing its beer, wine, and distilled spirits excises. States change their statutory sales tax rates on average two times during a time span of 26 years, which is smaller than the average for excises (three changes). Some

[22]The standard deviation of the tax rate of a particular state is divided by the mean of the tax rate of that state (known as the coefficient of variation) to arrive at a unit-free statistic, facilitating a comparison across states and tax categories.

states (e.g., Maryland) do not adjust their sales tax rates at all, whereas New Mexico changes its sales tax rate about six times. Increases in effective tax rates are much more common than tax rate reductions. More specifically, our data set reveals that only 17 of 96 changes (18 percent) in sales taxes pertain to tax rate reductions. We find roughly similar evidence for gasoline excises, for which we observe tax rate reductions in 16 percent of the cases. Hence, there is no indication of a race to the bottom in AETRs on consumption.

The center panel of Table 1 shows the mean size of tax changes (in absolute terms). The overall average change in the sales tax rate is very small (on the order of 0.07 percentage points). Once we exclude all observations where tax rates do not change, the average sales tax change is much higher; it amounts to 0.88 percentage points, which is roughly 20 percent of the overall average sales tax rate. Gasoline excises change more frequently and are of smaller size (15 percent of the average rate). The absolute change in the AETR is much larger than that of the sales tax, reflecting the contribution of revenue from excises.

The bottom panel shows that the average statutory sales tax rate in the United States amounts to 5.2 percent in 2002. It thereby exceeds the AETR (4.1 percent), owing to collection losses on sales taxes (reflecting tax evasion, exemptions, and the like) exceeding the additional revenue generated by excises. Average excise tax rates per gallon vary between US$ 0.19 (gasoline) and US$ 3.55 (distilled spirits). Florida sets the highest excises on distilled spirits and wine (US$ 2.25).

Table 2 shows that the average statutory sales tax across state groupings varies between 3.5 percent and 5.3 percent. Middle Atlantic States (New Jersey, New York, and Pennsylvania) have the highest statutory sales tax. The overall average statutory sales tax rate is slightly higher than the AETR, which is not necessarily true for particular state groups. For example, the Pacific Coast States (California, Oregon, and Washington) appear to have a higher AETR, possibly reflecting substantial excise revenue collections. In addition, AETRs are not necessarily more variable than statutory tax rates. In the aggregate, the variability of AETRs is similar to that of statutory sales tax rates. By state grouping the two measures differ, but there is no systematic pattern.

5 Empirical Results

5.1 Static Model

Table 3 shows estimation outcomes of the static tax reaction function (see equation (4)), using the three different weight matrices introduced above. The tax reaction coefficient can be interpreted as a "corrected tax elasticity," reflecting the logistic transformation of the AETR taken on both sides of the equation.[23] For all three weighting matrices, we find a positive slope of the tax reaction function in line with Hypothesis 1. All slope parameters are smaller than one, which ensures stationarity in the spatial lag model. The size of the slope parameter, however, varies with the weight matrix used. The contiguity weight matrix, $\mathbf{W}_C$, produces the highest slope coefficient (i.e., 0.65), whereas the δ of the population density weight matrix, $\mathbf{W}_P$, is lowest (i.e., 0.49). The home state's population size enters the model with a positive sign and the weighted size of neighboring states with a negative sign.[24] Both outcomes are in accordance with Hypotheses 2(a)–(b). The significance of the tax structure and per capita public expenditure, both lagged one period, complete the model. Both coefficients show the expected sign. Lagged unemployment and a state's political orientation did not prove to be significant.

Our IV estimates of the parameters are consistent even in the case of spatial error dependency. Correcting for spatial error dependence does not invalidate our finding of a significant tax interaction coefficient. Table A.2 in the Appendix reports the results. The

[23]The corrected elasticity is defined as $\frac{\partial \bar{\tau}_{it}}{\partial w_{ij} \bar{\tau}_{jt}} = \frac{\partial \hat{\tau}_{it}}{\partial \hat{\tau}_{jt}} \frac{\hat{\tau}_{jt}}{\hat{\tau}_{it}} \frac{1}{w_{ij}} \equiv \delta$, where $\bar{\tau}_{it} \equiv \ln(\hat{\tau}_{it})$ and $\hat{\tau}_{it} \equiv \frac{\tau_{it}}{1-\tau_{it}}$. Note that $\gamma \equiv \frac{1}{\hat{\tau}_{it}} \frac{\partial \hat{\tau}_{it}}{\partial \mathbf{Q}_{it}}$ and $\beta \equiv \frac{1}{\hat{\tau}_{it}} \frac{\partial \hat{\tau}_{it}}{\partial \mathbf{X}_{it}}$ are interpreted as semi-elasticities.

[24]We experimented with different measures of state size (i.e., surface area and labor force), which did not influence our conclusions.

headings of the columns refer to the weighting matrix employed in the spatial error model. In each of the columns, the AETRs and state size of neighboring jurisdictions are weighted by the population density matrix. Comparing the columns of Table A.2 to column (3) of Table 3, it can be seen that δ becomes more significant, reflecting a smaller standard error. More specifically, the error terms across states are negatively correlated (see the negative ρ in the bottom section of the table). Note that we find roughly the same set of significant variables as in Table 3. The tax interaction coefficients in the combined spatial error and spatial lag model differ somewhat in size from those in the spatial lag model, but these differences are not statistically significant. A negative ρ corresponds to an increase in the estimated δ. In the following, we do not pursue the spatial error model any further because it does not affect our conclusions.

To investigate Hypothesis 3, we include several spatial characteristics of states in the empirical tax reaction function where competitors' tax rates are weighted by the population density. Because it measures the density of *potential* cross-border shoppers, the population weighting matrix has the highest intuitive appeal.[25] We drop state fixed-effects from the model to avoid multicollinearity between time-invariant spatial characteristics and state fixed effects. Table 4 reports the outcomes. A direct consequence of replacing state fixed effects by spatial characteristics is a reduction in the adjusted R^2. Apparently, state fixed effects explain a larger share of the variation than the respective spatial variable that is included. Hypothesis 3 seems to hold. All spatial variables entering the tax reaction function separately have a significant impact on the tax rate. However, border curvature does not have the a priori expected negative sign. Border exposure, that is, the density of people living in counties near the state border, has a direct negative impact on the tax rate.

The inclusion of spatial characteristics does not affect the slope of the tax reaction

[25] Experiments with the other two weight matrices, however, yield the same qualitative conclusions.

function much, which stays close to 0.5. However, the parameters of state size and weighted size of neighboring states change sign, and the effect of lagged per capita public expenditure becomes much larger. In contrast to the previous table, lagged unemployment and a state's political orientation play a role. A higher lagged unemployment rate thus seems to push up a state's AETR via higher social security outlays. The political orientation dummy has the ex ante expected sign. These results suggest that most of the variation in the unemployment variable and political orientation dummy is cross-sectional in nature, which is picked up by the fixed effects in the benchmark regression (Table 3).

5.2 Dynamic Model

The static tax reaction function outcomes as presented in Tables 3 and 4 suffer from serial correlation, as can be seen from the Wooldridge (2002, pp. 282-83) serial correlation test for panel data models. Therefore, Table 5 presents estimates of the dynamic tax reaction function [equations (8)–(9)]. Here, we report the usual standard errors (instead of White diagonal standard errors) because they are robust to remaining serial dependency. The lagged dependent variable is highly significant for all specifications of the weighting matrix, with parameter estimates just above 0.5. Do our hypotheses still hold for the dynamic tax reaction function? The slopes of the tax reaction functions are significantly positive, but become less steep compared to the static model. Therefore, Hypothesis 1 is confirmed. The evidence does not support Hypothesis 2, which is not surprising given that the population sizes of states do not change much over time. Notice that, as mentioned before, theoretically the interpretation of the coefficients does not change after a first differencing operation has been applied. A disadvantage of the Arellano-Bond DPD estimator is that time-invariant variables cannot be included explicitly in the model. Therefore, we cannot formally address Hypothesis 3 in this framework.

To investigate whether tax competition has changed over time, we split the sample into two subperiods, that is, 1977–1990 and 1991–2003 (no table is provided). For all weighting matrices, we find that the slope parameter is much larger in the first subperiod compared to the second subperiod. To illustrate, we will focus on the population density weight matrix.[26] The first period, we find a significant slope parameter of 0.72, which exceeds the value of $\delta = 0.38$ based on the complete sample. In the second subperiod, we find a significant slope parameter of 0.20, suggesting a larger degree of tax competition among states in the 1980s than in the 1990s. The drop in transaction costs associated with cross-border shopping—and the potentially larger tax elasticity of the consumption tax base—thus has not resulted in a greater degree of tax competition.

6 Conclusions

This paper measures tax competition between US states, using a panel data set of state-level consumption taxes (i.e., retail sales taxes on goods and services and excise taxes collected by state governments) for the period 1977–2003 covering 48 states. Rather than employing statutory tax rates (as is customary in the literature), we calculate average effective consumption tax rates. We estimate both static and dynamic tax reaction functions, where the the dynamic model corrects for serial correlation in the error term.

We find strong evidence of strategic interaction among US states. The dynamic model yields much smaller estimated tax interaction coefficients than the static model, indicating that the latter overstates the degree of tax interaction between states. Using the preferred dynamic model, we observe a larger degree of strategic interaction during the 1980s than the 1990s. This suggests that the fall in transaction costs of cross-border shopping does not give rise to a race to the bottom in average effective consumption tax rates.

Spatial characteristics can influence the slope as well as the intercept of the tax reac-

[26] The results for the other weighting matrices are available upon request from the authors.

tion function. Contiguity weight matrices yield the largest tax interaction effect in both static and dynamic models. Using the static model, which allows time-invariant spatial characteristics to be modeled, we show that states near the oceans and Mexican Gulf set higher average effective consumption tax rates than inland states. In addition, states with a larger population density along the border—and thus face a larger exposure to cross-border shopping—tax consumption at a lower average effective tax rate than states with less border exposure. We find mixed evidence on the relationship between state size and tax setting.

In future work, we intend to apply the analysis to a broad set of (more heterogeneous) countries, including OECD and non-OECD countries. To date, few empirical studies have examined tax competition among governments of developing countries.

Appendix

Data Sources Table A.1 sets out the variable definitions and data sources. Total retail sales reflects *net* sales (gross sales minus refunds and allowances for returns) for all establishments primarily engaged in retail trade, plus eating and drinking establishments. Receipts from repairs and other services (by retailers) are also included, but retail sales by wholesalers and service establishments are not. Note that sales for some establishments (e.g., lumber yards, paint, glass, and wall-paper stores, and office supply stores) are also included, even if they sell more to businesses than to consumers.

Estimation of Consumption Retail sales data do not include private consumption of durable goods. State-level spending on durable consumption goods is estimated. To this end, we assume a fixed share of private durable consumption goods across states. Aggregate US durable private consumption is approximated by the difference between aggregate US private consumption expenditures and aggregate US retail sales (both measured at market prices). Note that this also includes nondurable private consumption expenditures that are not included in retail sales (e.g., travel expenditures). We focus on private consumption only because we do not have state-level data on goods and services purchased by the government (i.e., total public consumption minus the wage bill). The latter amounts to roughly 5 percent of total goods and services consumption across states.

Combined Spatial Error and Spatial Lag Model Table A.2 estimates equation (4) while correcting for spatial error dependence. To this end, the disturbance term is assumed to be spatially autoregressive:

$$\varepsilon = \rho \mathbf{W}_l \varepsilon + \xi, \tag{A.1}$$

where ρ is the spatial autoregressive coefficient, $\mathbf{W}_l$ (where $l \in \{B, C, P\}$) is a matrix of spatial weights, and ξ is a well-behaved error term. We thus allow for the possibility that $\mathbf{W}_k \neq \mathbf{W}_l$. The estimation procedure consists of three steps. In the first step, an instrumental variable (IV) estimator is used to arrive at an estimate of ε. The second step employs the estimated residuals in a GMM procedure to estimate ρ. Finally, the estimate of ρ is used to transform the model. Applying IV to the transformed model yields the estimates reported in Table A.2.

Table A.1: Variable Definitions and Data Sources

Definition	Sources	Internet location
Inputs for AETR		
Retail sales at the state level (in thousands of US$)	Survey of Buying Power, 1978–2004	www.salesandmarketing.com
Aggregate consumption (in thousands of US$)	IMF's International Financial Statistics	www.ifs.apdi.net/imf
Sales tax revenue at the state level (in thousands of US$)	World Tax Database	www.wtdb.org
Excise tax revenue at the state level (in thousands of US$)	World Tax Database	www.wtdb.org
Inputs for spatial variables at the state level		
Border length of state (in miles)	Thomas J. Holmes's web site	www.econ.umn.edu/~holmes/data/borderdata.html
County population along border (number of individuals)	US Census Bureau	www.census.gov
Population of state (in millions)	Bureau of Economic Analysis	www.bea.gov
Geographic area (in square miles)	US Census Bureau	quickfacts.census.gov
Inputs for control variables at the state level		
Direct tax revenues (in thousands of US$)	World Tax Database	www.wtdb.org/index.html
Indirect tax revenues (in thousands of US$)	World Tax Database	www.wtdb.org/index.html
Public expenditure (in thousands of US$)	World Tax Database	www.wtdb.org/index.html
Unemployment rate (in percent)	Bureau of Labor Statistics	www.bls.gov or www.economagic.com
Party of the Governor (dummy)	Individual states	Web sites of individual states

Table A.2: Static Model With Correction for Spatial Dependency in the Error Term

Weighting matrix:	Contiguity	Border length	Population
Population weighted AETR of neighbors	0.654***	0.705***	0.733***
	(0.168)	(0.175)	(0.177)
Home state's population size	0.010**	0.015**	0.018**
	(0.005)	(0.005)	(0.005)
Weighted state size of neighbors	-0.019**	-0.023**	-0.027**
	(0.009)	(0.010)	(0.009)
Tax structure at $t-1$	-0.178***	-0.181***	-0.169***
	(0.022)	(0.022)	(0.022)
Per capita public expenditure at $t-1$	0.036***	0.034***	0.031***
	(0.015)	(0.015)	(0.014)
Unemployment rate at $t-1$	0.000	0.000	-0.001
	(0.004)	(0.004)	(0.004)
Political orientation dummy	-0.003	-0.003	-0.009
	(0.006)	(0.006)	(0.006)
Adjusted R^2	0.909	0.935	0.945
Observations	1,248	1,248	1,248
Spatial lag parameter of error term (ρ)	-0.429	-0.402	-0.450
Sargan test	0.013	0.006	0.002
	[0.909]	[0.938]	[0.965]
Wooldridge test	2.978***	2.978***	2.978***
	[0.000]	[0.000]	[0.000]

Notes: The dependent variable is the average effective tax rate (AETR) of state i in period t. Both time and state fixed effects are included (but are not reported). The weighted AETR is instrumented with the weighted (lagged) unemployment rate and the weighted (lagged) per capita public expenditure using the population density weighting matrix. The remaining variables are assumed to be exogenous and therefore also included in the instrument matrix. ***, **, * denote significance at the 1, 5 or 10 percent level, respectively. White diagonal standard errors are presented in parentheses below the parameter estimates. Figures between brackets are p-values. Reported values for the Wooldridge serial correlation test are t-statistics.

Table 1: Descriptive Statistics on Statutory and Effective Tax Rates, 1977–2002

	Statutory tax rates						AETR[a]
	Sales tax	Excises (in US$ per unit)					(in percent)
	(in percent)	Beer (per gallon)	Cigarettes (per pack)	Distilled spirits (per gallon)	Gasoline (per gallon)	Wine (per gallon)	
	Changes in tax rates across states and years						
Number of tax rate changes:							
Mean	2.18	1.38	3.38	2.10	6.46	1.93	–
Maximum	6	6	9	9	19	5	–
State(s) with maximum[b]	NM	NY	WA	NY, NM	NE	NY, MO	–
Coefficient of variation[c]	0.69	1.08	0.64	1.13	0.61	0.84	–
Absolute tax rate changes:							
Mean (including zeros)	0.066	0.003	0.016	0.068	0.006	0.012	0.190
Mean (excluding zeros)[d]	0.875	0.053	0.103	0.848	0.023	0.132	0.190
	Tax rates across states in 2002						
Mean	5.19	0.22	0.52	3.55	0.19	0.65[e]	4.08
Maximum	7.00	0.77	1.50	6.50	0.28	2.25	9.93
State(s) with maximum[b]	MS	SC	NY	FL	RI	FL	WA
Coefficient of variation[c]	0.19	0.68	0.79	0.38	0.27	0.72	0.36
Number of states[e]	44	48	48	30	48	29	48

Sources: Office of Tax Policy Research, *World Tax Database*; and authors' own calculations.

[a] The AETR denotes the average effective consumption tax rate.

[b] The state labels are follows: Florida (FL), Mississippi (MS), Missouri (M0), Nebraska (NE), New Mexico (NM), New York (NY), Rhode Island (RI), South Carolina (SC), and Washington (WA).

[c] The coefficient of variation (defined as the standard deviation divided by the mean) measures the average variation of the tax rate.

[d] The sample sizes per tax category differ because of the elimination of the zeros.

[e] Alaska and Hawaii are excluded because these states do not share a common border with other US states. The numbers per tax category may differ because not all states have a sales tax (e.g., Delaware, Montana, New Hampshire, and Oregon) or excise tax.

Table 2: Average Statutory and Effective Tax Rates by Region, 1977–2002

Region[a]	Average statutory sales tax rate		AETR[b]	
	Average	Variation[c]	Average	Variation[c]
Middle Atlantic states	5.28	0.031	4.09	0.084
Midwestern states	4.57	0.149	3.30	0.092
New England states	4.64	0.078	3.74	0.124
Pacific Coast states	3.79	0.082	4.13	0.098
Rocky Mountain states	3.52	0.114	3.77	0.141
Southern states	4.22	0.106	4.45	0.105
Southwestern states	4.53	0.190	4.46	0.119
Average	4.36	0.107	4.06	0.109

Sources: Office of Tax Policy Research, *World Tax Database*; and authors' own calculations.

[a] The grouping of states is as follows: *Middle Atlantic States* (New Jersey, New York, and Pennsylvania), *Midwestern States* (Illinois, Indiana, Iowa, Kansas, Michigan, Minnesota, Missouri, Nebraska, North Dakota, Ohio, South Dakota, and Wisconsin), *New England States* (Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont), *Pacific Coast States* (California, Oregon, and Washington), *Rocky Mountain States* (Colorado, Idaho, Montana, Nevada, Utah, and Wyoming), *Southern States* (Alabama, Arkansas, Delaware, Florida, Georgia, Kentucky, Louisiana, Maryland, Mississippi, North Carolina, South Carolina, Tennessee, Virginia, and West Virginia), and *Southwestern States* (Arizona, New Mexico, Oklahoma, and Texas). Alaska and Hawaii are excluded, yielding a total of 48 states.

[b] The AETR denotes the average effective consumption tax rate.

[c] The coefficient of variation (defined as the mean divided by the standard deviation) measures the average variation of the tax rate in the specific region.

Table 3: Static Model With Both State and Time Fixed Effects

Weighting matrix:	Contiguity	Border length	Population
Weighted AETR of neighbors	0.649***	0.624***	0.492***
	(0.218)	(0.208)	(0.198)
Home state's population size	0.018***	0.015***	0.013**
	(0.005)	(0.005)	(0.005)
Weighted state size of neighbors	-0.041***	-0.023***	-0.019*
	(0.010)	(0.009)	(0.009)
Tax structure at $t-1$	-0.213***	-0.201***	-0.210***
	(0.024)	(0.024)	(0.024)
Per capita public expenditure at $t-1$	0.039**	0.043***	0.046**
	(0.017)	(0.016)	(0.016)
Unemployment rate at $t-1$	-0.000	-0.000	0.001
	(0.004)	(0.004)	(0.004)
Political orientation dummy	-0.002	0.001	0.002
	(0.007)	(0.006)	(0.006)
Adjusted R^2	0.942	0.941	0.944
Observations	1,248	1,248	1,248
Sargan test	0.003	0.078	1.249
	[0.959]	[0.606]	[0.264]
Wooldridge test	20.538***	20.188***	23.027***
	[0.000]	[0.000]	[0.000]

Notes: The dependent variable is the average effective tax rate (AETR) of state i in period t. Period and state fixed effects are included (but are not reported). The weighted AETR is instrumented with the weighted (lagged) unemployment rate and the weighted (lagged) per capita public expenditure using the respective weight matrix reported in the column heading. The remaining explanatory variables are assumed to be exogenous and therefore also included in the instrument matrix. ***, **, * denote significance at the 1, 5 or 10 percent level, respectively. White diagonal standard errors are reported in parentheses below the parameter estimates. Figures between brackets are p-values. Reported values for the Wooldridge serial correlation test are t-statistics.

Table 4: Static Model with Time Fixed Effects and Various Spatial Characteristics

Spatial characteristics:	Sea bordered	Border curvature	Border length
Population weighted AETR of neighbors	0.416***	0.434***	0.483***
	(0.096)	(0.097)	(0.114)
Home state's population size	-0.006***	-0.006***	-0.007***
	(0.002)	(0.002)	(0.002)
Weighted state size of neighbors	0.007**	0.006**	0.004
	(0.003)	(0.003)	(0.003)
Tax structure $t-1$	-0.342***	-0.340***	-0.340***
	(0.011)	(0.011)	(0.011)
Per capita public expenditure $t-1$	0.298***	0.335***	0.316***
	(0.025)	(0.028)	(0.027)
Unemployment rate $t-1$	0.031***	0.030***	0.030***
	(0.005)	(0.005)	(0.005)
Political orientation dummy	0.042***	0.049***	0.053***
	(0.015)	(0.015)	(0.015)
Sea bordered dummy	0.075***	–	–
	(0.015)		
Border curvature	–	1.724***	–
		(0.419)	
Border length	–	–	-0.099***
			(0.027)
Border exposure	-0.033**	-0.063***	-0.046***
	(0.013)	(0.017)	(0.014)
Adjusted R^2	0.700	0.686	0.672
Observations	1,248	1,248	1,248
Sargan test	0.572	0.488	0.343
	[0.450]	[0.485]	[0.558]
Wooldridge test	24.809***	25.230***	23.015***
	[0.000]	[0.000]	[0.000]

Notes: The dependent variable is the average effective tax rate (AETR) of state i in period t. The AETRs of neighboring state are weighted by the population density matrix. Only year fixed effects are included (but are not reported). Following Kelejian and Robinson (1993, p. 302), the weighted AETR is instrumented by the (lagged) unemployment rate (weighted once) and lagged per capita public expenditures (weighted twice), both using the population density matrix. The remaining explanatory variables are considered to be exogenous and therefore also included in the instrument matrix. ***, **, * denote significance at the 1, 5 or 10 percent level, respectively. White diagonal standard errors are reported in parentheses below the parameter estimates. Figures between brackets are p-values. Reported values for the Wooldridge serial correlation test are t-statistics.

Table 5: Dynamic Model Estimated Using Arellano-Bond

Weighting matrix:	Contiguity	Border length	Population
Lagged AETR of home state	0.555***	0.516***	0.544***
	(0.030)	(0.045)	(0.033)
Weighted AETR of neighbors	0.413***	0.405***	0.384***
	(0.033)	(0.045)	(0.047)
Home state size	0.002	0.002	-0.002
	(0.006)	(0.008)	(0.007)
Weighted state size of neighbors	-0.015	-0.015	-0.011
	(0.010)	(0.010)	(0.008)
Tax structure $t-1$	-0.061***	-0.058***	-0.061***
	(0.004)	(0.007)	(0.005)
Per capita public expenditure $t-1$	0.013***	0.012***	0.012***
	(0.005)	(0.005)	(0.004)
Unemployment rate $t-1$	0.005***	0.004***	0.004***
	(0.001)	(0.001)	(0.001)
Political orientation dummy	0.001	0.002	0.005
	(0.010)	(0.010)	(0.010)
Adjusted R^2	0.546	0.537	0.541
Observations	1,200	1,200	1,200
Sargan test	41.459	39.243	39.956
	[0.407]	[0.504]	[0.472]

Notes: The dependent variable is the first differenced average effective tax rate (△AETR) of state i in period t. State fixed effects are included (but are not reported). The weighted AETR is instrumented by the lagged unemployment rate and the lagged per capita public expenditure (both weighted by the population density matrix). The remaining explanatory variables are considered to be exogenous and therefore also included in the instrument matrix. The time-lag of the dependent variable is instrumented with higher-order lags (starting at $t-2$) of the dependent variable and lags of the other explanatory variables in levels. ***, **, * denote significance at the 1, 5 or 10 percent level, respectively. White period standard errors are presented in parentheses below the parameter estimates. Note that the standard errors in the dynamic model are robust against remaining serial dependency in the error term. Figures between brackets are p-values.

Acknowledgements

The authors would like to thank Charlotte Ostergaard for providing part of the data on nondurable retail sales, Peter Egger, J. Paul Elhorst, Ben Heijdra, and László Paizs for helpful discussions, and participants at the Research Day of the Netherlands Network of Economics (NAKE), Amsterdam, October 2006, and the 63rd Congress of the IIPF, Warwick, United Kingdom, August 2007.

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