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ABSTRACT: This paper explores for the first time the consequences of centrally imposed local tax limitations on the modelling and estimation of spatial auto-correlation in local fiscal policies, and compares three spatial interaction estimators: a) the conventional maximum likelihood estimator that ignores censoring; b) a spatial Tobit estimator; c) a discrete hazard estimator. Implementation of the above empirical approaches on the case of local vehicle taxation in Italy provides a reasonably coherent picture in terms of the direction and size of the spatial interaction process, and offers a plausible spatial interpretation of the race to the top in provincial vehicle taxes.

JEL Codes: C23, C25, H72

Keywords: vehicle taxation, spatial auto-correlation, censored data

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1 Introduction

The nature of central-local relationships in M-form (multi-divisional form) public sector structures typically plays a crucial role in determining the observed degree of tax and public expenditure decentralization and the actual extent of local fiscal autonomy (Maskin et al. [28]). As documented in Joumard and Kongsrud [23] and Sutherland et al. [41], national governments around the globe exercise their command by imposing lower and/or upper bounds on local tax rates or by mandating types and levels of expenditures on local public services.

While the issue of central mandating and capping on local fiscal decisions has attracted considerable interest in the theoretical and empirical public economics literature,[1] little attention has been devoted to their implications on the empirical modelling of spatial dependence in local fiscal policies.

In fact, it has long been recognised that a key feature of decentralised fiscal policy-making is the interdependence among decision-makers due to a variety of spatial transmission mechanisms, with tax competition and yardstick competition motives having somewhat obscured the traditional benefit spill-over hypotheses in recent years (Brueckner [8], Allers and Elhorst [1], Revelli [36]). However, all of the existing empirical analyses of inter-jurisdictional competition rest on the often implausible assumption that local decision-makers are actually free to choose their preferred policy.[2] In most instances, this is not the case due to the existence of central government mandates or caps either on taxes or on public expenditures. The examples around the world are countless.[3]

This paper attempts at exploring for the first time the consequences of central capping on the modelling and estimation of a local fiscal policy reaction

[1] A theoretical formalisation of the genesis of central mandates is Cremer and Palfrey [11], while most of the empirical literature concerns tax and expenditure limitations in the US states (Nechyba [30]; Figlio [16], Downes et al. [12], Dye et al. [14]). Boadway [7] reviews and highlights the key issues in that regard.

[2] In an early study of property tax competition within the Boston metropolitan area, Brueckner and Saavedra [9] highlighted the link between local tax limitations and the intensity of tax competition. They pointed out that "reaction functions become flat once they encounter the levy-limit constraint" (Brueckner and Saavedra [9], p. 220) and acknowledged the difficulty of modelling spatial dependence in a censored dependent variable framework: "implementing this kind of double regime specification in a spatial lag context appears difficult" (Brueckner and Saavedra [9], p. 220). In fact, their analysis focused on the regime switch represented by the introduction of a local tax limitation known as Proposition $2\frac{1}{2}$ and found that the degree of spatial auto-correlation in local property taxes was somewhat lower in the presence of the tax cap.

[3] See Joumard and Kongsrud [23], Emmerson et al. [15], Sutherland et al. [41], Ambrosanio and Bordignon [2], Zodrow [44]. Wolman et al. [43] provide a comprehensive and detailed picture of local tax and expenditure limitations in the US states.

function. In particular, we compare three different estimators of the slope of an intergovernmental fiscal reaction function in the frequently encountered case of central government exercising its command by imposing an upper limit on local fiscal choices: a) the conventional maximum likelihood estimator of a spatial lag dependence specification that does not account for censoring; b) a spatial Tobit estimator that allows for simultaneous spatial dependence; c) a discrete hazard estimator of the probability of a local government hitting the upper censoring point augmented with a dynamic spatial effect.

The paper shows an application of the above empirical approaches to a panel dataset of the (100) Italian provinces. In particular, the empirical analysis focuses on the centrally imposed constraints on the main source of own revenue for Italian provincial governments - the vehicle registration tax - and attempts at identifying the factors that brought about the extraordinary race to the top in the provincial vehicle tax during the 2000s.

Actually, while almost entirely neglected in the empirical public economics literature,[4] local vehicle taxation is widespreadly employed in both developed and developing countries and it is of great interest from a public economics standpoint for at least two reasons. If properly designed, it is one of the most powerful instruments to reduce vehicle-related pollution, and can play a crucial role in internalizing the external costs of road transport in terms of environmental and human health effects.[5] Second, due to the high visibility of vehicle taxes and the widespread ownership of motor vehicles, vehicle taxation can work as a signal of a government's quality and competence, and could therefore foster accountability and yardstick competition between decentralized governments.[6]

In fact, the evidence emerging from the investigation of the Italian Provinces' tax setting behavior is generally consistent with the hypothesis of a process of inter-provincial interaction. Moreover, the three estimation approaches outlined above provide a coherent picture in terms of the direction and size of the spatial auto-correlation process. In particular, the conventional spatial ML (maximum likelihood) approach that does not account for centrally imposed censoring leads to an estimate of the spatial auto-correlation coefficient (0.08) that is remarkably close to the estimate that is obtained with a Bayesian spatial Tobit approach

[4] Mahadi et al. [26], Suter and Walter [40] and Solé Ollé [39] represent exceptions. On the other hand, the empirical literature on the related issue of gasoline taxation is far more developed (Fredriksson and Millimet [19]).

[5] Suter and Walter [40] offer a detailed evaluation of the Swiss experience.

[6] In addition, local vehicle taxation might give rise to tax competition as long as the tax base (motor vehicles) is mobile across jurisdictions.

that explicitly models both the clustering of authorities at the censoring point and the process of simultaneous spatial dependence (0.10). Similarly, the discrete hazard model provides evidence that the probability of an authority hitting the upper censoring point is affected by the (lagged) fiscal choices of neighboring authorities. Due to the strategic timing of vehicle tax increases by Provincial governments, with tax increases being generally implemented far from or right after an election year, we are inclined to attribute the cause of the observed spatial dependence in local fiscal choices to a political information spill-over, by which Provinces that are forced to raise the vehicle tax to cope with their growing spending needs make it less politically costly, and consequently more likely, for other Provinces to raise their taxes too.

The paper is organized as follows. Section **2** illustrates the three empirical approaches for the estimation of a local policy reaction function in the presence of centrally determined censoring. Section **3** turns to the application to the Italian Provinces' vehicle taxes and discusses a number of hypotheses for the growth of the provincial vehicle tax over time. Section **4** reports and discusses the estimation results, and section **5** concludes.

2 Centrally censored local fiscal policies

Following the enormous growth in tax and yardstick competition theoretical research in the past two decades, the econometric analysis of spatial autocorrelation in local governments' fiscal policies has recently surged as one of the most lively areas of research in applied public economics.[7]

Typically, the theory focuses on either of the following two constraints on local policy-makers' ability to raise revenues: the first is represented by the mobility of the tax base giving rise to tax competition; the second consists in the need for politicians to gain imperfectly informed taxpayers' consensus in the presence of cross-jurisdictional information spillovers making relative performance evaluation preferable to absolute performance evaluation (yardstick competition).

However, local governments around the world are hardly ever free to set the fiscal policies they see fit. In fact, they are frequently subject to stringent regulations and caps on their tax and spending decisions, making the ideal paradigm of intergovernmental competition sort of blurred in practice.

[7] See the reviews in Brueckner [8] and Revelli [36].

While the empirical tax competition and yardstick competition literatures have virtually universally ignored those constraints, this paper attempts at exploring for the first time the consequences of central capping on the modelling and estimation of a local fiscal policy reaction function.

In particular, the following two empirical issues should ideally be simultaneously tackled in empirical reserach. First, central command on local fiscal policies generates censoring in the dependent variable. A frequently encountered case is a cap ($\overline{\tau}$) on a local tax rate τ, meaning that $\tau \leq \overline{\tau}$, ideally calling for a corner solution model accounting for clustering at the censoring point. Second, in the presence of interdependence in local fiscal choices, the empirical model needs to properly allow for the simultaneous determination of those choices, with competing authorities exerting a reciprocal influence on each other.

2.1 The spatial lag dependence model

Let us consider first the conventional spatial lag specification of the tax reaction function that takes the vector of observed local tax rates τ as the continuous dependent variable. This model ignores altogether the censoring in the dependent variable ($\tau \leq \overline{\tau}$), and can be expressed as:

$$\tau_{it} = \rho \tau_{-it} + x'_{it}\beta + \varepsilon_{it} \tag{1}$$

where τ_{it} is the tax rate set by jurisdiction i in year t, and ρ (with $-1 < \rho < 1$ to ensure spatial stationarity) is the first-order spatial auto-regressive coefficient relating own tax rates to the spatially weighted average of other jurisdictions' tax rates:

$$\tau_{-it} = \sum_{j=1}^{N} w_{ij}\tau_{jt} \tag{2}$$

where w_{ij} are spatial weights that, according to the conventional binary contiguity criterion, equal $\frac{1}{n_i}$ if jurisdiction j is contiguous to jurisdiction i, and equal 0 otherwise, with n_i being the number of units being adjacent to unit i. Finally, ε_{it} is assumed to be independently and identically distributed across geographical units and over time.[8]

By further assuming that $\varepsilon_{it} \sim N(0, \sigma^2_\varepsilon)$, the spatial lag dependence model (1)-(2) can conveniently be estimated by standard spatial econometric maxi-

[8] In fact, residual spatial autocorrelation, i.e., the possibility of a spatial process in ε_{it}, should be properly tested for before estimating the spatial lag specification (1). See section **4** below.

mum likelihood (ML) techniques (Anselin [3]).[9] In matrix form, the spatial lag dependence model can be inverted and expressed as:

$$\tau = (I - \rho W)^{-1} X\beta + (I - \rho W)^{-1} \varepsilon \tag{3}$$

where I is the $(NT \times NT)$ identity matrix and $W = [I_T \otimes W_N]$ is the block-diagonal, row-standardized spatial weights matrix, with $W_N = \{w_{ij}\}$, $i, j = 1, ..., N$, and $\sum_j w_{ij} = 1$, ∇i.

However, disregarding censoring of the dependent variable leads to the following two problems. First, and similarly to the standard econometric problem that is encountered when the dependent variable is censored (Winkelmann and Boes [42]), not accounting for censoring leads to downward biased estimates of the parameters β. Second, and more importantly for our purposes, it leads to a potentially biased estimate of the crucial first-order spatial auto-regressive coefficient ρ measuring the slope of the reaction function.

In fact, assume that the ultimate objective of the empirical analysis consists in recovering the "true" slope parameter ρ^* in a reaction function where the "desired" - and partly unobserved due to capping - tax rates are spatially auto-correlated:

$$\tau^*_{it} = \rho^* \tau^*_{-it} + x'_{it}\beta + \varepsilon_{it} \tag{4}$$

$$\tau^*_{-it} = \sum_{j=1}^{N} w_{ij} \tau^*_{jt} \tag{5}$$

with the observed tax rate being generated as:

$$\tau_{it} = \begin{cases} \overline{\tau} & \text{if} \quad \tau^*_{it} \geq \overline{\tau} \\ \tau^*_{it} & \quad\;\; \tau^*_{it} < \overline{\tau} \end{cases} \tag{6}$$

In a way, model (1)-(2) is the observed counterpart of the latent model (4)-(5). Since both the own tax (τ_{it}) and the variable representing the average taxes in neighboring jurisdictions (τ_{-it}) are only observed after censoring, though, the direction and size of the bias deriving from estimation of model (1)-(2) instead of the true but unobserved process (4)-(5) are unkown a priori.

[9] Moreover, the spatial lag dependence specification can easily accommodate fixed time and authority effects.

2.2 A Bayesian spatial Tobit approach

With $-1 < \rho^* < 1$, the matrix form of equation (4) can be inverted and expressed as:

$$\tau^* = (I - \rho^* W)^{-1} X\beta + (I - \rho^* W)^{-1} \varepsilon \tag{7}$$

with variance-covariance matrix:

$$\Omega = (I - \rho^* W)^{-1} (I - \rho^* W)^{-1\prime} \sigma_\varepsilon^2 \tag{8}$$

The substantial difference of the latent variable model (7) with respect to a non-spatial specification ($\rho^* = 0$) is that the spatially correlated covariance structure (8) does not allow the simplification of the multivariate distribution into the product of univariate distributions. Moreover, the heteroskedasticity implied by the spatial covariance structure causes inconsistency of standard non-spatial discrete choice estimation methods (McMillen, [29]; Fleming [18]).

A number of approaches have recently been proposed to consistently estimate variants of model (7), particularly with reference to a binary dependent variable setting (spatial Probit), and where spatial dependence typically takes the form of a first-order autoregressive process in the residuals (Pinkse and Slade [35]):

$$\begin{aligned} \tau^* &= X\beta + v && (9)\\ v &= \lambda W v + \varepsilon && (10)\\ \varepsilon &\sim N(0, I) && (11) \end{aligned}$$

where λ, with $-1 < \lambda < 1$, is the auto-regressive coefficient in the spatial error process and W is as defined above.

The proposed estimation methods either focus on the heteroskedasticity induced by the spatial model structure and address it by making specific assumptions on the form of the spatial weights matrix (Case [10]) and the variance-covariance structure (Pinkse and Slade [35]), or make full use of the spatial information and rely on computationally complex techniques (the EM algorithm, simulation methods or Bayesian methods) to tackle the issue of multidimensional integration (Fleming [18]).

Within the latter class of models, the Bayesian spatial discrete choice method developed by LeSage [24] overcomes some drawbacks that arise in the EM algorithm when estimating standard errors (McMillen [29]), and has the advantage of allowing the errors to be heteroskedastic after controlling for spatial dependence. Moreover, it tends to be superior to simulation methods (Beron and Vijverberg [4]) in terms of computational requirements and flexibility (Fleming [18]). Most importantly, though, the LeSage Bayesian approach is the best suited to estimate a censored dependent variable - Tobit - model with simultaneous spatial dependence as in (7) above.

The Bayesian spatial Tobit approach is based on the principle that a likelihood function for model (7) can be formulated and optimized based on estimates of the unobserved latent variable τ^*.[10] In practice, the approach relies on the actual observed τ values for uncensored observations, while estimates of the unobserved latent variables τ^* are obtained through Gibbs sampling from a distribution of the latent variable (truncated at $\overline{\tau}$) conditional on all other parameters in the model.[11]

The idea underlying the Bayesian spatial Tobit approach is similar to the EM algorithm proposed by McMillen [29], where the censored or latent unobserved observations on the dependent variable are replaced by estimated values. Given estimates of the missing values, the EM algorithm proceeds to estimate the other parameters in the model using methods applied to non-truncated data samples. In other words, conditional on the estimated values, the estimation problem is reduced to a non-censored estimation problem that can be solved using maximum likelihood methods. Similarly, once a sample for the unobserved latent dependent variables has been generated via the Gibbs sampler, the Bayesian estimation of the censored model reduces to an heteroscedastic spatial auto-regressive model. Moreover, the Bayesian spatial Tobit procedure yields the mean and dispersion of all parameters, including the crucial spatial lag coefficient.

2.3 A discrete hazard approach

Finally, it could be argued that the specification (7) is not really capturing the intergovernmental competition process that is likely to be at work, for two main reasons.

First, in the presence of yardstick competition - or even more so if local juris-

[10] The LeSage [24] Bayesian spatial Tobit approach is extensively discussed in Fleming [18] and LeSage and Kelley Pace [25]. The LeSage procedure has recently been applied to a local public finance context by Fiva and Rattso [17].

[11] The Gibbs sampler is an algorithm to generate a sequence of samples from a joint probability distribution of two (or more) random variables. It can be shown that the sequence of samples constitutes a Markov chain, and the stationary distribution of that Markov chain is just the sought-after joint distribution (LeSage and Kelley Pace [25]).

dictions were competing for a mobile tax base in a tax competition framework - each jurisdiction cares about the *actual* policies enacted by its neighbors, not their (unobserved) desired ones. As a result, the empirical specification would require the ideal tax rate of each government to be affected by neighboring juridictions' observed tax rates.

Second, the specification (7) relies on the assumption that, in every period, a government elaborates its optimal tax rate as a function of its own characteristics and neighbors' fiscal choices. However, it is rarely the case in reality that a government hitting the upper bound (the tax cap) ever reverts from there in the future.[12]

As far as issue one above is concerned, a simultaneous dependence model where desired policies τ^* are allowed to depend on neighbors' actual policies $(W\tau)$ is known to be algebraically inconsistent and cannot therefore be implemented empirically (Beron and Vijverberg [4]). As a result, we follow here the spatial discrete choice approach developed by Dubin [13] and implemented, among the others, by Hautsch and Klotz [21], Paez and Scott [32] and Paez et al. [32], and allow own attitudes towards taxation (τ_t^*) to be affected by neighbors' *lagged* fiscal policies $(W\tau_{t-1})$. In fact, the lagged specification can be justified by the idea that the adjustment to neighboring authorities' policies does not take place instantly due to the sluggishness of the political process.

As for the second issue, we treat the occurrence of a government hitting the upper bound as a discrete and irreversible event. In particular, we explicitly account for the upper bound in the dependent variable that is generated by the central cap, and model the event of local government i hitting the threshold $\tau_{it} = \overline{\tau}$ at some point $t = 1, ..., T$ as a "failure." Therefore, we estimate the probability - or hazard - of "exit" from the inner interval $(0, \overline{\tau})$ in period t conditional on having "survived" until then (Jenkins [22]; Winkelmann and Boes [42]).

In particular, let $T_i \in t = \{1, 2, ..., T\}$ denote the discrete survival time of local government i, i.e, the number of years that elapse before the government sets the maximum tax rate. The authorities surviving until the end of the period with $\tau_{it} < \overline{\tau}$ have a censored duration of $T_i = T$. The hazard function of T_i is the probability that $T_i = t$, conditional on government i not having failed in previous periods and on a number of time-varying characteristics - the vector x_{it} discussed with reference to the reaction function (1) - plus a set of time dummies

[12] See section **4** below for evidence in this respect.

(z_t) capturing duration dependence. By choosing a normal distribution for the formulation of the probability of exit, we can estimate a Probit model where the dependent variable takes value 0 in the years preceding the occurrence of the event $\tau_i = \overline{\tau}$. Censored observations - that is governments for which the event never occurs in the period considered here - take value 0 in all years. When the event occurs ($y_{it} = 1$), the local government exits the sample:

$$y_{it} = \begin{cases} 1 & \text{if} \quad \tau^*_{it} \geq \overline{\tau} \\ 0 & \phantom{\text{if}} \quad \tau^*_{it} < \overline{\tau} \end{cases} \tag{12}$$

In order to ascertain whether neighboring governments' fiscal choices affect the probability that a government hits the upper bound, we include the one-year lag of the spatially weighted average of neighbors' tax rates (with coefficient ρ_{-1}) among the explanatory variables, obtaining the following formulation for the desired tax rate:

$$\tau^*_{it} = x'_{it}\gamma + \rho_{-1} \sum_{j=1}^{N} w_{ij}\tau_{jt-1} + z_t + \eta_{it} \tag{13}$$

The specification of the desired tax rate in equation (13) implies that the lagged choices of neighboring jurisdictions can be treated as exogenous, and the hazard model can consequently be estimated by standard Probit.

3 The provincial vehicle tax in Italy

The Italian system of local government is organized as a three-tier structure, with over 8,000 municipalities, 100 Provinces and 20 Regions.[13] While Regions are in charge of health care services, Municipalities and Provinces share responsibility in the provision of local public services in the environmental, transportation, education and personal social service areas. Provinces play an important role in planning and coordinating municipal policies, particularly as far as the decisions that transcend strictly municipal boundaries - such as the control of industrial, car and heating pollution, as well as the management and disposal

[13] There also exist three "autonomous" Provinces in the upper North mountaneous bilingual regions, with special features and competencies. Due to their peculiarities, they are not considered in the rest of the analysis. In addition, seven new Provinces were recently created by redrawing the boundaries of the existing ones, thus bringing the number of Provinces to 107. However, we disregard them by focusing on the 2000-2006 time span, when the relevant data are available and consistent.

of waste - are concerned. Moreover, Provinces have exclusive responsibility for the construction and maintenance of intermunicipal roads, local transportation systems and secondary education schools and buildings.

Provincial expenditures rose considerably in recent years due to the devolution of (mostly) administrative duties from the regional and national governments. In fact, average per capita spending increased by about 40% in real terms between 2000 and 2006 - the time period on which the empirical analysis of this paper focusses.

Provincial authorities fund their expenditures through three main sources of revenues: 1) grants from upper levels of government; 2) revenue sharing arrangements; 3) own tax revenues.

First, grants to Provinces cover more than half of total current spending, with the proportion of grant-funded local spending slightly increasing over the 2000-2006 period. Grants from the central and regional governments are predominant and correspond to about $\frac{1}{6}$ and $\frac{1}{3}$ of local provincial spending respectively, the remaining grants being represented by specific transfers from other public bodies and international organizations.

Second, about $\frac{1}{4}$ of spending is funded via sharing of central government revenues, namely the personal income tax revenues based on taxpayers' province of residence (1% of it accruing to Provinces), and the motor-vehicle insurance tax, that - given the province-based territorial structure of the vehicle registration archive and bureaucracy - is entirely devolved to provincial authorities based on the province where the vehicle is registered.

Finally, slightly less than $\frac{1}{4}$ of provincial current spending is funded by own tax revenues, that are mainly constituted by the provincial vehicle registration tax (over 60% of own revenues), with the other own sources of revenues granting provincial governments either an extremely limited degree of fiscal autonomy or an almost negligible size of actual revenues.[14] As a result, the actual degree of fiscal autonomy of the Provinces basically rests on their ability to vary the vehicle registration tax rate.

The provincial vehicle registration tax was introduced in the year 2000 in order to attribute the provincial level of government an own source of revenues so as to reduce the reliance on external funding, and to foster the accountability of provincial administrations to their electorates. All brand new vehicles - as

[14] The other own sources of revenue include a surcharge on the national tax on the consumption of electricity, a surcharge on the municipal refuse collection charge, and a number of other minor sources of revenue.

well as used vehicles in case of change of ownership - are liable to the payment of the provincial vehicle registration tax the first time they are registered in the provincial archive under a given owner's name. The total tax due is made of a lump-sum amount plus a variable component that is related to the size, power and destination of the vehicle.

Central government establishes a lower and an upper bound on the vehicle tax parameters that Provinces can set, with the upper bound corresponding to a 20% higher tax burden than the one corresponding to the lower bound. Consequently, not having the power to alter the centrally set structure of the tax, the decision of each Province basically consists in determining autonomously the percentage tax spread (τ from here onwards, with $0 \leq \tau \leq 20$) relative to the lower bound ($\underline{\tau}$) set by central government.

Table 1 The provincial vehicle registration tax: key statistics

	Average τ	τ_t=$\underline{\tau}$	τ_t=$\overline{\tau}$	τ_t=$\overline{\tau}\,\vert\tau_{t-1} < \overline{\tau}$
2000	11.9	31	56	56
2001	14.5	19	68	12
2002	16.5	9	79	11
2003	16.7	8	80	1
2004	16.9	8	81	1
2005	17.6	6	86	5
2006	18	4	88	2
2000-2006	16.5			

Notes: 100 Provinces; τ is the provincial vehicle registration tax rate spread. $\underline{\tau}$ and $\overline{\tau}$ are the lower (0%) and upper (20%) bounds set by central government respectively.

Table 1 reports the average τ along with the number of Provinces setting the minimum ($\tau = \underline{\tau} = 0$) and maximum ($\tau = \overline{\tau} = 20$) tax spreads in each of the seven years following the introduction of the provincial vehicle tax (2000-2006). The table shows that Provinces steadily raised their tax spreads over time, with almost 90% of them hitting the upper bound by the year 2006.[15] A similar picture emerges from figure 1 in the Appendix, where the evolution of the geographical pattern of the provincial vehicle tax during the period under examination is depicted.

[15]Starting from 2007, the cap has been raised to 30%.

3.1 A race to the top in vehicle taxation? The null hypothesis

Before turning to the discussion and test of the hypothesis of "yardstick competition" as the driving force of the race to the top in provincial vehicle taxation, it is fair to briefly review some alternative interpretations of the observed phenomenon. In fact, the widespread growth of provincial vehicle tax rates might be exhaustively explained by the evolution of the underlying common determinants of provincial fiscal decisions (e.g., changes in the socio-economic environment, growth of spending needs, or decrease in external funding), even in the absence of inter-jurisdictional spill-overs. In correlated environments, local decision-makers acting in isolation might give the false impression of interacting with each other - the so-called "reflection problem" (Manski [27]).

There exist three plausible explanations of the observed evolution of the provincial tax pattern that are compatible with the the null hypothesis of absence of inter-provincial competition.

The first one has to do with the features of the central constraints imposed onto local decisions. Actually, the lower tax bound set by central government in the year 2000 remained fixed in nominal terms for all subsequent years.[16] As a result, provincial governments were in a way forced to raise their tax rate spreads in order to preserve vehicle tax revenues in real terms, inevitably ending up hitting the upper bound $\overline{\tau}$ at some point.

Second, the widespread increase in the provincial tax might have been fostered by the increase in public spending responsibilities of provincial governments as a result of the process of devolution of administrative functions by upper levels of government (Regions and State) during the 2000s. In fact, the late 1990s' reforms implied that the devolution of central and regional responsibilities to local governments had to be accompanied by an adequate transfer of resources (grants) to cope with the novel spending requirements, based on estimates of the additional costs for local governments. However, if central and regional grants systematically underestimated the new spending needs of provincial administrations, the latter would be subject to a mounting pressure to raise own sources of revenue. From 2000 to 2006, total financial resources transferred by the national and regional governments to the Provinces increased

[16] The lump-sum component of the vehicle registration tax was set in 2000 at euro 150 for cars and up to euro 650 for heavy trucks, while the variable component depended on the engine power and use of vehicles, and ranged from 1.75 euro to 3.50 euro per engine kW. Those tariffs remained unchanged over the subsequent decade.

by about 50% in real terms. While it is hard to say whether such a growth was commensurate to the enlargement in provincial functions and responsibilities, the financial inadequacy argument was frequently and forcefully put forward by provincial and municipal governments during the devolution process.

While the two arguments above rely on exogenous influences on provincial decisions, there might exist a third, endogenous cause of the increase in the provincial tax that shares the feature of requiring no competition among Provinces. In fact, during the years 2000s the Italian population rose from less than 57 to over 59 million and the total stock of vehicles expanded from less than 40 to over 46 million. Such a considerable change in the socio-economic environment might well have been responsible for an increase in the demand for public services, thereby pushing provincial governments to raise their own revenues and necessarily hit the upper bound $\overline{\tau}$ at some point, depending on the pace of growth of the provincial economy.

According to all of the above arguments, Provinces would simply face similarly evolving circumstances and would consequently react in similar ways, that is by manoeuvring their only autonomous means of raising revenues. In order to check whether those arguments can exhaustively explain the growth of provincial tax rates, we first speculate on the consequences of vehicle taxation in terms of government popularity. Afterwards, we will turn to the estimation of the provincial tax spread determination process.

3.2 Vehicle taxation and government popularity

One longstanding contribution of the "public choice" view of government recently revitalized by the so-called "political economy" literature (Persson and Tabellini [34], Besley [5]) was to demean the secular revenue-raising function of taxes and to stress instead their role as signals of politicians' quality and competence. In conventional political agency models, high tax levels are generally perceived by voters-principals as symptoms of inefficient use of resources, waste, transfers to special interests, or corruption. Consequently, if politicians know that voters have a tendency towards "fiscal conservatism," they might use their tax instruments strategically in order to minimize their loss of popularity (Peltzman [31]).

In particular, two circumstances are likely to affect the degree to which a tax increase translates into a loss of consensus and pluralities.

First, the popularity loss following a tax increase is arguably higher the closer is the tax rise to the next election. In fact, even without having to rely on sort of out-of-fashion assumptions on voters' memory, the media are likely to devote more attention to a government's fiscal policies in the proximity of an election.[17]

Second, the adverse popularity consequences of a tax rise are likely to be less severe if other governments in correlated economic environments are raising their taxes too. According to the "yardstick competition" hypothesis (Besley and Case [6]), a voting strategy based on relative fiscal performance evaluation is optimal in the presence of asymmetric information between taxpayers and policymakers and with multiple agents facing correlated fiscal shocks.

In the light of the above arguments, one could wonder whether the evolution of the provincial vehicle tax was influenced at all by electoral considerations. Provincial elections in Italy occur every five years, with direct popular election of the president of the Province, typically out of four to five candidates, and the members of the provincial Council. Councillors are elected by a proportional electoral system with a majority premium ensuring that the elected President is backed by at least 60% of the Council members. The elected President is the head of the provincial government, directly nominates the members of the government, and sets the provincial policy strategy and targets.

Table 2 summarizes the provincial election schedule in the 2000-2006 period. $\frac{3}{4}$ of the 100 Provinces held an election around the middle of the period (2003-2004), while some Provinces had an election in the early 2000s and went again to the polls five years later. Overall, we have 119 election occurrences over seven years.

Over the same period, table 2 also shows that the provincial vehicle tax spread was raised by provincial governments 113 times. Interestingly, though, the last column in table 2 shows that in only four of the 113 instances the tax rise occurred in a year when a provincial election was scheduled to take place, with the remaining 109 tax rises being decided in "safer" non-election years.

Furthermore, while the chances of success (re-election) of the incumbent in the overall sample exceed 75%, only 50% of the incumbents that raised the tax in election years managed to be re-elected.[18]

The above stylized facts provide a piece of suggestive evidence of oppor-

[17] See the evidence and references in Revelli [37].

[18] This evidence must of course be taken with caution due to the very few occurrences of tax rises in election years. In fact, the small number of observations precludes us from explicitly estimating a re-election function, and forces us to be content with these suggestive descriptive statistics.

tunistic setting of vehicle taxes. It seems that electoral considerations play a role, with provincial governments timing tax increases in order to minimize their adverse consequences in terms of popularity.

Table 2 Vehicle tax policy and provincial elections

	$el = 1$	$\Delta\tau > 0$	$\Delta\tau > 0 \vert el = 1$
2000	6	69	2
2001	9	14	0
2002	10	15	1
2003	12	2	1
2004	63	3	0
2005	6	7	0
2006	13	3	0
2000-2006	119	113	4
% re-elected	76.5		50

Notes: 100 Provinces; $el = 1$ in year t if a provincial election is held in that year. $\Delta\tau = \tau(t) - \tau(t-1)$ is the change in the provincial vehicle registration tax rate spread τ from year $t-1$ to year t.

4 The provincial vehicle tax setting process

The *prima facie* evidence from figure 1 along with the political economy suggestions in section **3.2** call for an empirical model that explicitly allows provincial tax setting polices to be interdependent. As long as electoral considerations play a role in local tax setting, the empirical model should allow the vehicle tax determination process in a Province to be affected both by "internal" determinants and by "external" determinants capturing the wider political environment in which local decisions are made.

According to the yardstick competition hypothesis, a particularly important role in this respect is likely to be played by the vehicle tax policies implemented by the Provinces that face a similar macroeconomic environment and are therefore most likely to face correlated shocks. The level of the provincial tax spread might convey a signal of the "quality" of policymakers, thereby generating a process of yardstick competition between consensus-seeking provincial governments. If information on a Province's vehicle tax policy spills over into neighboring Provinces, vehicle tax increases in nearby Provinces make it less costly in terms of popularity, and consequently more likely, for a Province to raise its tax too.

We start from the spatial lag specification (1) of section **2.1**. The vector of time-varying explanatory variables x_{it} includes grants per capita, income (value added) per capita, the stock of vehicles registered in the province in the previous year, a dummy that equals 1 in election years, a dummy that equals 1 if the government is right-wing, and the weighted average of neighboring provinces' tax spreads defined in equation (2).[19]

Since τ_{it} and τ_{-it} are determined simultaneously, equation (1) is estimated by spatial maximum likelihood techniques (Anselin [3]). Based on the fact that the (700×1) vector of average neighboring provinces' tax spreads τ_{-} equals $[I \otimes W]\,\tau$, where I is the (7×7) identity matrix and $W = \{w_{ij}\}$ is the (100×100) exogenous spatial weights matrix, the matrix form of equation (1) can be inverted as in (3) and estimated by maximum likelihood techniques.[20]

The results of estimation of a parsimonious specification of the spatial lag dependence specification (3) that includes no control variables ($\beta = 0$) are reported in table 3, while table 4 reports the results with all of the above explanatory

[19] The size of population residing in the province cannot be included because it is almost perfectly linearly correlated with the stock of vehicles (correlation coefficient > 0.99).

[20] The likelihood function is maximized via the MAXLIK procedure in GAUSS.

variables included. Both specifications include fixed Province (q_i) and time (z_t) effects.

In tables 3 and 4, the first column reports the results of OLS estimation of a non-spatial specification ($\rho = 0$); the second column reports the OLS estimates of the spatial lag specification; the third column shows the ML results of the spatial lag specification, under the hypothesis that ε_{it} is normally distributed; finally, the fourth column contains the ML estimation results of a spatial error dependence model (Anselin [3]), where $\rho = 0$ and nearby Provinces are allowed to be hit by spatially auto-correlated shocks:

$$\begin{aligned} \tau_{it} &= x'_{it}\beta + q_i + z_t + v_{it} & (14) \\ v_{it} &= \lambda v_{-it} + \varepsilon_{it} & (15) \end{aligned}$$

where, similarly to equation (2), v_{-it} is defined as:

$$v_{-it} = \sum_{j=1}^{100} w_{ij} v_{jt} \tag{16}$$

The tests for spatial auto-correlation in column (a) of table 3 and column (e) of table 4 point rather consistently towards positive spatial auto-correlation in the residuals of a non-spatial specification.[21] In the fully specified equation, the LM (Lagrange Multiplier) tests tend to favour the spatial lag dependence model over the spatial error dependence one: the LM test cannot reject the null of no spatial auto-correlation in the errors, while the corresponding LM test against a spatial lag of τ points to positive spatial auto-correlation in provincial tax spreads.[22]

[21] The Moran test is asymptotycally distributed as a standard normal under the null hypothesis of absence of spatial auto-correlation, while the LM tests against the hypotheses of a spatial lag - model (1) - or a spatial error - model (14)-(15) - are both distributed as $\chi^2(1)$ (Anselin [3]). All tests are performed in GAUSS.

[22] Clearly, the two LM tests take on the same value in table 3, where no exogenous variables are included, since the spatial lag and spatial error models are indistinguishable in that case, as shown by the ML estimation results in columns (c) and (d).

Table 3 Vehicle tax spread determination: baseline linear specification

	(a)	(b)	(c)	(d)
	ML	OLS	ML	ML
constant	-4.224***	-3.281***	-3.658***	-4.161***
	(0.373)	(0.464)	(0.429)	(0.422)
year 2001	2.580***	1.985***	2.222***	2.528***
	(0.528)	(0.553)	(0.540)	(0.596)
year 2002	4.740***	3.674***	4.099***	4.664***
	(0.528)	(0.611)	(0.578)	(0.596)
year 2003	4.970***	3.870***	4.309***	4.902***
	(0.528)	(0.616)	(0.581)	(0.597)
year 2004	5.100***	3.970***	4.421***	5.030***
	(0.528)	(0.621)	(0.584)	(0.596)
year 2005	5.890***	4.585***	5.107***	5.809***
	(0.528)	(0.651)	(0.604)	(0.597)
year 2006	6.290***	4.880***	5.444***	6.192***
	(0.528)	(0.671)	(0.615)	(0.597)
ρ		0.201***	0.121***	
		(0.060)	(0.047)	
λ				0.121***
				(0.047)
Fixed Province effects	yes	yes	yes	yes
Log likelihood	-1914.49		-1911.21	-1911.21
Observations	700	700	700	700
Moran test	3.358			
(p value)	(0.001)			
LM lag test	8.628			
(p value)	(0.003)			
LM error test	8.628			
(p value)	(0.003)			
LR test			6.552	6.552
(p value)			(0.010)	(0.010)

Notes: dep. var. = provincial tax spread $(0 \leq \tau \leq 20)$; standard errors in parentheses; *, **, *** (p-value < 0.10, 0.05, 0.01).

Table 4 Vehicle tax spread determination: full linear specification

	(e)	(f)	(g)	(h)
	ML	OLS	ML	ML
constant	-5.053***	-4.296***	-4.606***	-4.965***
	(0.505)	(0.612)	(0.567)	(0.531)
election dummy$_{it}$	-0.641	-0.580	-0.605	-0.632
	(0.443)	(0.436)	(0.428)	(0.447)
grants$_{it}$	-2.966***	-2.630***	-2.769***	-2.790***
	(0.729)	(0.743)	(0.733)	(0.757)
income$_{it-1}$	-0.446*	-0.417*	-0.429*	-0.435*
	(0.244)	(0.245)	(0.247)	(0.245)
stock of vehicles$_{it-1}$	1.774**	1.728**	1.747**	1.751**
	(0.718)	(0.714)	(0.717)	(0.716)
right-wing dummy$_{it-1}$	-0.575	-0.481	-0.519	-0.502
	(0.866)	(0.826)	(0.825)	(0.830)
ρ		0.135**	0.080*	
		(0.061)	(0.047)	
λ				0.062
				(0.048)
Fixed Province effects	yes	yes	yes	yes
Log likelihood	-1898.87		-1897.45	-1898.08
Observations	700	700	700	700
Moran test	1.882			
(p value)	(0.060)			
LM lag test	3.690			
(p value)	(0.055)			
LM error test	2.019			
(p value)	(0.155)			
LR test			2.840	1.570
(p value)			(0.091)	(0.210)

Notes: dep. var. = provincial tax spread ($0 \leq \tau \leq 20$); standard errors in parentheses; *, **, *** (p-value $<$ 0.10, 0.05, 0.01); time effects included.

The parsimonious specification in table 3 yields an OLS estimate of the spatial auto-correlation coefficient ρ of about 0.20, and an ML estimate of 0.12. After controlling for a number of exogenous local characteristics (table 4), evidence of a significant spatial dependence process in τ persists. While the (upward biased) OLS estimate of ρ in column (f) is 0.135, the ML estimate of ρ in column (g) of table 4 is an admittedly not overwhelming value of 0.08, pointing towards a pretty flat reaction function. However, as argued above, this is what one could expect given the censoring in the dependent variable, and $\widehat{\rho}_{ML}$ might be suffering from a downward bias.

On the other hand, λ - the auto-regressive coefficient of the error dependence process (15) - is not estimated to be significantly different from zero in column (h). The LR (Likelihood Ratio) test results reported at the bottom of columns (g) and (h) sort of redundantly summarize the above findings in terms of the overall performances of the spatial lag and spatial error models relative to a non-spatial specification.

As far as the other variables are concerned, right-wing ideology, proximity to elections, per capita income and grants from upper levels of government all tend to be associated with lower tax spreads.[23] Finally, the stock of vehicles circulating in the province has a positive effect on the provincial tax rate, possibly reflecting increasing marginal costs of providing transport-related public services due to congestion.

Table 5, columns (i) and (j), reports the estimation results of the Bayesian spatial Tobit model.[24] The estimate of the spatial auto-correlation coefficient is 0.15 and it is highly significant in the basic specification with $\beta = 0$, while it is around 0.10 when the effect of the explanatory variables on the provincial tax rate is accounted for. Overall, the Bayesian spatial Tobit results provide a very similar picture of the spatial pattern as the spatial lag specification that ignores censoring, and suggest that the latter is only weakly affected by a donward bias. Moreover, the coefficient estimates on the other explanatory variables are similar in the two models.

Finally, table 5 also reports the Probit estimates of the discrete hazard model discussed in section **2.3**. Estimation is performed on an unbalanced panel data set of 150 observations for the years 2001 to 2006, and exploits two interesting

[23] The negative effect of income is probably due to the fact that wealthier provinces are able to raise revenues from other tax sources, without having to increase the vehicle tax rate.

[24] Estimation is performed in Matlab based on the routines for a spatial auto-regressive Tobit model (sart_g function) provided by James LeSage (www.spatial-econometrics.com).

features of the data. The first consists - as shown in table 1 - in the fast growth in the provincial tax spreads that led 88/100 Provinces to hit the centrally set upper bound by 2006. Second, the decision to set the maximum tax spread seems to be an irreversible one. None of the Provinces that happened to end up in the upper corner solution $\tau = \overline{\tau}$ ever moved away from there. In terms of table 1, the 56 Provinces that set $\tau = \overline{\tau}$ in 2000 were stuck there for the rest of the period, with a varying number of further Provinces (from 1 to 12) joining them in each of the subsequent years. Since the 2000 cross-section is lost in taking the lag of neighboring Provinces' tax rates, and due to the fact that Provinces leave the sample when hitting the upper bound, we are left with 150 observations, 32 of which are censored. The resulting data structure for estimation of the discrete hazard model is depicted in figure 2 in the Appendix.

Column (k) in table 5 reports in particular the partial probability effects computed at the sample means. Partial effects for dummy variables are computed as the change in probability when a dummy variable shifts from 0 to 1, so that, for instance, the probability that a right-wing government hits the upper threshold is ten percentage points lower than it is for a left-wing government. The coefficient on the election year dummy has a similar size, but it is not statistically significant. As far as the effect of lagged neighboring Provinces' tax policies is concerned, it is estimated that an increase by 2 percentage points in the average tax rate τ of neighboring Provinces raises the probability of a Province hitting the upper bound $\overline{\tau}$ in the subsequent year by around 3 percentage points.

Overall, the evidence is generally consistent with the hypothesis of a process of inter-provincial interaction in the setting of vehicle taxes. In particular, the interaction process exhibits a geographical pattern, with the size of the spatial auto-correlation coefficient lying most likely in the vicinity of 0.10. In particular, the standard maximum likelihood estimate of ρ from a spatial lag dependence model that ignores censoring turns out to be remarkably close (0.08) to the estimate obtained in a bayesian spatial Tobit model that explicitly takes censoring into account (0.10). Interestingly, a similar picture emerges when estimating a discrete hazard specification that models the probability of a local government hitting the centrally set upper bound as a function of the tax rates chosen by adjacent governments in the previous period.

Table 5 Vehicle tax spread determination: alternative specifications

	(i)	(j)	(k)
	spatial Tobit	spatial Tobit	hazard Probit
election dummy$_{it}$		-0.609	-0.100
		(0.810)	(0.078)
grants$_{it}$		-1.245**	-0.001
		(0.643)	(0.001)
income$_{it-1}$		0.231***	-0.017***
		(0.056)	(0.006)
stock of vehicles$_{it-1}$		-0.645***	-0.030
		(0.064)	(0.051)
right-wing dummy$_{it-1}$		-2.589***	-0.109*
		(0.599)	(0.070)
ρ^*	0.153***	0.098**	
	(0.051)	(0.050)	
ρ_{-1}			0.015**
			(0.007)
year 2001	12.132***	12.985***	
	(1.023)	(1.510)	
year 2002	14.231***	15.526***	0.371**
	(1.134)	(1.609)	(0.197)
year 2003	14.443***	15.860***	0.438**
	(1.108)	(1.625)	(0.206)
year 2004	14.575***	16.169***	0.006
	(1.171)	(1.669)	(0.178)
year 2005	15.242***	16.458***	0.338*
	(1.203)	(1.596)	(0.234)
year 2006	15.626***	16.774***	0.173
	(1.199)	(1.585)	(0.237)
Observations	700	700	150
Right-censored ($\tau_{it}=\overline{\tau}$)	538	538	32

Notes: standard errors in parentheses; *, **, *** (p-value $<$ 0.10, 0.05, 0.01).

5 Concluding remarks

The empirical tax and yardstick competition literature that has grown impressively in the past two decades relies on the implicit and universal assumption that decentralised governments are free to set their tax policy instruments. However, local governments around the globe are hardly ever able to set the policies they see fit. In most instances, local governments' observed fiscal policies are censored due to existence of central government mandates or caps either on taxes or on public expenditures.

This paper has explored for the first time the consequences of central capping on the modelling and estimation of a local fiscal policy reaction function. By means of an empirical application on provincial vehicle taxation in Italy, we have employed three empirical approaches to the estimation of the inter-jurisdictional spatial interaction coefficient in the frequently encountered case of central government exercising its command by imposing upper limits on local fiscal choices.

It turns out that the three approaches provide a similar suggestion as to the direction and size of the interaction, and the conventional spatial lag dependence model that ignores censoring provides an estimate of the spatial autoregressive coefficient (0.08) that is reassuringly close to the Bayesian spatial Tobit estimate (0.10). Interestingly, a similar picture emerges when estimating a discrete hazard specification that models the probability of a local government hitting the centrally set upper bound as a function of the tax rates chosen by adjacent governments in the previous year. According to the lagged hazard model, an increase by two percentage points in the average tax rate τ of neighboring Provinces raises the probability of a Province hitting the upper bound $\overline{\tau}$ in the subsequent year by around three percentage points. Due to the strategic timing of vehicle tax increases by Provincial governments, we are inclined to attribute the cause of the observed spatial dependence to a political information spill-over, by which Provinces that "fail" - i.e., are forced to set the maximum vehicle tax rate to cope with their rising spending needs - make it less politically costly, and consequently more likely, for other Provinces to fail too.

While the empirical exercise performed in this paper might therefore tend to suggest that explicitly allowing for local tax censoring yields an admittedly small gain (and requires a high computational cost) relative to more standard estimation approaches, it should be taken into account that the direction and size of the bias caused by ignoring the censored nature of a local fiscal policy in a spatial lag dependence model are unknown a priori. In fact, since they plausibly depend on the intensity of the spatial interaction process, on the sample distribution of the latent variable, and on the structure of the spatial weights matrix, it seems that further theoretical and empirical research in this area is necessary in order to evaluate the relative merits of alternative estimation approaches in the presence of spatially dependent censored fiscal policies.

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Appendix

Table A1 Variables used in the analysis: descriptive statistics

	obs.	mean	s.d.	min	max
SPATIAL LAG DEPENDENCE AND TOBIT MODELS					
Vehicle tax spread (%)	700	16.5	6.9	0	20
Stock of vehicles (,000)	700	426.9	486.1	57.1	3514.2
Newly registered vehicles (,000)	700	31.4	50.5	3.1	471.2
Income (value added per capita; ,000 €)	700	20.1	5.0	10.8	34.3
Current spending per capita (€)	700	155.4	50.8	55.6	327.3
Grants per capita (€)	700	93.8	49.0	3.5	243.2
Right-wing control (%)	700	36			
DISCRETE HAZARD MODEL					
Vehicle tax spread (%)	150	8.9	8.0	0	20
Stock of vehicles (,000)	150	704.5	822.0	59.5	3383.1
Newly registered vehicles (,000)	150	56.6	86.5	3.2	414.6
Income (value added per capita; ,000 €)	150	19.7	6.1	10.9	34.3
Current spending per capita (€)	150	136.0	40.8	69.7	251.3
Grants per capita (€)	150	79.9	44.8	11.9	196.7
Right-wing control (%)	150	55			

Table A2 Variables used in the analysis: data sources

	SOURCE	YEARS
Vehicle tax	Automobile Club Italy - Quattroruote	2000-2006
Stock of vehicles	Public Registry of Vehicles	1999-2006
Vehicle registrations	Public Registry of Vehicles	1999-2006
Income	National Statistics Institute	1999-2005
Current spending	Italian Government, Home Office	2000-2006
Grants	Italian Government, Home Office	2000-2006
Right-wing control	Italian Government, Home Office	1999-2006

Figure 1: Vehicle tax spatial pattern

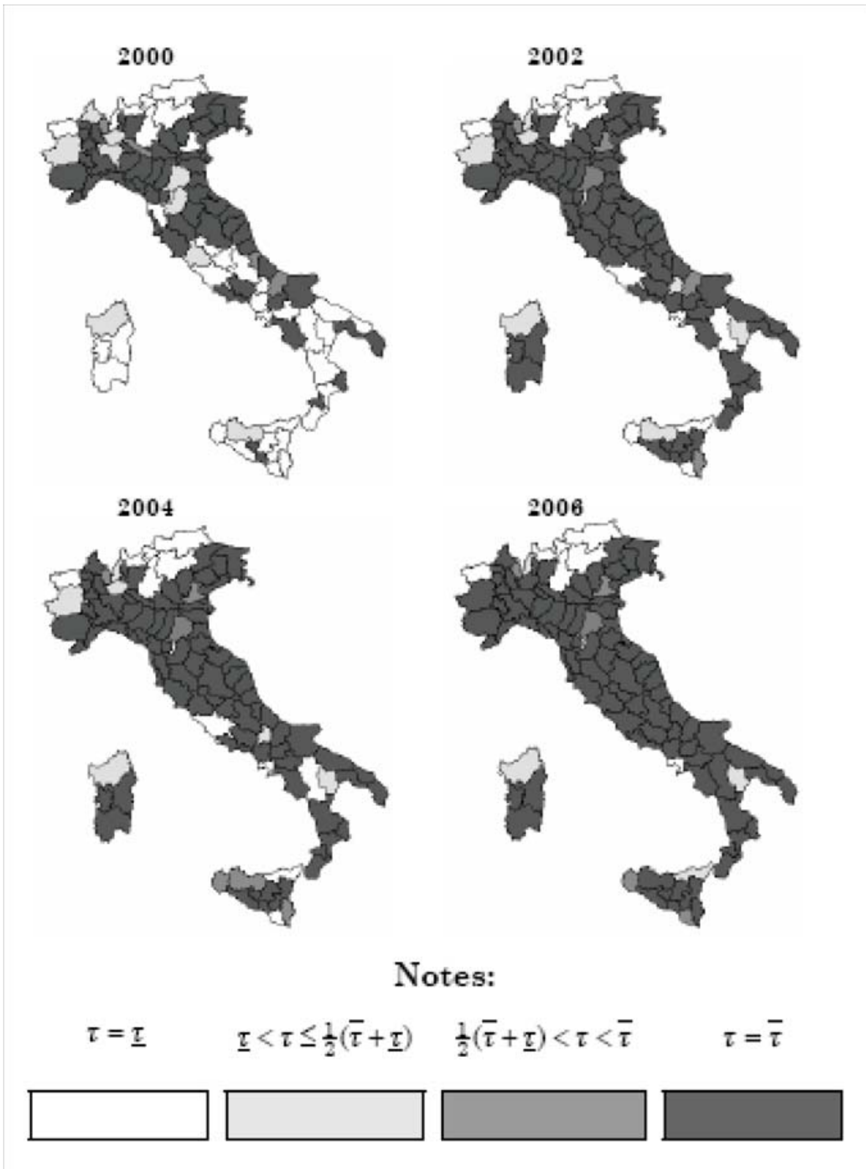
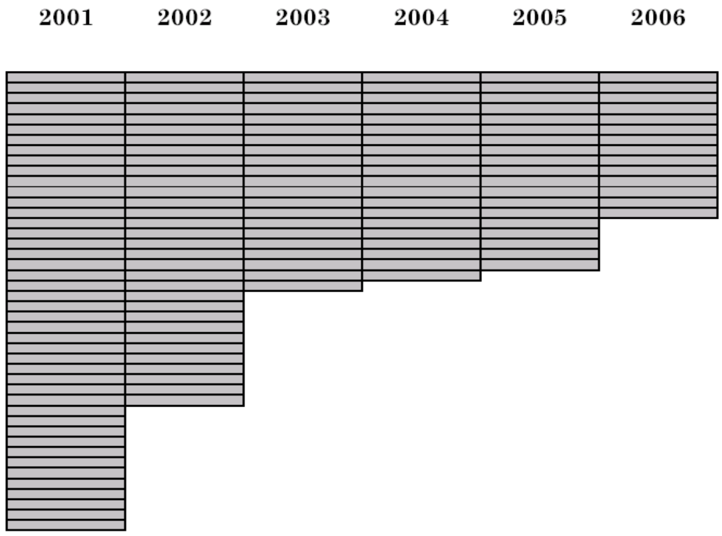


Figure 2: Data structure for discrete hazard model



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