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Spatial dependence through local yardstick competition: theory and testing

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Abstract

We propose a model of contracting for natural monopolies in which yardstick evaluation of performance can be optimal. Where principals have partially unobservable objective functions and agents are risk averse an externality is generated which can be observed in patterns of spatial dependence. Imposing standard contracting rules on principals can eliminate the externality and spatial dependence. We test this prediction using spatial econometrics on UK data covering a regime shift from independent contracting to compulsory competitive tendering rules. © 1997 Elsevier Science S.A.

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1. Introduction

This paper develops a model of yardstick competition[1] in a world where local principals contract with local agents. Yardstick competition emerges when it is optimal to condition local contracts on the performance of neighboring agents. An externality arises when particular principals pursue idiosyncratic and unobservable policies which then distort neighbors' incentive contracts. Because of these spillovers the results of our model can be seen as a new justification for imposing common competitive tendering rules on local principals by higher jurisdictions e.g. by federal government in the US on local government, or by the European Commission on member states of the European Union.

The externality described in our model also generates spatial dependence in the realized cost of the service under contract. We test our model using data on English garbage collection contracts before

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[1] The theory of yardstick competition was developed for the case of utility regulation in the work of Shleifer (1985). However, the notion of comparative performance evaluation through contracts had already been analyzed in the work of Lazear and Rosen (1981); Green and Stokey (1983), among others.

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and after the introduction of a law on compulsory competitive tendering (CCT) in 1988. Using spatial analysis techniques systematized by Cliff and Ord (1981), we show that there is a clear spatial dependence in the distribution of cost data pre-CCT. However, we also demonstrate that this dependence is substantially attenuated post-CCT. We also show that this change is associated with the increasing significance of market wage rates in explaining local variations in costs.

In Section 2 we describe the background to CCT in England. Section 3 outlines a model which generates spatial dependence. In Section 4 we explain the derivation of the spatial weights matrix used to identify spatial dependence. In Section 5 we report specification tests for the presence of spatial dependence and OLS and MLE estimates of a model with spatial dependence.

2. Garbage collection in England

Garbage collection is a very simple economic service with relatively few sources of variation. Previous studies have shown that the main causes of variations in cost are the number of properties from which garbage is collected, the proportion of these which are ordinary households, their density (units per square mile), whether they are urban or rural and local wage rates (see e.g. Stevens (1978) for US data, McDavid (1985) for Canadian data, Domberger et al. (1986) and Szymanski and Wilkins (1993) for English data). In addition, this research has shown that franchise monopolies subject to competitive tendering tend to achieve the lowest cost of service (compared to open competition or public service provision).

Domestic garbage collection is the statutory responsibility of local government in England, which is divided between 365 geographically distinct local authorities. Costs are covered by local taxes and grants from central government. Before 1988 over 90% of authorities fulfilled their obligations by employing their own staff and purchasing their own equipment, the unit responsible being called the DSO (Direct Service Organization). A DSO would typically consist of around 50 staff. DSOs are not allowed to take on business activities in the private sector, nor to work on behalf of other local authorities. In 1988 the central government imposed a law requiring all services to be put to tender under standard terms and conditions at regular intervals, thus allowing private contractors to compete against DSOs. DSOs were permitted to bid for contracts in their own authority after 1988 (but not in any other authority) and over 70% of tenders were won by DSOs[2].

In this paper we are interested in the spatial dimension of changes brought about under the CCT regime. Our explanation of the data is based on a specific principal-agent model of yardstick competition which can create externalities between neighboring authorities when principals are free to pursue idiosyncratic policies[3]. The model entails three critical assumptions. Firstly, agents are risk averse. We believe this assumption is justified because garbage collection is a small-scale activity (contracts have an average value of about £1 000 000 per year and last for 5 years). The agents are typically DSOs, small groups of public sector employees operating as a cooperative whose primary objective is to maintain their income and preserve their jobs. Even where contracts have been let to private sector firms, these firms have often been small local operations whose managers might plausibly be characterized as risk averse. Secondly, contracts are based on comparisons of

[2] However, service costs fell by about 20% on average (Szymanski, 1996).

[3] It may be possible to construct alternative explanations for the spatial correlations in our data. However, we believe our explanation is consistent with anecdotal evidence on the effect which CCT had on the structure of contracts.

performance against neighboring regions. Whilst there are no formal rules governing the way in which contracts are administered, comparison has been the main method for establishing service performance and has been encouraged by governmental advisory bodies such as the Audit Commission (1984), (1993). Most comparison tends to take place on a local basis, since managers and elected officials can monitor the performance of services of neighbors more easily than services located at the other end of the country. Thirdly, the imposition of standard rules under CCT in the UK reduced the scope for local authorities to pursue idiosyncratic policies. Prior to CCT, it is widely believed that a significant minority of local authorities engaged in feather-bedding of their own DSO employees by paying above market wage rates. Under CCT authorities were obliged to accept the lowest bid unless they could demonstrate that a higher bid offered better value for money (see e.g. Walsh (1991) p. 42). To retain the DSO against an outside bid typically meant reducing wages to the market rates.[4]

3. Spatial dependence through local yardstick competition

We develop a simple model to show how spatial dependence can arise through an externality in the contracting process between maximizing agents. Our assumptions are

- (A1) There are four principals, each of which independently contracts with one of four distinct agents to supply a local natural monopoly service.
- (A2) Principals are risk neutral but agents are assumed to be risk averse. Agents are assumed to have a CRRA utility function which takes the form $U=(w-e^2/2)^\theta$ where w is the payment from the principal, e is effort and $0<\theta<1$. Each agent has a reservation utility Z.
- (A3) For each principal, the service cost is $C=x-e+\varepsilon$, where x is a fixed and known parameter, e is the effort contributed by the agent, which reduces costs, and ε is a random variable. Effort is unobservable to the principal. The total cost to the principal is the service cost plus the (wage) payment w to the agent.
- (A4) For any service the random variable E can take one of two values, ε^H with probability p and ε^L with probability $1-p$. $E(\varepsilon)=0$. The random variables associated with the services for principals 1 and 2 are assumed to be correlated and the correlation is parameterised by r, so that the unconditional probabilities are $pr(\varepsilon_1^H, \varepsilon_2^H)=p[1-(1-p)(1-r)]$, $pr(\varepsilon_1^L,\varepsilon_2^H)=pr(\varepsilon_1^H,\varepsilon_2^L)=p(1-p)(1-r)$ and $pr(\varepsilon_1^L,\varepsilon_2^H)=(1-p)[1-p(1-r)]$. $r=0$ implies zero correlation, $r=1$ implies perfect correlation. However, these random variables are assumed to have zero correlation with the random variables associated with the services in authorities 3 and 4. The random variables for services 3 and 4 are defined in the same way, mutatis mutandis. Intuitively, principals 1 and 2 are neighbors, authorities 3 and 4 are neighbors, but neither pair has anything in common with the other pair.[5]
- (A5) Principals are restricted to offering one of two types of linear incentive contracts:

[4] We model the problem as one where a cost minimising principal (the local authority) offers an incentive contract to an agent (garbage-collection team). As pointed out by the referee, this might as easily be modelled as a problem where the local authority maximises the weighted average of consumer surplus and profit accruing to the garbage collection team.

[5] Neighborhood correlation in random shocks may arise because of local economic shocks. In the case of garbage collection a local plant closure may affect local incomes and therefore volumes of garbage generated, but have negligible national effects.

(1) independent: $w_i = a - bC_i$ where if $i = 1$, $j = 2$, if $i = 2$, $j = 1$,
(2) yardstick: $w_i = a - b(C_i - C_j)$ and if $i = 3$, $j = 4$, if $i = 4$, $j = 3$.

From assumption 4 it is clear that if yardstick comparison is ever optimal, it is only optimal with a neighbor whose random shock is correlated. It is straightforward to show that there exists a critical value for the correlation parameter r^* such that if $r > r^*$ a yardstick contract is strictly preferred to independent contract while for $r < r^*$ an independent contract is strictly preferred. To see this note first that under full information the optimal solution is to set the incentive parameter $b = 1$ and offer the agent full insurance. With asymmetric information the optimal value of the incentive parameter under the independent contract is

$$b^* = 1 - \frac{E(U'\varepsilon)}{EU'}$$

while under the yardstick contract it is

$$b_i^{**} = 1 - \frac{E[U_i'(\varepsilon_i - \varepsilon_j)]}{EU_i'}$$

Comparing b_i^* and b_i^{**} the only difference is that the covariance term depends on the composite random variable $\varepsilon_j - \varepsilon_i$ rather than ε_i alone. Given A4 these can be rewritten as

$$E(U'\varepsilon) = p(1-p)(U_H' - U_L')(\varepsilon^H - \varepsilon^L) > 0 \text{ (independent contract)}$$

(where U_H' means the marginal utility of income when $\varepsilon = \varepsilon^H$), while

$$E[U'(\varepsilon_j - \varepsilon_i)] = (1-r)p(1-p)(U_{LH}' - U_{HL}')(\varepsilon^H - \varepsilon^L) \text{ (yardstick contract)}$$

(where U_{LH}' means the marginal utility of income when $\varepsilon_i = \varepsilon_i^L$ and $\varepsilon_j = \varepsilon_j^H$). As r goes to unity the expectation of the covariance term goes to zero and b^{**} goes to unity, which is the full information value of the incentive parameter. With sufficient correlation between the two random variables the principal can achieve the first best solution.

This is simply a special case of the general case analyzed by Green and Stokey (1983). One implication of yardstick competition under these conditions is that even if the final cost realizations are correlated on a regional basis, these correlations will only arise from correlations in input costs. In particular, in a regression of service costs, a neighbor's input cost would not constitute an extra explanatory variable for own service costs (given that own input cost data is available). Thus once own input costs are allowed for there will be no residual correlation in costs between principals.

The notion of yardstick competition has often been criticized on the grounds that it is informationally demanding. In the current model, complete knowledge by, say, principal and agent 2 about the characteristics of the service of principal and agent 1 is a strong assumption. Here we illustrate the implications of incomplete knowledge by allowing some uncertainty on the part of principal and agent 2 about the reservation wage Z_1. This might occur for many reasons. For example principal 1 might be willing to cede high wages to a particular interest group which it favors, but be unwilling to reveal this fact to the tax-paying electorate. The neighboring principal and agent will therefore face some extra uncertainty when designing a yardstick contract.

- (A6) Suppose all parameters are identical for principals and agents 1–4, except that principal and agent 2 are uncertain about the wage policy of principal 1, and, therefore, about the reservation wage of agent 1. Suppose also that the correlation between the random variables for principals 1 and 2 and for principals 3 and 4 is sufficient to ensure that yardstick contracts dominate independent contracts. Suppose that principal and agent 2 have an expectation about agent 1's reservation wage which is simply $qZ^h + (1-q)Z^1$, where Z^1 is the reservation wage associated with the market wage rate and Z^h is a wage rate reflecting insider power of the incumbent. Thus principal and agent 2 now have an expectation of b_1 which is $E(b_1) = qb_1^h(Z_1^h) + (1-q)b_1^1(Z_1^1)$.

Proposition: Principal 1 will have the highest expected cost of service. If $r \leq r^*$ then principal 2 will strictly prefer an independent contract and will achieve the same level of costs as principals 3 and 4. If $r > r^*$, principal 2 will have the second highest expected cost, reflecting the externality created by uncertainty about principal 1's policy, while principals 3 and 4 will have identical and lower expected costs.

Proof: Firstly, all else equal, the higher expected reservation wage of agent 1 ensures a higher expected cost for principal 1. Now suppose $r \leq r^*$. In this case all principals will write independent contracts so that principals 2, 3 and 4 will have identical costs.[6] However, if $r > r^*$, agents 3 and 4 will definitely prefer to write a yardstick contract and agent 2 may want to. To compare contracts in this case we solve the constrained optimization problem. For principals 3 and 4 the Lagrangean can be written as

$$\min L_i = x + b_i(b_i - b_j - 1) + a_i - \lambda\{E[a_i - b_i(b_i - b_j + \varepsilon_j - \varepsilon_i) - b_i^2/2]^\theta - Z\} \tag{1}$$

(Incentive compatibility for each agent requires $b = e$ from the first-order condition for agent's effort). The first-order conditions for Eq. (1) are

$$\frac{\partial L_i}{\partial a_i} = 1 - \lambda EU_i = 0 \quad \frac{\partial L_i}{\partial b_i} = 2b_i - b_j - 1 - \lambda E[U_i'(b_i - b_j + \varepsilon_j - \varepsilon_i)] = 0 \quad \frac{\partial L_i}{\partial \lambda} = EU_i - Z = 0 \tag{2}$$

The optimal value of the incentive parameter b for agents 3 and 4 is given by

$$b_i^{**} = 1 - \frac{E[U_i'(\varepsilon_i - \varepsilon_j)]}{EU_i'} \tag{3}$$

However, principal and agent 2 know only the expected value of agent 1's incentive parameter and hence the first-order condition with respect to the incentive parameter b will be

$$\frac{\partial L_2}{\partial b_2} = b_2 - Eb_1 - 1 - \lambda\{E(U_2'b_1) + E[U_2'(\varepsilon_1 - \varepsilon_2)\} = 0 \tag{4}$$

Solving for the optimal value of the incentive parameter

[6] We assume that when $r = r^*$ principals choose the independent contract.

$$b_2^{***} = 1 - \frac{\mathrm{cov}(U_2'b_1)}{EU_2'} - \frac{E[U_2'(\varepsilon_2 - \varepsilon_1)]}{EU_2'} \tag{5}$$

Thus comparing Eqs. (5) and (3), $b_2^{***} < b_i^{**}$, $i=3$, 4. Principal 2 is constrained to offer a lower-powered incentive scheme than principals 3 and 4 and hence will face a higher expected cost of service, all else equal.

An important corollary of this result is that for $r > r^*$, the externality will manifest itself in spatial correlation between observed costs. In a regression of costs on observable characteristics, the effect of unobservable policies will manifest itself through a spatial autoregressive process (see below).

4. The data and the spatial weights matrix

Data was obtained on garbage collection costs and local characteristics for 324 out of the 365 English local authorities which are responsible for the service. A spatial weights matrix was used in testing the spatial dependence of the variables, the OLS regression residuals, and in the maximum likelihood estimation of the spatial autoregressive processes of the error terms. Here the matrix was constructed as a row-standardized binary contiguity matrix, where local authorities were defined as neighbors if they were linked in a Delaunay triangulation of the 365 districts in the complete data set after the removal of spurious links crossing estuaries and coastlines; 41 districts with missing data were then deleted from the contiguity matrix (for detailed presentation of Voronoi diagrams and the closely related Delaunay triangulation see Okabe et al., 1992). The observations were represented in the plane by the coordinates of the administrative seats of the local authorities. While the choice of an operationalization of the neighborhood relationship is arbitrary, spatial weights matrices of this kind are generally considered to assume least about the spatial processes under analysis (Cliff and Ord, 1981). Formally, the elements of the spatial weights matrix W are:

$$w_{ij} = \frac{w'_{ij}}{\sum_{j=1}^{n} w'_{ij}} \text{ where: } w'_{ij} = \begin{cases} 1 & \text{if } i \text{ linked to } j \\ 0 & \text{otherwise} \end{cases} \tag{6}$$

5. Estimation and testing

The basic regression used in the literature is simply an OLS regression of the logarithm of cost on the logarithm of the explanatory variables. The explanatory variables may be seen as reflecting structural factors affecting the technology employed by particular local authorities. We here take the OLS residuals as an indication of the service provision cost after having taken differences in technology into account. These differences relate to factors such as distance of haul to landfill site, landfill or incineration as disposal method, and many others, which we attempt to capture in the explanatory variables chosen. The results of OLS estimation are shown in Table 1. Over 80% of the variance of the logarithm of real net cost is accounted for by the explanatory variables in both the pre-CCT and the post-CCT models. Interpretations of the coefficients themselves have been given in other work, and are not our chief focus here (Szymanski and Wilkins, 1993). As can be seen, the

Table 1
Results of OLS estimation of pre-CCT and post-CCT models

	Pre-CCT	Post-CCT
Constant	−4.841	−5.918
	(4.1)	(5.6)
Log properties	0.962	0.924
	(25.4)	(26.0)
Houses%	−1.654	−1.403
	(5.2)	(4.7)
Log density	−0.013	−0.047
	(1.0)	(3.9)
London authority	0.164	0.109
	(2.5)	(1.8)
Other metropolitan authority	0.033	0.052
	(0.5)	(0.9)
Log wage rate	0.544	0.795
	(2.5)	(4.1)
Labour majority dummy	0.178	0.193
	(5.2)	(6.1)
σ^2	0.05371	0.04740

t-values in parentheses.

coefficient value and significance of the wage variable increase sharply, in accord with our expectations.

The theoretical model suggests that the service provision cost, net of technological differences should display more spatial dependence under neighborhood yardsticks, that is pre-CCT, than under competitive tendering (i.e. post-CCT). There is significant spatial patterning in all the explanatory variables, and in both the pre-CCT and post-CCT service provision cost variables (Moran's I significant at $\alpha=0.001$ or better; see Cliff and Ord (1981), for Moran's I). Any spatial patterning in the OLS residual term is the correlation between service provision costs for neighboring local authorities, when differences in technology have been taken into account. Using a two-directional Lagrange Multiplier test for spatial error and/or lag misspecification (Anselin et al., 1996), we find highly significant spatial dependence for the pre-CCT residuals, and less significant spatial dependence for the post-CCT error term. The results of this test on the residuals of the regression models shown in Table 1, using the spatial weights matrix described above were: pre-CCT model $LM=16.61$, prob. value$=0.0002$; post-CCT model $LM=6.19$, prob. value$=0.0452$.

We then sought to model the dependence by estimating the spatial autoregressive process of the error term directly for the pre-CCT and post-CCT cases. For brevity, we will not report on alternative formulations of the spatial process model considered in our work, nor on the issue of heteroskedasticity (see Anselin (1988) on spatial econometric techniques, and Anselin and Hudak (1993) for estimation methods). Table 2 displays the results of Maximum Likelihood estimation of the spatial error model for both pre-CCT and post-CCT cases (estimation methods based on Ord, 1975). As can be seen the spatial autoregression coefficient is larger in the pre-CCT case than in the post-CCT case. Specifically, for the pre-CCT case, $\rho=0.295$, significant at the $\alpha=0.00005$ level (two tailed), while for the post-CCT case, $\rho=0.199$, significant at the $\alpha=0.0093$ level. Summarizing, the coefficient

Table 2
Results of ML spatial error estimation of pre-CCT and post-CCT models

	Pre-CCT	Post-CCT
Constant	−4.866	−6.082
	(3.6)	(5.2)
Log properties	0.995	0.947
	(27.4)	(27.2)
Houses%	−1.749	−1.393
	(5.7)	(4.7)
Log density	−0.023	−0.052
	(1.8)	(4.3)
London authority	0.137	0.092
	(1.9)	(1.4)
Other metropolitan authority	−0.038	0.016
	(0.6)	(0.3)
Log wage rate	0.513	0.785
	(2.0)	(3.6)
Labour majority dummy	0.155	0.182
	(4.4)	(5.6)
ρ	0.294	0.199
	(4.1)	(2.6)
σ^2	0.05067	0.04617

Asymptotic t-values in parentheses.

values and significance of the wages variable rise from pre-CCT to post-CCT, while the spatial autocorrelation clearly falls.

We would not expect our results from modeling empirical data to switch from total absence to total presence (or vice versa) of these effects, but rather to trend in the predicted directions which they clearly do.

6. Conclusions and further research

This paper has shown how local yardstick regulation of natural monopolies can generate externalities and spatial dependence when principals are free to pursue idiosyncratic and unobservable policies. These effects can be reduced by imposing standard contracting rules as happened in the UK when compulsory competitive tendering was introduced. We have shown, using spatial econometric techniques, that the change in regime brought about by compulsory competitive tendering in the UK had the predicted effect. One direction for future research is to analyze the nature of the externality implied by local yardstick competition. The model suggests a new justification for imposing competitive tendering. In our model it is desirable not only because of its impact on a given jurisdiction, but because of the spillover benefits for its neighbors.

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