BANCA D'ITALIA
EUROSISTEMA

Temi di discussione

(Working papers)

Agglomeration within and between regions: Two econometric based indicators

by Valter Di Giacinto and Marcello Pagnini

Number 674

June 2008

The purpose of the Temi di discussione *series is to promote the circulation of working papers prepared within the Bank of Italy or presented in Bank seminars by outside economists with the aim of stimulating comments and suggestions.*

The views expressed in the articles are those of the authors and do not involve the responsibility of the Bank.

AGGLOMERATION WITHIN AND BETWEEN REGIONS: TWO ECONOMETRIC BASED INDICATORS

by Valter Di Giacinto* and Marcello Pagnini**

Abstract

We propose two indexes to measure the agglomeration forces acting within and between different regions. Unlike the existing measures of agglomeration, our model-based indexes allow for simultaneous treatment of both aspects. Local plant diffusion in a given industry is modelled as a spatial error components process (SEC). Maximum likelihood inference on model parameters is dealt with, including the problem of data censoring. The statistical properties of standard agglomeration indexes in the data environment provided by our SEC model are then treated. Finally, our methodology is applied to Italian census data for both manufacturing and service industries.

JEL Classification: R12, L70, C19.
Keywords: agglomeration, spatial autocorrelation, spatial error components model.

Contents

* Bank of Italy, L’Aquila Branch, Economic Research Unit.
** Bank of Italy, Bologna Branch, Economic Research Unit.

1. Introduction[1]

Economic theory emphasizes several economic mechanisms explaining the propensity of firms to co-locate. In the urban economics tradition, the main engine of agglomeration is given by the externalities produced by repeated and direct interactions between workers acting within a very narrowly defined local labour market. They include the so-called human capital externalities and knowledge spillovers. A more neo-classically oriented school underlies the role played by the exogenous distribution of local productive endowments in determining agglomeration, sometimes defined as natural advantages. Finally, the so called New Economic Geography explains the propensity to agglomerate with the input and output externalities endogenously generated by the joint action of increasing returns to scale internal to the firm, imperfect competition and transport costs.[2]

Some of these forces - like knowledge spillovers or natural advantages - may produce their effects within a very narrow spatial range, while others - like input or output linkages - display their effects on a wider spatial scale, though their intensity will decay with distance. Far from being mutually exclusive, these different factors are likely to co-exist within the same industry and therefore are very difficult to disentangle.

Mainly reflecting data availability, empirical literature on agglomeration usually assumes a partition of economic space into a finite number of regions. In this setting Ellison and Glaeser (also referred to as “EG” in what follows) or Gini Locational indexes measure spatial concentration or inequality, assuming that economic forces explaining agglomeration produce their effects within each region only. A second group of indicators, including the well-known Moran’s I index, abstracts from within region concentration and focuses on the behaviour of the process across regions located closely in space (polarization v. dispersion). While both theoretical considerations and empirical evidence would suggest a joint treatment of within and between regions agglomeration, the two statistical measures are derived under the assumption that only

---

[1] The authors wish to thank Stefano Iezzi and Andrea Lamorgese as well as participants at seminars held at the Bank of Italy and University of Cagliari and at the conference on Agglomeration and Growth in Knowledge-based societies (Kiel) for their helpful comments and suggestions. The views expressed herein are the authors’ and do not necessarily reflect those of the Bank of Italy.

[2] For the literature on urban economics see Belleflamme et al (2000) and Soubeyran and Weber (2002); for explanations based on local factor endowments see Kim (1995) and Ellison and Glaeser (1999). Finally the NEG school is surveyed in Fujita et al (1999).

one agglomeration process underlies the observed location patterns. To overcome this shortcoming, in this paper a new methodological approach is set forth, aiming at providing a simultaneous assessment of spatial concentration and polarization within a unified setting. In other words, spatial concentration will be estimated, conditional on polarization effects, without having to resort to the assumption that plant location decisions are only affected by internal market conditions. Mutatis mutandis, the same mechanism is at work when between region agglomeration forces are evaluated.

To this purpose, the proposed methodology assumes that the number of plants in a given sector/area can be treated as the realization of a spatial error components process (see Kelejian and Robinson, 1993). Such a process is the sum of: 1) a within (or local) component, modelled as an heteroskedastic white noise process, that is responsible for within region agglomeration, and 2) a between (or global) component, modelled as a spatial moving average process, that is responsible for polarization effects.

While our formulation is introduced mainly as a statistical tool useful to make joint inferences on agglomeration/polarization patterns, it is shown in the paper that it can be derived from a plant location choice model. We show that in our setting EG's and Moran's indexes are no longer unbiased measures of within and between agglomeration. Moreover, it is also shown how the bias affecting the two statistics can induce a spurious correlation pattern between the two.

Inference on model parameters is based on maximum likelihood, under a joint normality assumption for both components. However, since in many cases the observed number of plants in a given region/sector is zero, inferential procedures are subsequently adapted to deal with this censoring problem, by implementing the EM approach proposed by McMillen in the case of the spatial probit model.

Our methodology, while clearly more demanding from a computational point of view, has several advantages over existing indexes. First, it increases comparability of agglomeration indicators across sectors and time, given that each kind of agglomeration effect is measured controlling for the other, thus reducing possible interference effects. Second, statistical testing of the hypotheses for the presence of agglomeration and polarization is allowed for. Third, by dealing with the censoring problem, a straightforward treatment of cases where some regions host no plants is possible.

Finally, unlike some recently proposed agglomeration measures that are based on the knowledge of individual plant locations, our econometric set-up can be implemented using aggregate regional data.

An empirical evaluation of our proposed methodology is performed by resorting to a very detailed Italian census data set, reporting number of plants and employees broken down by sector of economic activity and geographic area. Specifically, we compute agglomeration indexes for several industrial and service sectors according to our methodology and compare results with the evidence provided by EG's and Moran's I indexes. Finally, a comparative analysis of agglomeration in manufacturing and service sectors is also carried out, given that spatial concentration and polarization may exhibit a different strength in these two fields of economic activity.

Results confirm that moderate levels of spatial concentration and polarization are present and coexist in an overwhelming majority of sectors. Manufacturing activities exhibit a stronger propensity to agglomerate within regions, while plants in service sectors are more likely to cluster in nearby geographical units. Differently from the proposed measures of within and between region agglomeration, EG's and Moran's indexes appear to be correlated across sectors, an occurrence that we are able to motivate within our unified econometric approach.

The remainder of the paper is organized as follows. Section 2 briefly reviews the literature on agglomeration and polarization indexes. Section 3 describes the modelling approach underlying the proposed simultaneous assessment of within/between markets agglomeration patterns and deals with maximum likelihood inference. Section 4 discusses the statistical properties of standard agglomeration indexes in the data environment provided by our econometric model. Section 5 presents our main empirical findings obtained from Italian census data and reports our own agglomeration indicators as well as EG's and Moran's I measures computed at three digit industry level. Section 6 compares spatial concentration and polarization between services and manufacturing. Finally, some concluding remarks are given in Section 7.

2. Measuring geographical agglomeration and polarization[3]

In this section we will survey two of the most popular agglomeration indexes, i.e. EG's γ for within agglomeration (1997) and Moran's *I* for polarization. Assume that the economy is made of several sectors and that economic space is divided into a given number of regions indexed by *i* *(i=1,….M)*. In a given sector there are *K* plants *(j=1,..,K)*, each representing a share z_j of sector employment. Let also x_i and s_i denote, respectively, the region *i*'s share of aggregate economic activity, as measured by total employment, and the share of industry's employees working in region *i* .

To measure within region agglomeration in each sector, Ellison and Glaeser (1997) propose the following agglomeration index:

$$\gamma_{EG} = \frac{G-(1-\sum_i x_i^2)\cdot H}{(1-\sum_i x_i^2)\cdot(1-H)} \qquad (1)$$

where $G = \sum_{i=1}^{N}(s_i - x_i)^2$ represents a relative geographic concentration index and takes on values ranging from 0 to 2, then $H = \sum_{J=1}^{K} z_j^2$ is the Herfindhal index at plant level in that particular industry. Positive values for this index indicate that plants in an industry tend to agglomerate beyond the level of concentration that would be generated by plant size and by the randomness in plant distribution across regions.[4]

Between regions agglomeration can be computed by relying on a standard spatial autocorrelation index like Moran's *I*. Specifically, let $div_i = s_i - x_i$, this index equals:

---

[3] Henceforth we also refer to "within agglomeration" as geographical or spatial concentration and to "between agglomeration" as spatial autocorrelation or polarization.

[4] Ellison and Glaeser show that geographic concentration generated by the effects of chance and plant size is equal to:

$$(1-\sum_i x_i^2)\cdot H$$

hence it follows that:

$$\gamma \geq 0 \quad if \quad G \geq (1-\sum_i x_i^2)\cdot H$$

In other words, a positive value for G is not a sufficient condition to generate agglomeration within a sector.

$$I = \frac{(N/S_o)\sum_i \sum_j w_{ij} \cdot t_i t_j}{\sum_i t_i^2} \qquad (2)$$

where $t_i = div_i - mean(div_i)$, w_{ij} equals 1 if regions i and j are 'neighbours' and zero otherwise ($w_{ii} = 0$ by convention) and where $S_o = \sum_i \sum_j w_{ij}$.

Moran's I index has a range of variation that depends on the spatial weighting scheme. In particular, when the spatial weights matrix, i.e. the matrix whose elements are the w_{ij}, is row normalized the maximum possible value of I is equal to 1. Spatial autocorrelation takes on positive and increasing values whenever regions with a high (low) specialization in a sector tend to cluster in space. In other words, positive spatial autocorrelation occurs when neighbouring regions exhibit similar values in their sectoral specialization. Alternatively, this index takes on negative values when neighbouring regions are dissimilar: highly specialized regions alternate with despecialized regions in space.[5]

These agglomeration measures certainly represent a substantial advance over previous indicators but at the same time they also exhibit some problematic aspects. The aim of this section is to illustrate the pros and the cons that may derive from their use.

A1) *Testing*

The two indicators allow testing for the presence of agglomeration and this is certainly an advantage over previous indicators. As for the γ index, Ellison and Glaeser recognize that randomness in plant distribution across regions and differences in plant size may generate geographic concentration that is not due to genuine agglomerative forces. Their index is precisely motivated by the need to net out the agglomeration measures from these effects and through that to increase the comparability of agglomeration indexes across sectors and time.

---

[5] As an alternative to the Moran's index, Cliff and Ord (1981) propose to measure spatial autocorrelation by means of the standard Pearson correlation coefficient. Letting $Lz_i = \sum_j w_{ij} z_j$ denote the spatially lagged value of z_i, the spatial autocorrelation index is given, in this case, by the usual expression:

$$P = \frac{COV(z_i, Lz_i)}{[VAR(z_i)VAR(Lz_i)]^{1/2}}$$

where P now ranges from –1 to 1 whatever the spatial weights specifications and, hence, provides more easily interpretable results compared to Moran's I.

A2) *Relative measures*

A second aspect shared by the two indexes is that they both represent relative measures of agglomeration, i.e. they both measure agglomeration in a sector relative to that observed for the overall economy.

A3) *Links to theory*

Finally, the γ index is grounded on economic theory. EG show how it can be derived from a location choice model, where profit maximizing firms choose their locations evaluating the strength of plant spillovers and natural advantages. They also show that their index cannot distinguish between the two agglomeration forces.

We will now move on to illustrate some of the drawbacks of the traditional agglomeration measures that motivate our proposed alternative indicators.

B1) *Simultaneity and the scope of spatial externalities*

A major problem that currently used indexes do not address is that they estimate one kind of externality between plant location decisions while ignoring the other. Consider first the γ index. Agglomeration forces, either natural advantages or plant spillovers, act within each specific region: natural advantages do not spread into other neighbouring regions; spillovers between plants located in different (even close) regions are ruled out by assumption. On the other hand, Moran's I index considers only connections between plants located in different (possibly close) regions ignoring the existence of within-type agglomeration forces. This circumstance may seriously bias the assessment of agglomeration forces. Considering that our alternative methodology is mainly motivated by this type of criticism, we will illustrate our argument through an example.

Assume that two sectors, *I1* and *I2*, have the same number of plants and that in both industries a given area, let us call it a_j, can potentially attract many plants due to natural advantages or other agglomeration forces. Further, assume that the strength of these localization economies is the same in the two sectors. Between region spillovers however are assumed to be very weak in *I1* and strong in *I2*. As a result, in *I1* economic activity will concentrate in that area and very few plants will be located outside its borders, including the nearby regions given that across region externalities are weak. In other words, there would be high spatial concentration and low polarization.

Now consider industry *I2* where agglomeration economies within a specific area coexist with across regions agglomeration benefits. Further assume that the intensity of the latter decreases with the distance to the region a_j. Now a firm can compare within agglomeration benefits deriving from locating in region a_j with those accruing by locating a plant in region a_k, where the latter is relatively close to a_j. As a result, many plants will still locate in region a_j as in sector *I1* but nearby regions will also attract a significant number of plants. Under this circumstance, one would observe lower spatial concentration (the same number of plants is spread over a larger number of locations) and higher polarization with respect to sector *I1*. Figure 1 gives an example of plant distribution across regions in the two industries, the grey area denotes region a_j.

Comparing the two cases, it can be observed that spatial concentration is higher in sector *I1* with respect to *I2*. Thus, a researcher could come to the conclusion that within agglomeration economies are stronger in *I1*. However, the difference in spatial concentration may only reflect the relative strength of across region agglomeration economies in the two cases. At the same time, the spikes in the spatial distribution of plants implied by the operating of region-specific factors would introduce outliers in the observed spatial distribution causing an underlying smooth pattern of polarization to stand out less clearly. Both types of limitations could be overcome by embedding the measures of the two different sources of agglomeration economies within a unified and consistent setting.

B2) *Spatial aggregation issues*

Both EG's γ and Moran's I index assume that the economy is partitioned into a finite set of regions. By doing so these indices transform dots on a map (plants) into units in a box (what we called regions and that are spatial units defined at a given level of aggregation).[6] This makes the agglomeration measures dependent on the chosen spatial scale. A variation in this scale can deeply influence the results based on a comparison of agglomeration forces across sectors or time. Moreover, spatial units at a given spatial scale may be defined according to administrative criteria, again distorting the analysis. Finally, spatial units are treated symmetrically in the EG approach, i.e. their index is insensitive to any spatial permutation of spatial units (Arbia, 2001).

[6] See Duranton and Overman (2005).

Moving from this criticism, Duranton and Overman (2005) and Marcon and Puech (2003) propose alternative estimators based on distance methods. They treat space as continuous and compute distances between plant pairs locations. These methods for computing agglomeration are very promising but they are also very demanding in terms of data availability as they require knowledge of each individual plant's exact location. [7]

B3) *More regions than plants*

Kim, Barkley and Henry (2000) recently argued that EG's index does not correctly assess the expected of value for G in the absence of agglomeration economies for those sectors in which the number of regions is greater then the number of plants. In particular the γ index overstates concentration in sectors for which there are more regions than plants.

B4) *Use of plant vs. employment counts*

The strength of spillover effects and natural advantages in the EG model does not depend on plant size. However, Holmes and Stevens (2002) empirically show that plant size increases with employment concentration in a specific region. Baldwin and Okubo (2006) give a theoretical underpinning to the positive correlation between geographic concentration and plant size by showing that large plants have a stronger propensity to locate in highly agglomerated areas. This spatial sorting mechanism introduces an upward bias in the measure of agglomeration.

To address this problem, a bunch of recent papers argue that an EG index should be computed using plant counts instead of employment shares.[8] Furthermore, building on Maurel and Sédillot (1999), Guimaraes, Figueiredo and Woodward (2004) show that an EG agglomeration index based on plant counts has the same expected value as the γ_{EG} index based on employment shares while achieving a lower variance.

B5) *Distance and between region spatial spillovers*

The weights matrix plays a central role in computing the Moran's I index. The weights define how distance affects the interactions of plants located in different regions. It is possible to assume that spatial interactions occur only between contiguous regions as in the example discussed above. Alternatively, it can be assumed that the

[7] For an alternative perspective on agglomeration indexes using aggregate data and based on Kullback-Leibler divergence, see Mori, Nishikimi and Smith (2005).
[8] See Lafourcade and Mion (2007).

strength of spillovers between any two regions increases with their proximity as measured by the inverse of distance. In any case, to compute *I* it is necessary to define a priori the structure of spatial interactions or connections between different regions. This may be a serious shortcoming, especially when the structure of these spatial interactions is not well-known a priori or when it changes according to the characteristics of different industries.

Following previous discussion, in the next section we propose two agglomeration indexes, one measuring spatial concentration and the other polarization, that should preserve the desirable properties A1-A3 shared by the two previous agglomeration measures and at least partially solve the problems described under B1-B5. These indicators: (1) simultaneously estimate spatial concentration and polarization, imposing relatively mild data requirements compared to plant level distance-based measures; (2) take into account the censoring problem caused by the existence of regions without plants; (3) are based on plant counts rather than on employment shares; and (4) allow a more flexible setting where the effects of distance on spillovers between different regions can be estimated from the data and not simply postulated.

3. The methodological approach

3.1 The unobserved components model

Our methodological approach to the statistical modelling of the spatial pattern of economic activity in a given sector is based on the assumption that the observed spatial distribution is shaped by the joint operation of three factors:

1. the overall attractiveness of the area for economic activity, measured by the share of the total number of plants located within the area;
2. a purely local unobservable random term, representing the influence on location choices exerted by natural advantages or localized information spillovers;
3. a second unobservable stochastic term, accounting for the spatial propagation of idiosyncratic random shocks originating within a given area and affecting plant location decisions also outside the area of origin.

The spatial error component (SEC) model, first introduced in Kelejian and Robinson (1993; referred to as KR in what follows), provides a methodology allowing for the joint consideration of these three factors and, with small adjustments, can form the basis for a simultaneous statistical analysis of both within and across market agglomeration.

The SEC model can be stated as follows:

$$s_i = x_i + \eta_i + \zeta_i \tag{3}$$

where η_i and ζ_i represent the *local* (or *within*) and *global* (or *between*) unobservable random components, assumed to be orthogonal. [9]

To complete the model, proper distributional assumptions, allowing for parameter identifiability, must complement expression (3).

While in the original specification the local error component is assumed to be the realization of a homoskedastic white-noise process, in analogy with the natural advantage firm location model set forth in Ellison and Glaeser, we model η_i as a zero mean random variable with continuous and unbounded support, variance $E[\eta_i^2] = \gamma x_i(1-x_i)$, $\gamma \geq 0$, and covariance $E[\eta_i \eta_r] = 0$ for $i \neq r$. For given x_i, as the non negative parameter γ increases, empirical realizations of s_i will be more likely to include positive outliers, hence displaying within region concentration, as a consequence of higher heteroskedasticity. At the same time, the restriction $\gamma=0$ implies, as in the EG model, the absence of agglomerating forces within regions.

Following the standard SEC approach, the global component ζ_i is subsequently modeled as a process of spatial diffusion of local disturbances u_i, assumed to be homoskedastic and uncorrelated across space. More specifically, we set:

$$\zeta_i = \sum_{m=1}^{M} w_{im} u_m \tag{4}$$

$$E(u_i) = 0\,,\ E(u_i^2) = \psi\,,\quad E(u_i, u_r) = 0 \ \text{when}\ i \neq r \tag{5}$$

[9] While, in the present context, we make reference to the SEC approach mainly for statistical modelling convenience, it is shown in Appendix A how such specification can be motivated as a linear approximation to the solution of a discrete choice problem closely related to the ones usually dealt with in the literature on plant location.

$$(\eta_i \perp u_i) \tag{6}$$

where w_{ir} is a non negative weight measuring the degree of spatial proximity of locations *i* and *r*.

Expressions (4) and (5) jointly define a spatial moving average (SMA) process (Haining, 1978; Anselin, 2003) that, differently from the local component η, is correlated across regions, its autocovariance function being given by:

$$E[\zeta_i \zeta_j] = \psi \sum_{m=1}^{M} w_{im} w_{jm}, \qquad i, j = 1,..., M \tag{7}$$

As usual with spatial econometrics specifications, the pattern of spatial propagation of local random disturbances reflects the selected weighting scheme. Two main approaches have been proposed in the literature to define spatial weights. The first assigns non zero weight only to linkages between nearest neighboring locations, usually defined as areas sharing a common border or separated by a distance not exceeding a given threshold. In the second, spatial weights are assumed to be a decreasing function of distance, taking positive values for each couple of locations and converging to zero as distance between locations diverges.

While the first weighting scheme is adopted in KR, in our implementation of the model we give preference to the second approach, since it can account for long range interactions of the kind implied, for example, by the existence of spatial externalities generated by a market potential mechanism. More specifically, letting d_{ij} denote geographical distance between locations *i* and *j*, we set:

$$w_{ij} = d_{ij}^{-\delta} / \sum_{r=1}^{M} d_{ir}^{-\delta} \tag{8}$$

where the coefficient δ, measuring the rate of distance decay of between markets spillover effects, can be set a priori or estimated from observed data.

In general, when dealing with aggregate area data, we will set $d_{ii}>0$, $i=1,\ldots,M$, thus allowing, differently from the KR approach, for global shocks to affect within market conditions as well.

To provide a graphical depiction of the spatial patterns that the model is able to replicate, Figures 2 and 3 display simulated series respectively for the η and ζ model components, using a 20x20 lattice as spatial reference. Figure 4 depicts the combination of the two components, i.e. a realization of the (*s*-*x*) process, assuming a variance ratio that allows for both components to be still separately detectable in the graph.

3.2 Maximum likelihood parameter estimation

The variance-covariance matrix of the vector $(s-x)$, $s=[s_1,\ldots,s_M]'$, $x=[x_1,\ldots,x_M]'$, takes the following form:

$$\Sigma = \gamma \Xi + \psi W W' \tag{9}$$

where:

$$\Xi = diag\{[x_1(1-x_1),\ldots,x_M(1-x_M)]\} \tag{10}$$

and where W is the MxM matrix with elements *wir*. For analytical purposes, it can be more conveniently reparametrized as:

$$\Sigma = \lambda[(\Xi^* + \tau W^* W^{*}{}'] = \lambda\Omega \tag{11}$$

$$\lambda = \gamma\kappa_x \tag{12}$$

$$\tau = (\psi/\gamma)(\kappa_w/\kappa_x) \tag{13}$$

$$\Xi^* = \kappa_x^{-1}\Xi;\;\; W^* = \kappa_w^{-1/2} W \tag{14}$$

where κ_x and κ_w denote, respectively, the median values of the diagonal terms of Ξ and WW'. In this setting, parameter τ measures the ratio of the variances of the global component to the local component, evaluated at the median point of the cross-sectional variance distribution of the two error components since both are heteroskedastic. This modified variance ratio provides, as in the KR model, an assessment of the relative strength of between markets interactions compared to within market externalities. A value of τ equal to 1 implies that, when evaluated at the centre of the respective cross-sectional distributions, within and between variance components contribute equally to the spatial variation of economic activity in a given sector. As τ approaches 0, the contribution of the between component to the process variance becomes negligible, while the opposite result holds for values of τ largely above 1.

Assuming a Gaussian distribution for both random components, the log-likelihood function of the model parameter vector $\theta = [\lambda;\tau]'$ takes the following form:

$$L(\theta) = -\frac{M}{2}\log(2\pi) - \frac{M}{2}\log(\lambda) - \frac{1}{2}\log(|\Omega|) + -\frac{1}{2\lambda}(s-x)'\Omega^{-1}(s-x). \qquad (15)$$

Details on the derivation of maximum likelihood estimators are given in Appendix B.

3.3 Testing hypotheses on model parameters

In bringing the model to data, a researcher will usually be interested in testing these two statistical hypotheses:

$$H_0^1 : \lambda = 0 \qquad (16)$$

$$H_0^2 : \tau = 0 \qquad (17)$$

respectively corresponding to the absence of within and between market agglomeration effects. While the test of the second hypothesis can be carried out in the usual fashion, i.e. by means of a *t*-test based on the asymptotic distribution of maximum likelihood estimators, in testing H_0^1 one has to take into account the violation of regularity conditions implied by fact that the parameter lies on the boundary of the admissible space under the null hypothesis.

Anselin (2001), tackling a similar problem in the context of the KR SEC model, develops a Rao’s score testing procedure, also known as the Lagrange multiplier (LM) test, and a similar approach can be implemented in the case of our modified SEC specification.

The general expression of the LM test statistic is given by the following quadratic form:

$$LM = d(\theta_0)'J(\theta_0)^{-1}d(\theta_0) \qquad (18)$$

where θ_0 is the parameter vector evaluated under the null hypothesis. The derivation of the tests of the two hypotheses listed above, both asymptotically distributed as a χ^2 with 1 degree of freedom, is carried out in Appendix C.

3.4 Dealing with censoring

In the empirical implementation of the proposed methodology it must be taken into account that, as the size of geographical units decreases and the sectoral detail of the industry classification increases, there might be frequent occurrences of regions with no plants. Such occurrences induce a censoring problem that complicates statistical inference on parameters of interest.

In line with the standard econometric approach, we will deal with the censoring issue by assuming that the left hand variable in (3) is a latent variable that is only observed when it takes on positive values, i.e. letting s^* denote the observed censored variable we set:

$$\begin{aligned} s_i^* &= s_i, \text{ if } s_i > 0 \\ s_i^* &= 0, \text{ if } s_i \leq 0. \end{aligned} \tag{19}$$

Under the normality assumption for the distribution of s^*, the model defined by expressions (3) and (19) can be interpreted as a heteroskedastic and spatially autocorrelated error Tobit model. It is well known that the latter yields a computationally intractable likelihood function, due to the presence of a multiple integral of high dimensionality (Fleming, 2004). To overcome such difficulties, in the strictly related case of the spatial probit model, McMillen (1992) proposes to base parameter estimation on an implementation of the EM algorithm, initially introduced by Dempster *et al.* (1977).

The EM approach is structured in two steps: Expectation and Maximization. In the present context the E step requires the computation of the predicted values of the unobserved latent variable (Expectation), conditional on observed explanatory variables and on a set of values for the model parameters. In the M step the objective function, given by the likelihood function of the underlying latent process, is maximized over the admissible parameter space taking as inputs the expected values obtained in E step. The two step are subsequently iterated until parameter estimates converge.

By making reference to the properties of the censored normal distribution, the following expression for the conditional expected value of s_i can be derived:

$$E[s_i \mid s_i^* = 0] = x_i - \sigma_i \frac{\phi(x_i/\sigma_i)}{1-\Phi(x_i/\sigma_i)} \tag{20}$$

where $\sigma_i^2 = E\left[(s_i - x_i)^2\right] = \gamma x_i(1-x_i) + \psi \sum_{j=1}^{M} w_{ij}^2$ is the variance of the composite error term $(\eta_i + \zeta_i)$ and where ϕ e Φ denote, respectively, the standard normal probability density and cumulative distribution functions.

The predicted latent values are then substituted for the corresponding censored values and the algorithm proceeds with the M step, involving the maximization of (15).

The computational intractability of the likelihood function due to censoring makes the derivation of LM test statistics unfeasible for the hypotheses H_0^1 and H_0^2 along the lines given in the previous section.

By analogy with the EM approach adopted for the problem of parameter estimation, we propose to carry out the test procedure using the test statistics derived for the case of uncensored observations, but replacing unobserved censored values with their respective expected values. The impact of this replacement on the asymptotic distribution of the test statistics is assumed to be negligible at the present stage, but will be subsequently analyzed by means of stochastic simulation techniques.

4. Properties of standard agglomeration indexes in the SEC environment

In this section we study some statistical properties of the EG's and Moran's indexes when we assume that the data are generated by a SEC process with the above stated features.

As regards the EG index, since we focus on plant counts, we can assume that the number of plants, K, tends to infinity and therefore $H \cong 0$. The EG index formula, in this case, simplifies to:

$$\gamma_{EG} = \frac{G_m}{(1-\sum_i x_i^2)} = \frac{\sum_i (s_i - x_i)^2}{(1-\sum_i x_i^2)} \tag{21}$$

From (4) we get $s_i - x_i = \eta_i + \zeta_i$. Substituting this expression in (21) and taking expected values yields:

$$\begin{aligned} E[\gamma_{EG}] &= \frac{\sum_i E(\eta_i + \zeta_i)^2}{(1-\sum_i x_i^2)} = \frac{\sum_i E(\eta_i)^2}{(1-\sum_i x_i^2)} + \frac{\sum_i E(\zeta_i)^2}{(1-\sum_i x_i^2)} = \\ &= \gamma \frac{\sum_i x_i (1-x_i)}{(1-\sum_i x_i^2)} + \psi \frac{\sum_i \sum_m w_{im}^2}{(1-\sum_i x_i^2)} = \gamma + \psi\phi \end{aligned} \tag{22}$$

showing how the EG index provides an unbiased estimator of the SEC model parameter γ only in the case when the variance of the between component is 0 (i.e. when ψ=0). Given that ϕ>0 by definition of the x, when ψ>0 the EG index overestimates γ with a bias increasing linearly with the value of ψ (or, equivalently, τ).[10]

To analyze the behaviour of the Moran's *I* index, we start by noting that in our SEC model, spatial autocorrelation in the observed process is entirely related to the between component, since the within component is uncorrelated over space. Applying the well-known equivalence of the Moran's *I* statistic to the OLS regression coefficient of the spatial lag of a given variable on the same variable not lagged in space, we have the following expression for the spatial autocorrelation of the within component:

$$I_\zeta = \frac{COV[L\zeta\zeta]}{VAR[\zeta]} \tag{23}$$

where *L* denotes, as above, the spatial lag operator. When evaluated on observable data Moran's index expression becomes:

$$\begin{aligned} I &= \frac{COV[L(\eta+\zeta)(\eta+\zeta)]}{VAR[(\eta+\zeta)]} = \frac{COV[L\zeta\zeta]}{VAR[\eta]+VAR[\zeta]} = I_\zeta R \\ R &= \frac{VAR[\zeta]}{VAR[\eta]+VAR[\zeta]} \le 1. \end{aligned} \tag{24}$$

As expected, given the analogy with a standard errors-in-variables problem, apart from the trivial case where γ = 0, the observable version of Moran's *I* appears to

[10] With a finite number of plants, H=1/K, and expression in (22) becomes:

$$E(\gamma_{EG}) = \left(\frac{k}{k-1}\right)(\gamma + \psi\phi) - \frac{1}{k-1}.$$

understate the actual underlying value by an amount that is increasing with the variance of the within region component.

It is quite common, in the empirical literature, to repeat the analysis of agglomeration on a selected range of economic sectors. In our setting this amounts to considering a number of different instances of the SEC process, each observed on the same set of spatial units. If model parameters are allowed to vary across processes, the biases affecting the EG's and Moran's indexes in the SEC environment will vary accordingly, possibly inducing a spurious correlation pattern between the two measures even when actual parameter variation is completely random.

Let us consider first the case when γ is positive and constant across sectors, while ψ is allowed to vary randomly. To qualify the mathematical relation linking γ_{EG} and I in this setting, let us consider the following derivatives:

$$\frac{\partial E[\gamma_{EG}]}{\partial \psi} = \phi > 0 \tag{25}$$

$$\frac{\partial I}{\partial \psi} = I_{\zeta} \frac{\partial R}{\partial \psi} = I_{\zeta} \frac{\partial VAR[\zeta]}{\partial \psi} \frac{VAR[\zeta]}{(VAR[\eta] + VAR[\zeta])^2} > 0 \tag{26}$$

Expression (25) states that, while γ is constant across processes by definition, γ_{EG} will display some variation, induced by the change in ψ. In particular, γ_{EG} is immediately shown to be an increasing function of ψ, for a given value of γ.

At the same time, inequality (26), stemming from the relation linking the variance of ζ to ψ (see expression (7)[11]), shows how *also* Moran's I computed on observable data is an increasing function of ψ.

Taken together, the two results imply that, when different instances of the SEC process (i.e. different sectors) are evaluated jointly and γ is fixed across instances, any variation in ψ will result in a positive correlation between γ_{EG} and I, since they both increase as ψ increases.

Turning to the case when ψ is constant while γ is allowed to vary randomly across sectors, we obtain, in a similar fashion:

[11] Since our SEC model can only accomodate positive values of spatial autocorrelation, we can rule out as irrelevant the case when $I_{\zeta} < 0$.

$$\frac{\partial E[\gamma_{EG}]}{\partial \gamma} = 1 > 0 \tag{27}$$

$$\frac{\partial I}{\partial \gamma} = I_{\zeta} \frac{\partial R}{\partial \gamma} = -I_{\zeta} \frac{\partial VAR[\eta]}{\partial \gamma} \frac{1}{(VAR[\eta] + VAR[\zeta])^2} < 0 \tag{28}$$

the latter inequality stemming from the fact that $\partial VAR[\eta] / \partial \gamma$ is positive by definition. Since γ_{EG} and I respond in opposite ways to an increase in γ, in this case the two measures will tend to display a negative correlation when evaluated across sectors.

In both situations, inaccuracies, due to the fact that one of the two sources of spatial variation in the model has been neglected, will result in a correlation between EG's and Moran's indexes that is totally spurious, i.e. not induced by an underlying pattern linking the change in model parameters across different instances of the process, that is assumed to be completely random.

When random parameter variation is allowed for, the SEC model can thus account for both positive and negative correlation between the estimates of γ_{EG} and I measured over different economic sectors. The actual sign of the correlation, when both γ and ψ are allowed to vary simultaneously across different instances of the process, will eventually reflect the sign implied by the prevailing source of parameter variation.

To provide some more insight on this aspect, a Monte Carlo experiment was carried out by simulating the process under different hypotheses regarding the variance of γ and ψ across realizations (Appendix D details the design of the experiment).

The simulation evidence shows that, although the values of γ and ψ were drawn independently, the correlation between γ_{EG} and I can be substantial, taking positive or negative values according to the prevailing source of parameter variation. In particular, as expected, for a given range of the distribution of parameter ψ, the value of the correlation turns from positive to negative as the dispersion of the distribution of γ increases.

Finally, it can be noted that changing the value of δ, while affecting the correlation pattern in a significant way, leaves qualitative findings essentially unchanged.

5. The empirical analysis

In this section, the agglomeration indexes reviewed in Sections 2 and 3 are computed on Italian census data, as collected by the Italian National Institute of Statistics (ISTAT) in 2001.

The database reports employees and number of plants with a very detailed sectoral and geographical breakdown. In this paper, we adopt a sectoral classification based on 103 manufacturing industries (3 digits) and 81 service sectors. Geographical breakdown is given by the 686 local labour market systems (LLMS) in which the Italian territory can be subdivided according to travel to work mobility flows recorded in the 2001 population census. This definition of spatial units is based on economic rather than administrative criteria and therefore should attenuate the impact of the aforementioned modifiable unit area problem.

Figures 5 and 6 graph the cross-sectional distribution of (s_i - x_i) for two manufacturing sectors, i.e. "Manufacture of vegetable and animal oils and fats" and "Manufacture of machine-tools". In both cases the plot features a number of sharp spikes, providing some preliminary evidence on the existence of very (low) high concentration of activity in a few (de)specialized areas. At the same time the spatial pattern of the data, displayed in Figures 7 and 8 for the same sectors, bears out the existence of smooth spatial trends, neighbouring markets showing mostly similar levels of the variable.

Such graphical evidence, that is broadly shared by all the sectors considered in our empirical analysis, appears to support our basic assumption that both local and global agglomerating forces play a role in shaping observed plant location patterns.

Our proposed measures for spatial concentration (denoted here as γ_w to distinguish it from the corresponding parameter in the Ellison and Glaeser model) and polarization (denoted by τ), were obtained following the econometric set-up described in Section 3. s_i and x_i are computed using the number of plants and not of employees, for the reasons mentioned in section 2.

The value of parameter δ was not imposed a priori, but was estimated from the data by maximum likelihood jointly with γ_w and τ.

When the fraction of LLMSs for which zero plants were recorded exceeded 2 per cent in a given sector, inference was carried out correcting for censoring. Censoring was observed for an overwhelming majority of manufacturing sectors, the share of censored observations being lower than 2 per cent in only 4 out of 103 industries. Service activities appear to be more evenly distributed across LLMSs. In 25 per cent of the service sectors, censored observations accounted for less than 2 per cent of the sample size.

Furthermore, two versions of the EG index are computed, one based on plant counts, γ_{EG1}, and another one based on employees, γ_{EG2}. This is done to allow a better comparison with other papers reporting γ_{EG} calculated on employment shares. Finally, we also computed the standard Moran's *I* spatial autocorrelation index.

We first test for the presence of within and between agglomeration according to our indexes. It turns out that all manufacturing and service sectors display positive values of γ_w. Moreover, all these values are significantly different from zero at the 95 per cent level of confidence. Parameter τ is also positive and significantly different from zero in an overwhelmingly majority of sectors, except for 5 manufacturing and 3 service industries.

Hence, according to this evidence there is a general propensity of economic activities both to concentrate in some specific regions and to cluster across nearby regions. These findings are similar to those obtained in other papers that measure agglomeration at sectoral level.[12] Ours however are obtained under the assumption of a possible coexistence of spatial concentration and polarization.

Although the propensity to concentrate in space is well spread across industries, most activities display low values of γ_w. The median is very close to zero both in the manufacturing and service sectors (see Table 1). A graphical analysis confirms that the distribution of γ_w across sectors is very skewed and that most industries display values between 0 and 0.02 (see Figure 9). Estimates for τ clearly indicate the prominence of within over between agglomeration forces in determining the spatial distribution of

[12] See, among others, Ellison and Glaeser (1997) and Rosenthal and Strange (2001) for the USA, Maurice and Sèdillot (1999) for France, Devereux, Griffith and Simpson (1999) for the United Kingdom and Pagnini (2003) for Italy.

several economic activities. Its median is well below unity both in manufacturing and in services. Moreover, the 75th percentile for the same variable is still lower than unity (0.71) in manufacturing, while it is above unity (1.75) in the service sector.

Estimated values for δ are well below unity, denoting a slow rate of spatial decay of between regions externalities. Consistently with a priori expectations based on the higher degree of tradability (and hence sensitivity to external market conditions) of the output in manufacturing compared to services, the average value of δ is lower for manufacturing than for services (0.21 and 0.31, respectively).

Overall, γ_{EG1} appears to take higher values compared to γ_W, the median across sectors being about twice as large. On the basis of the results shown in Section 4, such evidence can be motivated by the positive bias affecting the former when between regions agglomerating forces exist, as is the case for most of the sectors considered. However, rank correlations between the two indexes are positive and highly significant (Table 3), being equal to around 0.7 in both manufacturing and services. Hence the two indexes rank sectors in a rather similar way, although there are significant differences, as detailed below by comparing most (least) agglomerated sectors according to the two statistics. Correlations between τ and Moran I's are lower, especially in the service sector, a finding than can be explained on the grounds of the downward bias affecting the Moran spatial autocorrelation index when observations includes a number of outliers, as is the case when strong within region agglomeration forces coexist along with between regions spillovers.

While our measures of local and global agglomeration do not display any significant correlation across sectors, Moran and EG's indexes are significantly and negatively correlated, a feature that we can motivate on the basis of the spurious relationship induced by the simultaneous biases that have been shown to affect the two indexes when the observed spatial pattern results from the joint operation of both strictly local and global externalities.. Of course, correlation across sectors of the measures of within and between regions agglomeration could be an intrinsic feature of the actual regional economies, but economic theory does not provide any indication of the possible existence of such displacement mechanism across short and long range externalities and, as a matter of fact, when we compare our estimates of τ and γ, that do not suffer from such measurement bias, we do not find any evidence of a systematic relation linking the two agglomeration measures.

The pool of most spatially concentrated sectors according to our index does not seem to exhibit specific characteristics. In manufacturing, it includes activities for which natural advantages are important (e.g. shipbuilding, fish, wood and leather products; Table 3). However, it also encompasses activities in which knowledge spillovers are likely to be relevant (musical instruments, optical and photographic instruments). Finally, the most agglomerated sectors also include those manufacturing activities in which the Italian economy is traditionally specialized in and that feature the so-called industrial districts (leather products, textile, footwear and ceramic tiles).

A comparison with the ranking obtained according to the two versions of the EG indexes considered, shows that there is limited concordance. In particular, the latter seem to underestimate spatial concentration for those sectors in which natural advantages and knowledge spillovers play a central role.

The most dispersed manufacturing industries include those with a low value-to-weight ratio (e.g. concrete, paper and printing, metal products) and also some technologically advanced activities (see Table 4). Concordance with the ranking based on EG's indexes is much higher in this case.

In the service sector highly spatially concentrated activities are dominated by the presence of transport-related industries (Table 5). The same group also includes activities like hotels and camping sites that are obviously connected to tourism. For both transport and tourism-related industries, it is likely that natural advantages play a central role in explaining within agglomeration. As for the most agglomerated industries in manufacturing, service sectors that are at the top of the ranking according to γ_W are quite different from those selected through γ_{EG1}.

The service sectors that more closely follow the spatial distribution of aggregate economic activity include those industries for which proximity to the sources of local demand is crucial (Table 6). These are monetary and other financial intermediation services, bars, various repair and maintenance activities, retail sale of different goods including pharmaceuticals, and medical and business services.

The group of the most polarized manufacturing activities includes several industries that are also at the top of the ranking according to γ_W and for which it is likely that natural advantages are important (e.g. shipbuilding, fish, wood and leather products, see Table 7). We interpret this evidence as showing that abundant resource endowments associated to specific areas extend their benefits beyond the borders of a specific local labour market. This pattern may also be at work for some industrial sectors for which knowledge spillovers are important (e.g. the weapons and ammunition, tile and footwear sectors). As far as the set of less polarized industrial sectors is concerned, it is difficult to detect specific patterns (Table 8).

Service activities displaying the highest propensity to cluster in nearby regions include again those with natural advantages (e.g. hotels and camping sites; see Table 9). However, there are many other activities in this group that are among the least geographically concentrated and for which market potential is probably the main explanation for their propensity to cluster in space. Again the composition of less polarized services is more difficult to interpret (see Table 10).

6. Comparing agglomeration patterns in the manufacturing and service sectors.

Our data allow us to compare spatial concentration between manufacturing and services. Before carrying out this comparison, it should be kept in mind that the agglomeration indexes analyzed in this paper measure spatial concentration and polarization in relative terms, i.e. they assess concentration beyond the levels observed for aggregate economic activity.[13]

Several manufacturing goods are shipped long distances with respect to where they are produced. Moreover, in an open economy like Italy, exports represent a large share of total demand in many industrial sectors. As a consequence, one can expect that industrial activities are free to locate their plants in order to exploit natural advantages or knowledge spillovers without the need of being close to sources of local demand. However, several service activities are in fact non-tradable and therefore transaction

[13] Differences between absolute and relative spatial concentration indexes are illustrated by Amiti (1997) and Haaland et al (1999).

costs rapidly increase with distance. Hence, a priori these activities are expected to follow the spatial distribution of aggregate economic activity. But services also encompass some highly knowledge-intensive sectors for which the propensity to agglomerate could be stronger than the benefits obtained from proximity to local sources of demand. Thus, it could be that at least a proportion of service sectors may display relatively high values of spatial agglomeration. In any case, we expect that the latter should be higher in manufacturing than the levels observed for services.

It is more difficult to compare a priori plants' propensity to cluster in nearby regions in manufacturing and services. Differences in this propensity may depend on the spatial range of knowledge spillovers and natural advantages as well as on many other factors. In any case, values for τ are compared in the two fields of economic activity.

Evidence reported in Table 1 clearly confirm that manufacturing activities are much more spatially concentrated than are service activities. The median and the mean of γ_w in manufacturing are much higher than the corresponding statistics in the service sectors. The null hypothesis of the equality of means and medians between the two populations can be rejected at the 99 per cent level of confidence.

We replicate Figure 9 on a finer scale by dropping some sectors displaying extreme values i.e. those with $\gamma_w > .002$. This new evidence confirms that industrial activities are much more spatially concentrated than service sectors (Figure 10).

As regards polarization, we find the opposite result in that it is higher for the service sectors (see Table 2). The null hypothesis of equality of means cannot be rejected but a corresponding test on the difference between medians of τ clearly indicates that this difference is statistically different from zero.

7. Concluding remarks

In this paper, a new method to measure spatial concentration and polarization across industries is proposed. Its main advantage over previous indicators is that of measuring the two types of agglomeration simultaneously allowing for the coexistence of both strictly local and global spatial externalities. The proposed statistical methodology is based on a slightly modified spatial error components (SEC) model, a specification that has received some attention in the spatial econometric literature and whose choice can be motivated as a linear approximation to the solution of a discrete choice problem involving plant location. Having dealt with parameter estimation and hypothesis testing, we have discussed the statistical properties of the most common agglomeration and polarization indexes under the assumption that the data are generated by a SEC process with the proposed features. As expected, it is shown that both EG's and Moran's I indexes are biased when the process includes both within and between regions agglomerating forces. It is also shown how the interplay of the biases affecting EG's and I indexes can induce a spurious correlation between two statistics when they are evaluated over a set of sectors with varying intensities of local/global spillovers.

The proposed methodology is implemented using Italian census data on 103 and 81 manufacturing and service sectors, respectively. A graphical analysis of the spatial pattern of plant location across Italian LLMSs provided some preliminary evidence of the joint operating of strictly local agglomerating forces (resulting in the presence of outliers in the spatial plant distribution) and more wide-ranging externalities, inducing an underlying smooth trend of economic activity across space. Such preliminary evidence was subsequently confirmed by econometric estimates of the proposed measures of within and between markets agglomeration. It turns out that both spatial concentration and polarization are jointly quite well spread across Italian industries, most sectors displaying significant values of both measures. Manufacturing activities appear to be more spatially concentrated than non-manufacturing industries, while the opposite result is obtained as far as polarization is concerned. Comparing our estimates with the evidence provided by standard agglomeration and polarization indexes, we find some empirical support for the actual existence of the biases predicted under our SEC model. While a broad agreement between the results obtained implementing the proposed indexes and their standard counterparts is found, some significant differences are uncovered as well. In any case, avoiding the biases induced by the coexistence of short and long range agglomeration patterns increases comparability of results across sectors.

TABLES AND FIGURES

Table 1 Descriptive statistics on agglomeration

The table reports descriptive statistics for our within and between agglomeration indexes, respectively γ_W and τ, Ellison and Glaeser's γ_{EG1}, based on employees and γ_{EG2} based on plant counts, Moran's I, and δ, the parameter measuring the effects of distance on between regions spillovers. All these indexes are computed on Italian census data in 2001. Geographical units are defined by the 686 local labour systems. We use a three digit sectoral classification reporting 102 industries in manufacturing and 81 in services.

Variable	N	Mean	Std Dev	Median	Minimum	Maximum
Manufacturing						
γ_W	102	0.02033	0.06518	0.00436	0.0002393	0.54880
γ_{EG1}	102	0.02336	0.04438	0.00945	0.0009420	0.25736
γ_{EG2}	102	0.04226	0.13294	0.01248	0.00190	1.28080
τ	102	2.24617	7.89646	0.25470	0	58.34990
I	102	0.01904	0.02083	0.01549	-0.00472	0.12293
δ	102	0.21335	0.19028	0.12406	0.0101	0.91355
Services						
γ_W	81	0.00670	0.02663	0.0004875	0.0000398	0.19705
γ_{EG1}	81	0.01380	0.05601	0.00297	0.0001081	0.49492
γ_{EG2}	81	0.02191	0.05715	0.00415	-0.00109	0.34644
τ	81	4.92584	17.12187	0.72194	0	115.82920
I	81	0.01538	0.01743	0.01019	-0.00593	0.09338
δ	81	0.3090	0.2721	0.2586	0.0173	1.7682

Table 2. Cross-sectional rank correlations between different agglomeration indexes

	γ_w	γ_{EG1}	γ_{EG2}	τ	I
		Manufacturing			
γ_w	1				
γ_{EG1}	0.67974 (<.0001)	1			
γ_{EG2}	0.56631 (<.0001)	0.73582 (<.0001	1		
τ	-0.03369 (0.7368)	-0.14380 (0.1493	-0.12339 (0.2166	1	
I	-0.05788 (0.5633)	-0.28431 (0.0038)	-0.46364 (<.0001)	0.42274 (<.0001)	1
		Services			
γ_w	1				
γ_{EG1}	0.66005 (<.0001)	1			
γ_{EG2}	0.52075 (<.0001)	0.73374 (<.0001)	1		
τ	-0.22100 (0.0474)	-0.24776 (0.0257)	-0.20185 (0.0707)	1	
I	-0.11138 (0.3222)	-0.37078 (0.0007)	-0.39146 (0.0003)	0.29207) (0.0082)	1

(1) Spearman correlation coefficients (p-values are reported in brackets).

Table 3 – The 15 most agglomerated sectors according to γ_W: manufacturing

NACE	Sector	γ_W	γ_{EG1}	γ_{EG2}	Rank γ_W	Rank γ_{EG1}	Rank γ_{EG2}	Rank τ
231	Manufacture of coke oven products	0.549	0.014	1.281	1	36	1	100
201	Sawmilling and planing of wood, impregnation of wood	0.295	0.009	0.013	2	58	49	4
351	Building and repairing of ships and boats	0.161	0.016	0.028	3	28	28	5
191	Tanning and dressing of leather	0.139	0.206	0.212	4	3	3	9
263	Manufacture of ceramic tiles and flags	0.128	0.257	0.336	5	1	2	8
296	Manufacture of weapons and ammunition	0.086	0.157	0.139	6	5	5	12
160	Manufacture of tobacco products	0.046	0.054	0.015	7	10	43	92
334	Manufacture of optical instruments, photographic equipment	0.044	0.019	0.047	8	21	19	21
363	Manufacture of musical instruments	0.042	0.056	0.131	9	9	6	69
152	Processing and preserving of fish and fish products	0.039	0.020	0.019	10	20	35	2
171	Preparation and spinning of textile fibres	0.026	0.223	0.088	11	2	10	97
313	Manufacture of insulated wire and cable	0.025	0.012	0.007	12	39	76	62
193	Manufacture of footwear	0.021	0.038	0.035	13	14	23	14
181	Manufacture of leather clothes	0.019	0.035	0.052	14	15	17	39
177	Manufacture of knitted and crocheted articles	0.018	0.011	0.016	15	42	40	7

Table 4 – The 15 least agglomerated sectors according to γ_W: manufacturing

NACE	Sector	γ_W	γ_{EG1}	γ_{EG2}	Rank γ_W	Rank γ_{EG1}	Rank γ_{EG2}	Rank τ
212	Manufacture of articles of paper and paperboard	0.001	0.003	0.003	88	90	95	22
295	Manufacture of other special purpose machinery	0.001	0.005	0.003	89	77	93	17
266	Manufacture of articles of concrete, plaster, cement	0.001	0.003	0.005	90	88	81	71
312	Manufacture of electricity distribution and control apparatus	0.001	0.008	0.011	91	60	58	26
182	Manufacture of other wearing apparel and accessories	0.001	0.003	0.005	92	95	82	16
252	Manufacture of plastic products	0.001	0.003	0.002	93	87	101	30
322	Manufacture of television and radio transmitters and apparatus for line telephony and line telegraphy	0.001	0.003	0.030	94	91	25	78
316	Manufacture of electrical equipment n.e.c.	0.001	0.004	0.004	95	82	89	20
331	Manufacture of medical and surgical equipment and orthopaedic appliances	0.001	0.003	0.008	96	89	73	42
292	Manufacture of other general purpose machinery	0.000	0.002	0.003	97	99	97	13
222	Printing and service activities related to printing	0.000	0.007	0.010	98	71	65	77
158	Manufacture of other food products	0.000	0.002	0.002	99	100	100	10
203	Manufacture of builders' carpentry and joinery	0.000	0.003	0.005	100	96	85	1
287	Manufacture of other fabricated metal products	0.000	0.001	0.003	101	102	92	15
281	Manufacture of structural metal products	0.000	0.001	0.002	102	101	102	31

Table 5 – The 15 most agglomerated sectors according to γ_W: services

NACE	Sector	γ_W	γ_{EG1}	γ_{EG2}	Rank γ_W	Rank γ_{EG1}	Rank γ_{EG2}	Rank τ
712	Renting of other transport equipment	0.197	0.005	0.006	30	32	3	3
552	Camping sites, other provision of short-stay accommodation	0.107	0.015	0.012	11	21	2	2
612	Inland water transport	0.088	0.495	0.346	1	1	68	68
852	Veterinary activities	0.040	0.001	0.003	72	48	1	1
622	Non-scheduled air transport	0.013	0.006	-0.001	27	81	66	66
603	Transport via pipelines	0.012	0.001	0.008	65	26	69	69
611	Sea and coastal water transport	0.012	0.032	0.099	5	5	60	60
551	Hotels	0.008	0.012	0.007	16	27	4	4
925	Library, archives, museums, other cultural activities	0.007	0.004	0.008	34	25	42	42
621	Scheduled air transport	0.006	0.108	0.324	2	2	78	78
801	Primary education	0.005	0.025	0.026	7	15	54	54
927	Other recreational activities	0.004	0.010	0.010	21	23	79	79
512	Wholesale of agricultural raw materials, live animals	0.003	0.008	0.012	23	20	80	80
634	Activities of other transport agencies	0.002	0.011	0.025	18	16	76	76
601	Transport via railways	0.002	0.002	0.000	51	80	59	59

Table 6 – The 15 least agglomerated sectors according to γ_W: service

NACE	Sector	γ_W	γ_{EG1}	γ_{EG2}	Rank γ_W	Rank γ_{EG1}	Rank γ_{EG2}	Rank τ
511	Wholesale on a fee or contract basis	0.000	0.001	0.001	67	73	63	53
602	Other land transport	0.000	0.001	0.001	68	68	68	5
505	Retail sale of automotive fuel	0.000	0.002	0.006	69	49	31	49
502	Maintenance and repair of motor vehicles	0.000	0.001	0.004	70	54	37	27
804	Adult and other education	0.000	0.000	0.001	71	81	73	71
741	Legal, account., book-keeping & auditing activities; tax consult.; market research & public op. polling; business & management consulting; holdings	0.000	0.002	0.002	72	48	57	26
651	Monetary intermediation	0.000	0.001	0.000	73	69	76	10
523	Retail sale of pharmaceutical, medical goods, cosmetic	0.000	0.001	0.002	74	63	49	52
554	Bars	0.000	0.001	0.002	75	70	52	8
524	Other retail sale of new goods in specialized stores	0.000	0.001	0.003	76	59	46	25
527	Repair of personal and household goods	0.000	0.000	0.002	77	79	61	48
742	Architectural and engineering activities and related tech. consult.	0.000	0.000	0.001	78	78	70	81
851	Human health activities	0.000	0.001	0.001	79	67	65	28
672	Activities auxiliary to insurance and pension funding	0.000	0.000	0.001	80	80	75	13
930	Other service activities	0.000	0.000	0.002	81	77	50	6

Table 7 – The 15 most agglomerated sectors according to τ: manufacturing

NACE	Sector	τ	Moran's I	Rank τ	Rank Moran's I	Rank γ_W
203	Manufacture of builders' carpentry and joinery	58.350	0.032	1	21	100
152	Processing and preserving of fish and fish products	40.738	0.031	2	22	10
293	Manufacture of agricultural and forestry machinery	27.408	0.022	3	39	18
201	Sawmilling and planing of wood, impregnation of wood	23.318	0.032	4	20	2
351	Building and repairing of ships and boats	16.450	0.023	5	35	3
314	Manufacture of accumulators, primary cells and primary batteries	6.837	0.002	6	84	21
177	Manufacture of knitted and crocheted articles	5.940	0.027	7	28	15
263	Manufacture of ceramic tiles and flags	5.168	0.006	8	66	5
191	Tanning and dressing of leather	4.605	-0.002	9	94	4
158	Manufacture of other food products	3.053	0.075	10	3	99
154	Manufacture of vegetable and animal oils and fats	2.804	0.123	11	1	48
296	Manufacture of weapons and ammunition	2.746	0.009	12	62	6
292	Manufacture of other general purpose machinery	2.054	0.029	13	25	97
193	Manufacture of footwear	2.008	0.039	14	12	13
287	Manufacture of other fabricated metal products	1.485	0.036	15	17	101

Table 8 – The 15 least agglomerated sectors according to τ: manufacturing

NACE	Sector	τ	Moran's I	Rank τ	Rank Moran's I	Rank γ_W
286	Manufacture of cutlery, tools and general hardware	0.021	0.038	88	15	35
202	Manufacture of veneer sheets; manufacture of plywo	0.019	0.016	89	51	37
174	Manufacture of made-up textile articles, except apparel	0.019	0.005	90	70	76
267	Cutting, shaping and finishing of ornamental and building stone	0.018	0.027	91	29	55
160	Manufacture of tobacco products	0.017	0.021	92	41	7
172	Textile weaving	0.011	0.002	93	83	19
335	Manufacture of watches and clocks	0.008	0.008	94	64	32
176	Manufacture of knitted and crocheted fabrics	0.008	0.020	95	44	52
361	Manufacture of furniture	0.008	-0.001	96	93	73
171	Preparation and spinning of textile fibres	0.006	0.000	97	90	11
159	Manufacture of beverages	0.006	0.037	98	16	49
205	Manufacture of other products of wood; manufacture of articles of cork, straw and plaiting materials	0.001	0.018	99	47	61
231	Manufacture of coke oven products	0.001	-0.003	100	97	1
262	Manuf. of non-refractory ceramic goods other than for construction purposes; manuf. of refractory ceramic prod.	0.000	0.014	101	56	33
265	Manufacture of cement, lime and plaster	0.000	0.029	101	24	38

Table 9 – The 15 most agglomerated sectors according to τ: services

NACE	Sector	τ	Moran's I	Rank τ	Rank Moran's I	Rank γ_W
852	Veterinary activities	115.829	0.034	1	8	4
552	Camping sites, other provision of short-stay accommodation	87.127	0.052	2	4	2
712	Renting of other transport equipment	52.136	0.026	3	19	1
551	Hotels	28.244	0.030	4	13	8
602	Other land transport	19.871	0.025	5	21	68
930	Other service activities	18.204	0.004	6	59	81
641	Post and courier activities	5.809	0.000	7	68	46
554	Bars	4.448	0.004	8	58	75
514	Wholesale of household goods	3.419	0.005	9	55	48
651	Monetary intermediation	3.186	0.027	10	18	73
671	Activities auxiliary to financial intermediation, except insur. & pens. funding	2.902	0.014	11	35	63
803	Higher education	2.875	0.000	12	69	34
672	Activities auxiliary to insurance and pension funding	2.595	0.017	13	32	80
516	Wholesale of machinery, equipment and supplies	2.543	0.024	14	24	59
926	Sporting activities	2.426	0.093	15	1	53

Table 10 – The 15 least agglomerated sectors according to τ: services

NACE	Sector	τ	Moran's I	Rank τ	Rank Moran's I	Rank γ_W
711	Renting of automobiles	0.271	0.000	67	66	18
612	Inland water transport	0.203	-0.001	68	72	3
603	Transport via pipelines	0.190	0.001	69	64	6
900	Sewage and refuse disposal, sanitation and similar activities	0.150	0.007	70	48	44
804	Adult and other education	0.127	0.010	71	42	71
632	Other supporting transport activities	0.125	0.003	72	61	32
921	Motion picture and video activities	0.124	-0.002	73	78	19
555	Canteens and catering	0.119	0.016	74	33	35
633	Activities of travel agencies and tour operators; tourist assistance activities n.e.c.	0.083	-0.001	75	70	50
634	Activities of other transport agencies	0.069	0.001	76	65	14
713	Renting of other machinery and equipment	0.061	0.007	77	49	51
621	Scheduled air transport	0.055	-0.001	78	73	10
927	Other recreational activities	0.002	0.031	79	11	12
512	Wholesale of agricultural raw materials, live animals	0.000	0.007	80	50	13
742	Arch. and engineering activities and related technical consultancy	0.000	0.022	81	25	78

Figure 1. Two examples of plant location patterns

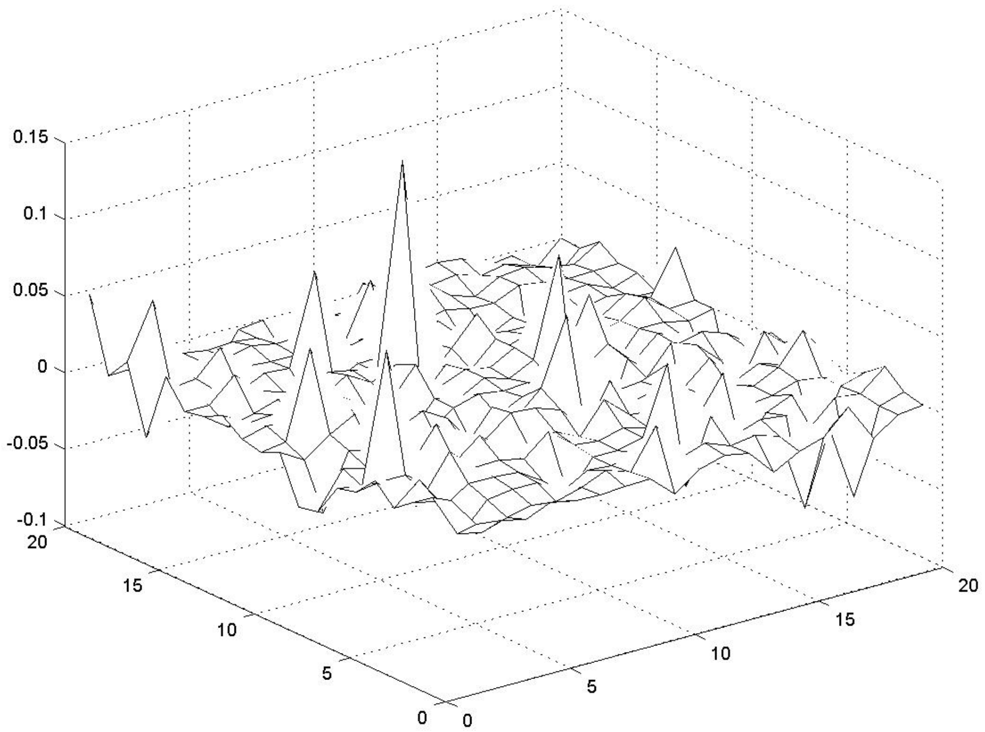
Industry I1

1			1		
					1
	1			1	
		15			
1				1	
1		1			1

Industry I2

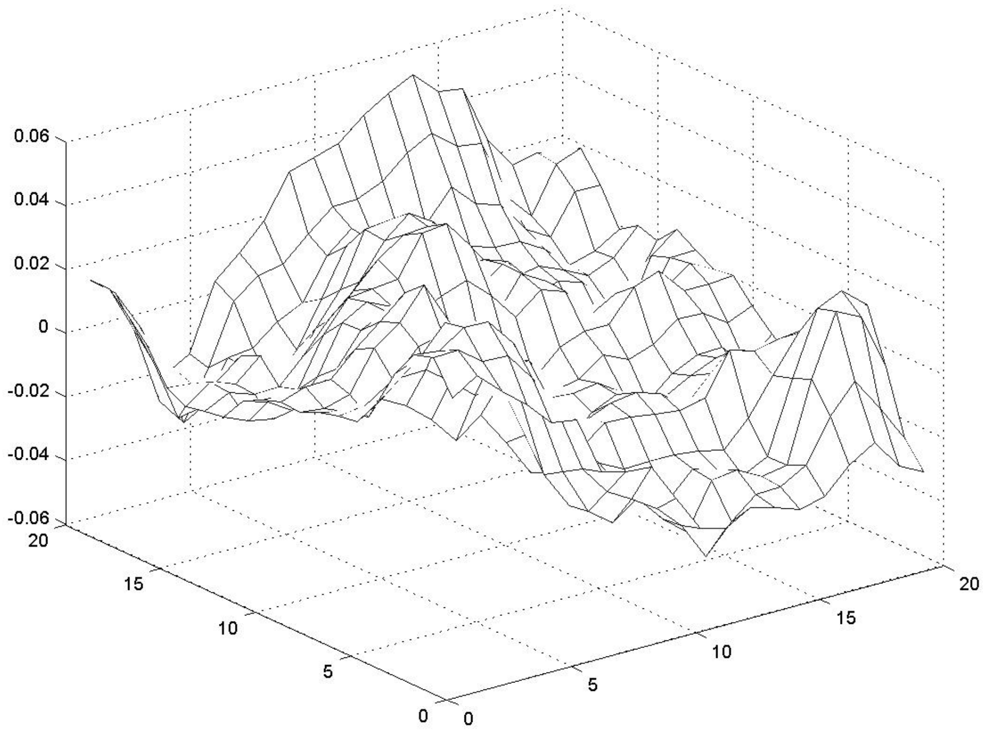
1		1			
	1	3	1		
	2	10	1		
	1	1	3		

Figure 2. Simulated pattern of the local SEC model component



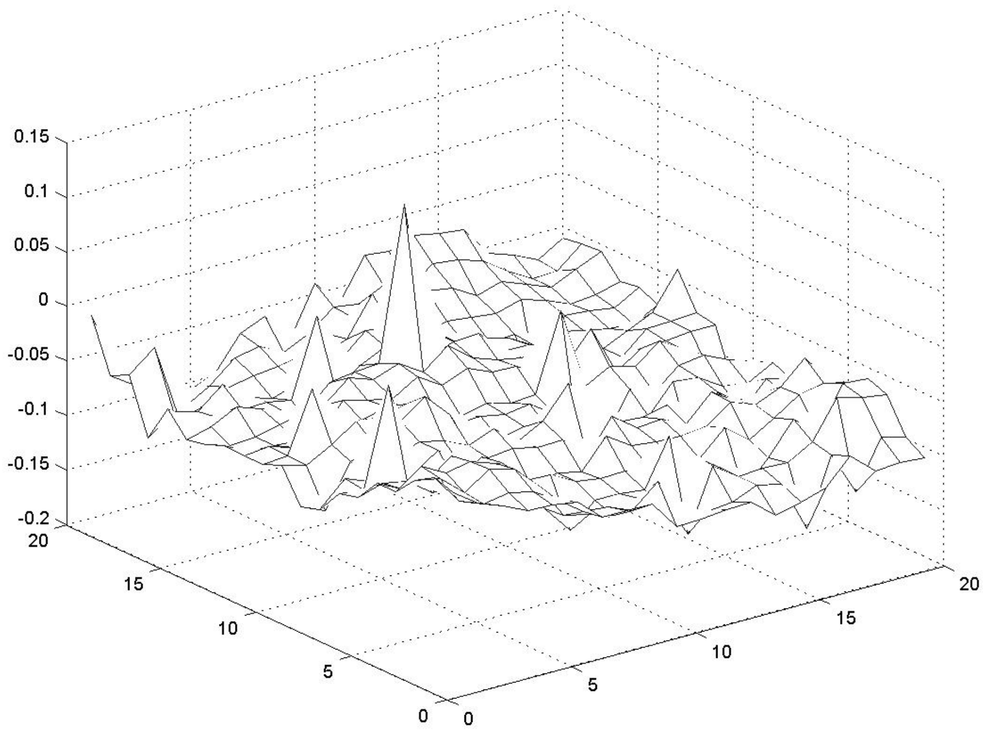
This graph displays simulated series for the η component, using a 20x20 lattice as spatial reference. The x values required to simulate the within component were drawn randomly from a lognormal distribution with parameters (0;2). As a consequence, the spikes, corresponding to locations with abnormally high (or low) plant concentration in a given industry, appear to be scattered randomly across the plane.

Figure 3. Simulated pattern of the global SEC model component



This graph displays simulated series respectively for the ζ model component, using a 20x20 lattice as spatial reference. The between component was simulated computing spatial weights on the basis of Euclidean distances between cells, assuming a within cell distance d_{ii}=0.5, and setting a value of δ=0.5. The slow weights decay implied by such a parameter value imposes a high persistence on shocks across space, a feature that induces the smooth, trending behaviour observed in this figure.

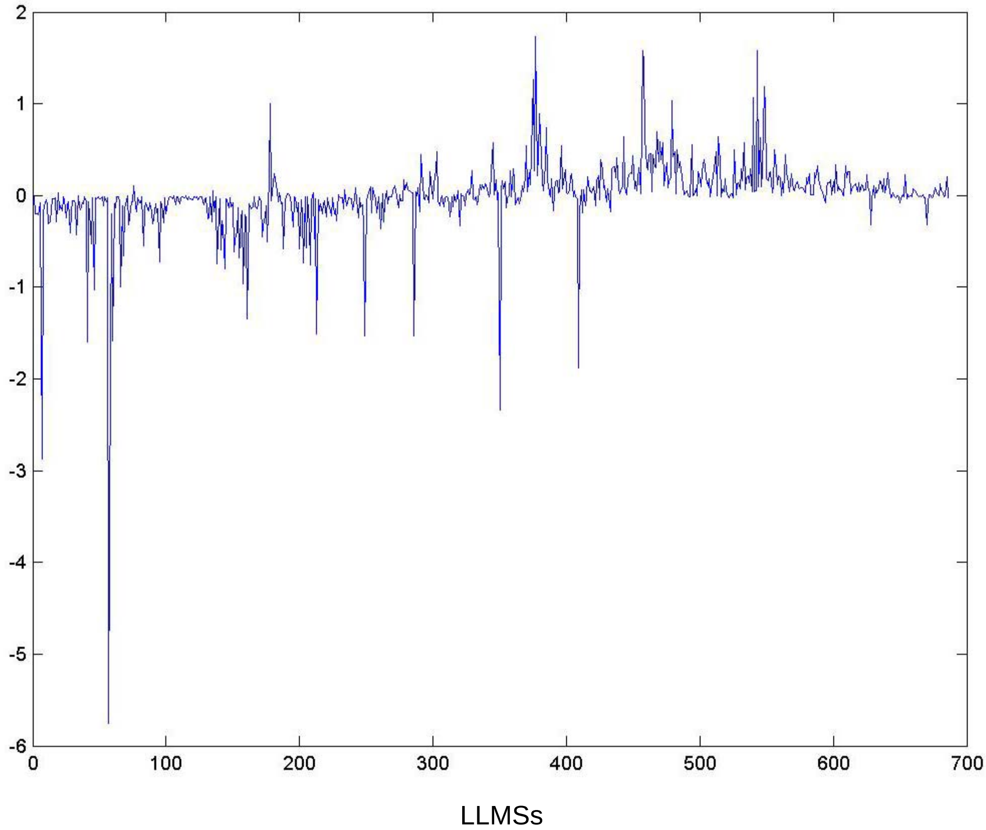
Figure 4. Combined local and global SEC model components



This graph depicts the combination of the two components, i.e. a realization of the (s-x) process, assuming a variance ratio τ about equal to 9.

Figure 5. Relative cross-sectional distribution of plants in the “Manufacture of vegetable and animal oils and fats” sector

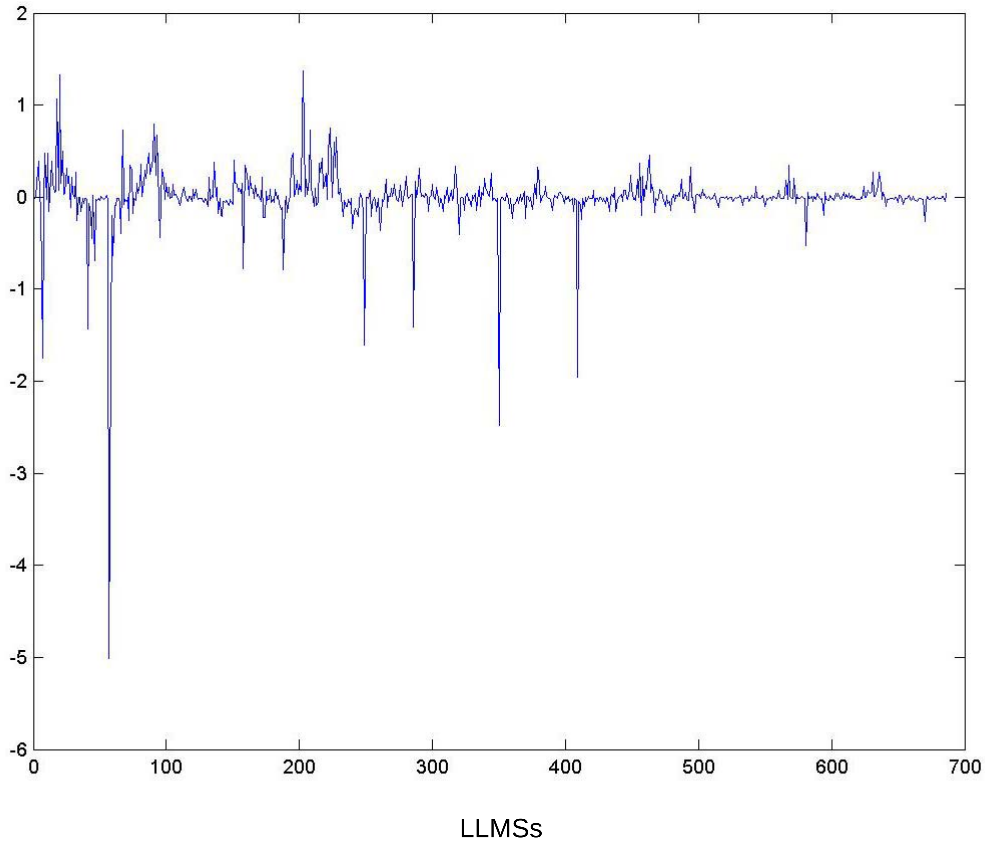
(percentage points)



For each of the 686 Italian LLMS the graph plots the value of (*s*-*x*) computed on the basis of local shares of the number of plants recorded on 2001 national Census of industry and services.

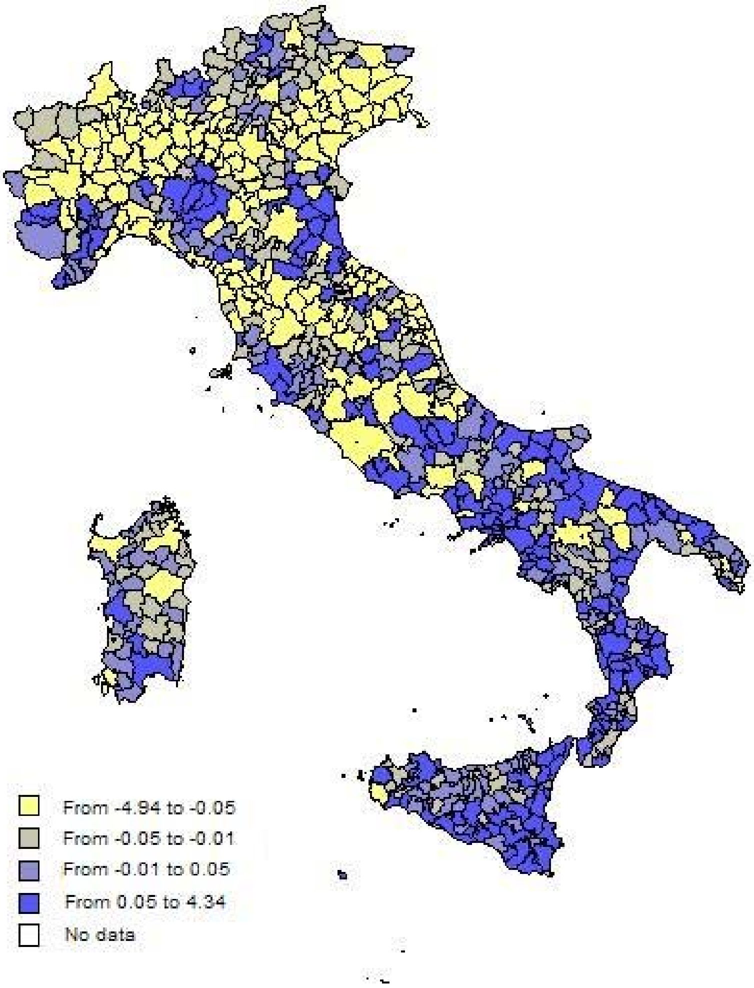
Figure 6. Relative cross-sectional distribution of plants in the “Manufacture of machine-tools” sector

(percentage points)



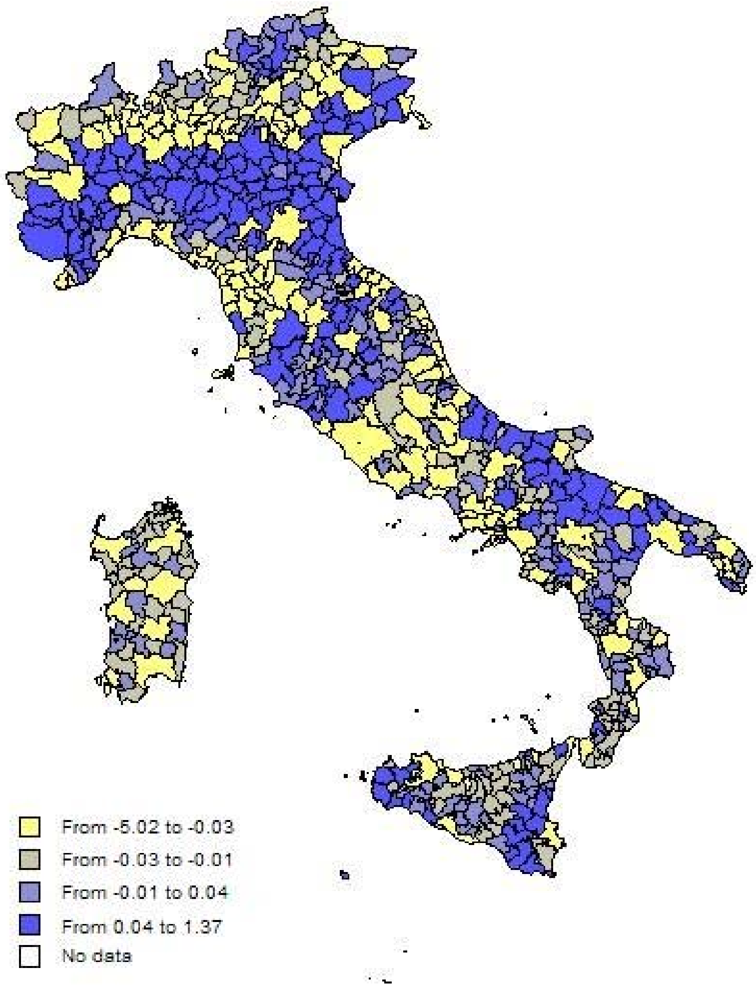
For each of the 686 Italian LLMS the graph plots the value of (*s*-*x*) computed on the basis of local shares of the number of plants recorded on 2001 national Census of industry and services.

Figure 7. Relative spatial distribution of plants in the “Manufacture of vegetable and animal oils and fats” sector:



For each of the 686 Italian LLMS the map displays the value of (*s*-*x*) computed on the basis of local shares of the number of plants recorded on 2001 national Census of industry and services.

Figure 8. Relative spatial distribution of plants in the “Manufacture of machine-tools” sector:



For each of the 686 Italian LLMS the map displays the value of (*s*-*x*) computed on the basis of local shares of the number of plants recorded on 2001 national Census of industry and services.

Figure 9.

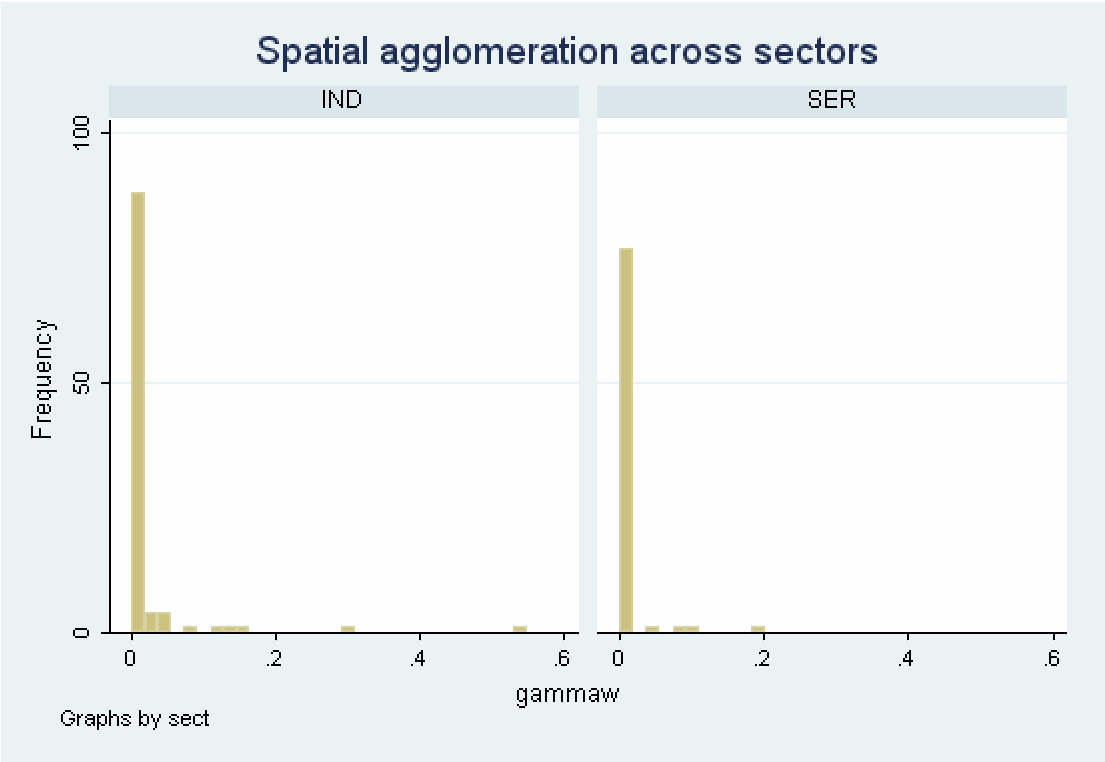
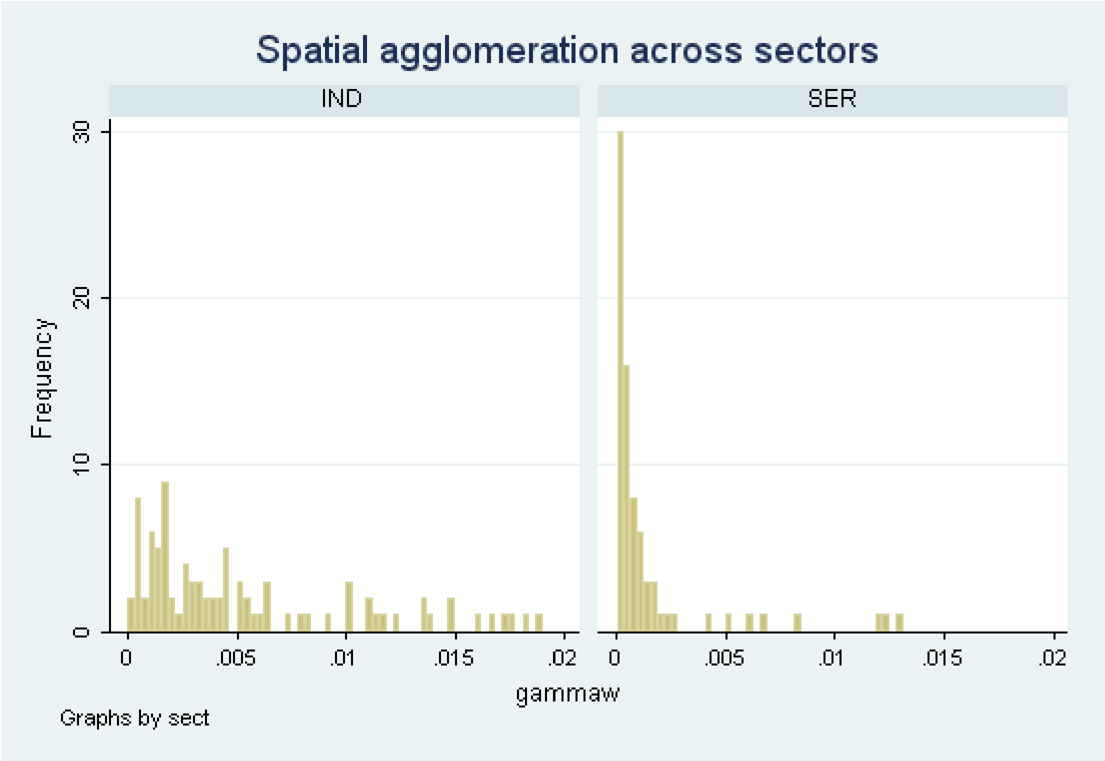

Spatial agglomeration across sectors
IND
SER
Frequency
0
50
100
0
.2
.4
.6
gammaw
Graphs by sect

Figure 10.



Spatial agglomeration across sectors
IND
SER
Frequency
0
10
20
30
0
.005
.01
.015
.02
gammaw
Graphs by sect

APPENDIX A: A discrete choice motivation of the methodological approach

Profits accruing to firm *j*, belonging to sector *q*, from the decision to set up a plant on the *i*-th location, (i=1,…,M), are assumed to be a function of a set of observable features of each site μ_i, of an unobservable random area/sector effect η_i, representing the general suitability of location *i* for industry *q*, and of an idiosyncratic random disturbance ε_{ij} capturing the specific appropriateness of the given site for firm *j*, due to firm specific characteristics. Assuming a log-linear specification and letting π denote profits, we have:

$$\log \pi_{ij} = \mu_i + \eta_i + \varepsilon_{ij}, \qquad (i=1,\dots,M) \tag{a1}$$

In selecting plant location, the firm faces the following discrete choice problem:

$$\max_{i \in \{1,\dots,M\}} \log \pi_{ij}. \tag{a2}$$

It is well known (McFadden, 1974) that, when the distribution of the idiosyncratic shock is of the *extreme value* type, the solution of (a2) implies the following conditional probabilities of a (sector *q*) firm locating on site *i* :

$$p_i \mid \mu_i + \eta_i = \frac{\exp(\mu_i + \eta_i)}{\sum_{i=1}^{M} \exp(\mu_i + \eta_i)} \tag{a3}$$

defining a multinomial logistic density function.

Guimarães et al. (2004) show how, assuming that $\exp(\eta_i)$ has a gamma(δ^1, δ^1) distribution and marginalizing with respect to η, the joint sample density function is Dirichlet-Multinomial, a result that is subsequently utilized to obtain maximum likelihood estimators of unknown model parameters.

Specification (a1) can be extended by allowing profits to be affected also by a third unobservable stochastic factor (ζ_i). This represents a random disturbance originating both within and outside the area and propagating across space according to a given diffusion mechanism, an example being demand (market potential) shocks transmitted via trade across areas, as recently hypothesized by Head and Meyer (2002) in a study of the location of foreign firms in the European Union. In this extended setting, we have:

$$\log \pi_{ij} = \mu_i + \eta_i + \zeta_i + \varepsilon_{ij} \tag{a4}$$

yielding, under the same distributional assumptions for ε_{ij} as above, the following conditional logistic probabilities:

$$p_i \mid \mu_i + \zeta_i + \psi_i = \frac{\exp(\mu_i + \eta_i + \zeta_i)}{\sum_{i=1}^{M} \exp(\mu_i + \eta_i + \zeta_i)}. \tag{a5}$$

In this case, the inclusion of a third unobservable stochastic effect, that is *cross-sectionally* correlated due to the underlying diffusion mechanism, makes the analytical derivation of the sample likelihood function overly complex. Assuming a linear approximation to the logistic probabilities given by (a5), i.e. a *linear probability model*, would yield in this case:

$$p_i \mid \mu_i + \zeta_i + \psi_i = \mu_i + \eta_i + \zeta_i \tag{a6}$$

that, by replacing probabilities with observed frequencies and substituting x_i for μ_i, yields the error component specification given in expression (3) in the text.

APPENDIX B: Maximum likelihood estimation

First order conditions for the maximization of (15) in the text are:

$$\frac{\partial L}{\partial \lambda} = -\frac{M}{2\lambda} + \frac{1}{2\lambda^2}(s-x)'\Omega^{-1}(s-x) = 0 \tag{a7}$$

$$\begin{aligned} \frac{\partial L}{\partial \tau} &= -\frac{1}{2} tr\left\{\Omega^{-1}\frac{\partial \Omega}{\partial \tau}\right\} + \frac{1}{2\lambda}(s-x)'\Omega^{-1}\frac{\partial \Omega}{\partial \tau}\Omega^{-1}(s-x) = \\ &= -\frac{1}{2} tr\left\{\Omega^{-1}W^{*}W^{*}{}'\right\} + \frac{1}{2\lambda}(s-x)'\Omega^{-1}W^{*}W^{*}{}'\Omega^{-1}(s-x) = 0 \end{aligned} \tag{a8}$$

A closed form solution exists only for (a7) from which it is possible to derive the following expression for the maximum likelihood estimator of λ:

$$\hat{\lambda} = \frac{1}{M}(s-x)'\Omega^{-1}(s-x). \tag{a9}$$

Substituting the above expression in (15), the following concentrated log-likelihood ensues:

$$L_c(\tau) = const. - \frac{M}{2}\log[(s-x)'\Omega^{-1}(s-x)] - \frac{1}{2}\log(|\Omega|) \tag{a10}$$

that upon maximization, to be achieved by means of standard iterative techniques, gives the ML estimator of τ.

Finally, ML estimators of the original model parameters γ and ψ can be obtained by solving equations (12) and (13) with respect to those variables, yielding:

$$\hat{\gamma} = \hat{\lambda} / \kappa_x \tag{a11}$$

$$\hat{\psi} = \hat{\tau}\hat{\gamma} / (\kappa_w / \kappa_x). \tag{a12}$$

The above derivations assume that the parameter δ, measuring the rate of spatial decay of global spillovers, is a constant known to researchers (perhaps from previous studies). In many empirical applications this situation will not hold, and as a consequence it could be necessary to treat δ as an unknown parameter to be estimated from the data. The first order condition for the maximum of the log-likelihood function with respect to δ is highly non linear and does not allow for a closed form solution. In consequence the parameter will have to be estimated by iterative numeric techniques, a task that can be performed, along with the estimation of τ, on the basis of the concentrated log-likelihood function given by (a10).

APPENDIX C: Derivation of the LM test statistics.

The expression of the unrestricted information matrix is the following:

$$J = \begin{bmatrix} \frac{M}{2\lambda^2} & \frac{1}{2\lambda} tr\{\Omega^{-1}W^*W^{*}{}'\} \\ \frac{1}{2\lambda} tr\{\Omega^{-1}W^*W^{*}{}'\} & \frac{1}{2} tr\{\Omega^{-1}W^*W^{*}{}'\Omega^{-1}W^*W^{*}{}'\} \end{bmatrix} \qquad \text{(a13)}$$

We first proceed to derive the LM test for the H_0^2 null hypothesis. The relevant element of the score vector and the information matrix, evaluated under H_0^2, becomes:

$$\tilde{d}_\tau = -\frac{1}{2} tr\{\Xi^{*-1}W^*W^{*}{}'\} + \frac{1}{2\lambda}(s-x)'\Xi^{*-1}W^*W^{*}{}'\Xi^{*-1}(s-x) = 0 \qquad \text{(a14)}$$

And:

$$\tilde{J} = \begin{bmatrix} \frac{M}{2\lambda^2} & \frac{1}{2\lambda} tr\{\Xi^{*-1}W^*W^{*}{}'\} \\ \frac{1}{2\lambda} tr\{\Xi^{*-1}W^*W^{*}{}'\} & \frac{1}{2} tr\{\Xi^{*-1}W^*W^{*}{}'\Xi^{*-1}W^*W^{*}{}'\} \end{bmatrix} \qquad \text{(a15)}$$

By making the positions $tr\{\Xi^{*-1}W^*W^{*}{}'\} = T_1$ and $tr\{\Xi^{*-1}W^*W^{*}{}'\Xi^{*-1}W^*W^{*}{}'\} = T_2$, the expression for the test statistic immediately follows thus:

$$LM_\tau = \left[\frac{1}{\lambda}(s-x)'\Xi^{*-1}W^*W^{*}{}'\Xi^{*-1}(s-x) - T_1\right]^2 / 2\left(T_2 - \frac{(T_1)^2}{M}\right) \qquad \text{(a16)}$$

To derive the LM test of H_0^1 it is more convenient to start from the initial model parameterization, based on the couple $[\gamma, \psi]$ instead of $[\lambda, \tau]$. The expression of the log-likelihood in this case becomes:

$$L(\gamma;\psi) = -\frac{M}{2}\log(\pi) - \frac{1}{2}\log(|\Sigma|) - \frac{1}{2}(s-x)'\Sigma^{-1}(s-x)\,. \qquad \text{(a17)}$$

and the relevant component of the score is equal to:

$$d_\gamma = \frac{\partial L}{\partial \gamma} = -\frac{1}{2} tr\left\{\Sigma^{-1}\frac{\partial \Sigma}{\partial \gamma}\right\} + \frac{1}{2}(s-x)'\Sigma^{-1}\frac{\partial \Sigma}{\partial \gamma}\Sigma^{-1}(s-x) =$$
$$= -\frac{1}{2} tr\left\{\Sigma^{-1}\Xi\right\} + \frac{1}{2}(s-x)'\Sigma^{-1}\Xi\Sigma^{-1}(s-x) \tag{a18}$$

Applying the Harville (1977) formula, we derive the following equation:

$$J_{rs} = \frac{1}{2} tr\left[\Sigma^{-1}\frac{\partial \Sigma}{\partial \theta_r}\Sigma^{-1}\frac{\partial \Sigma}{\partial \theta_s}\right] \tag{a19}$$

Individual entries of *J* can be shown to have expressions:

$$J_{\gamma\gamma} = \frac{1}{2} tr\left[\Sigma^{-1}\Xi\Sigma^{-1}\Xi\right] \tag{a20}$$

$$J_{\gamma\psi} = \frac{1}{2} tr\left[\Sigma^{-1}\Xi\Sigma^{-1}(WW')\right] \tag{a21}$$

$$J_{\psi\psi} = \frac{1}{2} tr\left[\Sigma^{-1}(WW')\Sigma^{-1}(WW')\right] \tag{a22}$$

A restricted version of the score and information matrix can subsequently be obtained by setting:

$$\Sigma = \tilde{\Sigma} = \psi WW' \tag{a23}$$

In turn, this transformation yields:

$$\tilde{d}_\gamma = -\frac{1}{2\psi} tr\left\{(WW')^{-1}\Xi\right\} + \frac{1}{2\psi^2}(s-x)'(WW')^{-1}\Xi(WW')^{-1}(s-x) \tag{a24}$$

And:

$$\tilde{J}_{\gamma\gamma} = \frac{1}{2\psi^2} tr\left[(WW')^{-1}\Xi(WW')^{-1}\Xi\right] \tag{a25}$$

$$\tilde{J}_{\gamma\psi} = \frac{1}{2\psi^2} tr\left[(WW')^{-1}\Xi\right] \tag{a26}$$

$$\tilde{J}_{\psi\psi} = \frac{M}{2\psi^2} \tag{a27}$$

Finally, making the positions $tr\left\{(WW')^{-1}\Xi\right\} = T_3$ and $tr\left\{(WW')^{-1}\Xi(WW')^{-1}\Xi\right\} = T_4$, the relevant element of the partitioned inverse of *J* will be:

$$\tilde{J}^{\gamma\gamma} = 2\psi^2\left[T_4 - \frac{(T_3)^2}{M}\right]^{-1} \tag{a28}$$

and the LM test statistics expression will take the following form:

$$LM_{\gamma} = \left[\frac{1}{\psi}(s-x)'(WW')^{-1}\Xi(WW')^{-1}(s-x) - T_3 \right]^2 / 2\left(T_4 - \frac{(T_3)^2}{M} \right) \qquad \text{(a29)}$$

APPENDIX D: Simulation exercise.

The design of the Monte Carlo experiment is the following. Q instances of the SEC process are assumed to be observed simultaneously on the spatial reference set provided by a 20x20 regular lattice. Each instance of the process assumes the same realization of the x vector, drawn from a lognormal distribution, and is denoted by a different parameter vector $\theta_h = [\gamma_h, \psi_h]'$, h=1,...,Q, whose elements are drawn randomly and independently from uniform distributions of ranges $[0, U_{\gamma}]$ and $[0, U_{\psi}]$. A fixed value of parameter $\delta=\delta^*$ is assumed for all $h \in \{1,2,..,Q\}$. Letting $\boldsymbol{s}^h = [s_1^h, s_2^h, ..., s_M^h]'$ denote a realization of the h-th SEC process on the M=20^2 sites, N independent replications of $\boldsymbol{s}^h$, h=1,...,Q, were simulated by randomly drawing components η^h and ζ^h from two independent normal distributions with zero mean and covariance matrices $\Sigma_{\eta} = \gamma_h \Xi$ and $\Sigma_{\zeta} = \psi_{\zeta} WW'$. On each replication of $\boldsymbol{s}^h$ the value of the EG's and Moran's statistics were computed. Considering all Q processes jointly, at iteration $k \in \{1,2,..,N\}$ the procedure yields vectors $\gamma_{EG}(k) = [\gamma_{EG}^1(k), \gamma_{EG}^2(k), ..., \gamma_{EG}^Q(k)]'$ and $I(k) = [I^1(k), I^2(k), ..., I^Q(k)]'$ and a value $\rho(k)$ of the correlation coefficient between $\gamma_{EG}(k)$ and $I(k)$.

To study the influence of different levels of dispersion in the two model parameters, the simulation experiment was repeated considering increasing values of U_{γ} and U_{ψ}, and Table D1 reports the average value of $\rho(k)$ across replications, $\overline{\rho} = N^{-1}\sum_k \rho(k)$, for Q=N=100 and for different combinations of U_{γ} and U_{ψ}. Three different values of δ were considered, allowing for different degrees of spatial diffusion of disturbances in the between component.

Table D1. Correlation between the EG's and Moran's indexes within the SEC model: evidence from a simulation exercise

Parameter range					
γ	ψ				
	0 - 0.0001	0 - 0.001	0 - 0.01	0 - 0.1	0 - 1
			δ=0.5		
0 - 0.0001	0.578	0.606	0.547	0.462	0.468
0 - 0.001	-0.228	0.523	0.602	0.496	0.482
0 - 0.01	-0.349	-0.201	0.620	0.598	0.509
0 - 0.1	-0.162	-0.297	-0.131	0.599	0.614
0 - 0.25	-0.176	-0.273	-0.314	0.303	0.645
			δ=1		
0 - 0.0001	0.678	0.643	0.592	0.503	0.502
0 - 0.001	-0.162	0.643	0.668	0.541	0.527
0 - 0.01	-0.382	-0.149	0.699	0.622	0.551
0 - 0.1	-0.196	-0.325	-0.105	0.661	0.633
0 - 0.25	-0.180	-0.308	-0.357	0.393	0.682
			δ=2		
0 - 0.0001	0.442	0.233	0.197	0.126	0.163
0 - 0.001	0.458	0.437	0.272	0.153	0.157
0 - 0.01	-0.425	0.474	0.491	0.228	0.190
0 - 0.1	-0.278	-0.358	0.505	0.418	0.238
0 - 0.25	-0.271	-0.415	0.070	0.479	0.297

References

Amiti, M. (1997), "Specialisation patterns in Europe", CEPR, Discussion Paper n. 363.

Anselin, L. (2001),"Rao's Sore Test in Spatial Econometrics", *Journal of Statistical Planning and Inference*, 97(1), pp. 113-139.

Anselin, L. (2003), "Spatial Externalities, Spatial Multipliers, and Spatial Econometrics", *International Regional Science Review*, 26(2), pp. 153-166.

Arbia, G. (2001), "The Role of Spatial Effects in the Empirical Analysis of Regional Concentration", in *Journal of Geographical Systems,* 3, pp. 271-281.

Baldwin R. E. and T. Okubo (2006), "Heterogeneous Firms, Agglomeration and Economic Geography: Spatial Selection and Sorting", *Journal of Economic Geography*, 6, pp. 323-346.

Belleflamme, P., P. Picard, and J. F. Thisse (2000), "An Economic Theory of Regional Clusters", *Journal of Urban Economics*, 48, pp. 158-84.

Cliff, A. D. and J. K. Ord (1981), *Spatial Processes: Models and Applications*. London, Pion.

Dempster A. P., N. M. Laird and D. Rubin D. (1977), "Maximum Likelihood from Incomplete Data via the EM algorithm", *Journal of the Royal Statistical Society* B, 39(1), pp. 1-38.

Devereux M. P., R. Griffith, H. Simpson (1999), "The Geographic Distribution of Production Activity in the United Kingdom", The Institute for Fiscal Studies, working paper n. 26

Duranton, G., and H. G. Overman. (2005), "Testing for Localisation Using Micro-Geographic Data", *Review of Economic Studies*, 72(4), 1077-1106.

Ellison G.,. E. L. Glaeser (1994), "Geographic concentration in U.S. Manufacturing Industries: a Dartboard approach", NBER, working paper n. 4840.

Ellison G.,. E. L. Glaeser (1997), "Geographic Concentration in U.S. Manufacturing Industries: a Dartboard Approach", in *Journal of Political Economy,* 105(3), pp. 889-928.

Ellison, G., and E. L. Glaeser (1999), "Geographic Concentration of Industry: Does Natural Advantage Explain Agglomeration?", in "The American Economic Review Papers and Proceedings", vol. 89, pp. 311-16.

Fleming, M. M. (2004), "Techniques for Estimating Spatially Dependent Discrete Choice Models", in L. Anselin, R. J. G. M. Florax and S. J. Rey (eds.) *Advances in Spatial Econometrics*, Springer, Berlin.

Fujita, M., P. Krugman, and A. J. Venables. (1999), *The Spatial Economy: Cities, Regions, and International Trade*. Cambridge, Mass., The MIT Press.

Guimarães P., O. Figueiredo and D. Woodward (2004), “Measuring the Localization of Economic Activity: A Random Utility Approach”, CEMPRE, working paper n. 161.

Haaland J. I., H. Jarle Kind, K. H. Midelfart Knarvik and J. Torstensson (1999), “What Determines the Economic Geography of Europe?”, CEPR, Discussion Paper n. 2072.

Harville, D. A. (1977), “Maximum Likelihood Approaches to Variance Component Estimation and to Related Problems”, *Journal of the American Statistical Association*, 72, pp. 320-340.

Head, K. and T. Mayer (2002), “Market Potential and the Location of Japanese Investment in the European Union”, CEPR, Discussion Paper n. 3455.

Haining, R. (1978), “The Moving Average Model for Spatial Interaction”, *Trans. Inst. British Geographers*, 3, pp. 202-225.

Holmes T. J. and J. J. Stevens (2002), “Geographic Concentration and Establishment Scale”, *Review of Economics and Statistics*, 84(4), pp. 682-690.

Kelejian, H. H. and D. P. Robinson (1993), “A Suggested Method of Estimation for Spatial Interdependent Models with Autocorrelated Errors”, *Papers in Regional Science*, 72, pp. 297-312.

Kim, S. (1995), “Expansion of Markets and the Geographic Concentration of Economic Activities: the Trends in U.S. Regional Manufacturing Structure, 1860–1987”, *Quarterly Journal of Economics* 110(4), pp. 881–908.

Kim Y., D. L. Barkley, M. S. Henry (2000), “Industry Characteristics Linked to Establishment Concentrations in Nonmetropolitan Areas”, *Journal of Regional Science*, 40(2), pp. 231-59.

Lafourcade M. and G. Mion (2007), “Concentration, Agglomeration and the Size of Plants”, *Regional Science and Urban Economics*, 37(1), pp. 46-68.

Marcon E. and F. Puech (2003), “Evaluating the Geographic Concentration of Industries Using Distance-Based Methods”, *Journal of Economic Geography*, 3(4), pp. 409-428.

Maurel, F. and B. Sedillot (1999), “A measure of the Geographic Concentration in French Manufacturing industries”, *Regional Science and Urban Economics,* 29(5), pp. 575-604.

McFadden, D. L. (1974), “Conditional Logit Analysis of Qualitative Choice Behavior”, in P. Zarembka (ed.) *Frontiers in Econometrics*, London, Academic Press.

McMillen D.P. (1992), “Probit with Spatial Autocorrelation”, *Journal of Regional Science*, 32, pp. 335-348.

Mori T., K. Nishikimi and T. E. Smith, (2005) “A Divergence Statistic for Industrial Localization”, *Review of Economics and Statistics*, 87(4), pp. 635-651.

Pagnini M. (2003), “Misura e Determinanti dell’Agglomerazione Spaziale nei Comparti Industriali in Italia”, *Rivista di Politica Ecomica,* 93(3-4), pp. 149-196.

Rosenthal S. S. and W. C. Strange (2001), “The Determinants of Agglomeration”, *Journal of Urban Economics,* 50, pp. 191-229.

Soubeyran A. and S. Weber (2002), “District Formation and Local Social Capital: A (Tacit) Co-opetition Approach”, October, *Journal of Urban Economics,* 52(1), pp. 65-92.

RECENTLY PUBLISHED "TEMI" (*)

N. 643 – *The producer service sector in Italy: Long-term growth and its local determinants*, by Valter Di Giacinto and Giacinto Micucci (September 2007).

N. 644 – *Aggregazioni bancarie e specializzazione nel credito alle PMI: peculiarità per area geografica*, by Enrico Beretta and Silvia Del Prete (November 2007).

N. 645 – *Costs and benefits of creditor concentration: An empirical approach*, by Amanda Carmignani and Massimo Omiccioli (November 2007).

N. 646 – *Does the underground economy hold back financial deepening? Evidence from the Italian credit market*, by Giorgio Gobbi and Roberta Zizza (November 2007).

N. 647 – *Optimal monetary policy under low trend inflation*, by Guido Ascari and Tiziano Ropele (November 2007).

N. 648 – *Indici di bilancio e rendimenti di borsa: un'analisi per le banche italiane*, by Angela Romagnoli (November 2007).

N. 649 – *Bank profitability and taxation,* by Ugo Albertazzi and Leonardo Gambacorta (November 2007).

N. 650 – *Modelling bank lending in the euro area: A non-linear approach,* by Leonardo Gambacorta and Carlotta Rossi (November 2007).

N. 651 – *Revisiting poverty and welfare dominance,* by Gian Maria Tomat (November 2007).

N. 652 – *The general equilibrium effects of fiscal policy: Estimates for the euro area,* by Lorenzo Forni, Libero Monteforte and Luca Sessa (November 2007).

N. 653 – *Securitisation and the bank lending channel,* by Yener Altunbas, Leonardo Gambacorta and David Marqués (November 2007).

N. 654 – *The cyclical response of fiscal policies in the euro area. Why do results of empirical research differ so strongly?*, by Roberto Golinelli and Sandro Momigliano (January 2008).

N. 655 – *What's behind "inflation perceptions"? A survey-based analysis of Italian consumers*, by Paolo Del Giovane, Silvia Fabiani and Roberto Sabbatini (January 2008).

N. 656 – *The effects of fiscal policy in Italy: Evidence from a VAR model*, by Raffaela Giordano, Sandro Momigliano, Stefano Neri and Roberto Perotti (January 2008).

N. 657 – *Excess money growth and inflation dynamics*, by Barbara Roffia and Andrea Zaghini (January 2008).

N. 658 – *R&D and market structure in a horizontal differentiation framework*, by Davide Fantino (January 2008).

N. 659 – *Housing market spillovers: Evidence from an estimated DSGE model*, by Matteo Iacoviello and Stefano Neri (January 2008).

N. 660 – *Real exchange rate volatility and disconnect: An empirical investigation*, by Riccardo Cristadoro, Andrea Gerali, Stefano Neri and Massimiliano Pisani (April 2008).

N. 661 – *The effect of investment tax credit: Evidence from an atypical programme in Italy*, by Raffaello Bronzini, Guido de Blasio, Guido Pellegrini and Alessandro Scognamiglio (April 2008).

N. 662 – *Accounting for sampling design in the SHIW*, by Ivan Faiella (April 2008).

N. 663 – *Delayed privatization*, by Bernardo Bortolotti and Paolo Pinotti (April 2008).

N. 664 – *Portfolio selection with mononotone mean-variance preferences*, by Fabio Maccheroni, Massimo Marinacci, Aldo Rustichini and Marco Taboga (April 2008).

N. 665 – *Directed matching with endogenous Markov probability: Clients or competitors?*, by Emanuela Ciapanna (April 2008).

N. 666 – *What are borders made of? An analysis of barriers to European banking integration*, by Massimiliano Affinito and Matteo Piazza (April 2008).

N. 667 – *Innovation driven sectoral shocks and aggregate city cycles*, by Andrea R. Lamorgese (April 2008).

N. 668 – *On applying synthetic indices of multidimensional well-being: Health and income inequalities in selected EU countries*, by Andrea Brandolini (April 2008).

N. 669 – *Values, inequality and happiness*, by Claudia Biancotti and Giovanni D'Alessio (April 2008).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Studi di struttura economica e finanziaria – Divisione Biblioteca e Archivio storico – Via Nazionale, 91 – 00184 Rome – (fax 0039 06 47922059). They are available on the Internet www.bancaditalia.it.

2005

L. DEDOLA and F. LIPPI, *The monetary transmission mechanism: Evidence from the industries of 5 OECD countries,* European Economic Review, 2005, Vol. 49, 6, pp. 1543-1569, **TD No. 389 (December 2000)**.

D. Jr. MARCHETTI and F. NUCCI, *Price stickiness and the contractionary effects of technology shocks*. European Economic Review, Vol. 49, 5, pp. 1137-1164, **TD No. 392 (February 2001)**.

G. CORSETTI, M. PERICOLI and M. SBRACIA, *Some contagion, some interdependence: More pitfalls in tests of financial contagion,* Journal of International Money and Finance, Vol. 24, 8, pp. 1177-1199, **TD No. 408 (June 2001)**.

GUISO L., L. PISTAFERRI and F. SCHIVARDI, *Insurance within the firm*. Journal of Political Economy, Vol. 113, 5, pp. 1054-1087, **TD No. 414** (**August 2001**)

R. CRISTADORO, M. FORNI, L. REICHLIN and G. VERONESE, *A core inflation indicator for the euro area,* Journal of Money, Credit, and Banking, Vol. 37, 3, pp. 539-560, **TD No. 435 (December 2001)**.

F. ALTISSIMO, E. GAIOTTI and A. LOCARNO, *Is money informative? Evidence from a large model used for policy analysis,* Economic & Financial Modelling, Vol. 22, 2, pp. 285-304, **TD No. 445 (July 2002)**.

G. DE BLASIO and S. DI ADDARIO, *Do workers benefit from industrial agglomeration?* Journal of regional Science, Vol. 45, (4), pp. 797-827, **TD No. 453 (October 2002).**

G. DE BLASIO and S. DI ADDARIO, *Salari, imprenditorialità e mobilità nei distretti industriali italiani,* in L. F. Signorini, M. Omiccioli (eds.), *Economie locali e competizione globale: il localismo industriale italiano di fronte a nuove sfide*, Bologna, il Mulino, **TD No. 453 (October 2002).**

R. TORRINI, *Cross-country differences in self-employment rates: The role of institutions*, Labour Economics, Vol. 12, 5, pp. 661-683, **TD No. 459 (December 2002)**.

A. CUKIERMAN and F. LIPPI, *Endogenous monetary policy with unobserved potential output,* Journal of Economic Dynamics and Control, Vol. 29, 11, pp. 1951-1983, **TD No. 493 (June 2004)**.

M. OMICCIOLI, *Il credito commerciale: problemi e teorie*, in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 494 (June 2004)**.

L. CANNARI, S. CHIRI and M. OMICCIOLI, *Condizioni di pagamento e differenziazione della clientela,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 495 (June 2004)**.

P. FINALDI RUSSO and L. LEVA, *Il debito commerciale in Italia: quanto contano le motivazioni finanziarie?,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 496 (June 2004)**.

A. CARMIGNANI, *Funzionamento della giustizia civile e struttura finanziaria delle imprese: il ruolo del credito commerciale,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 497 (June 2004)**.

G. DE BLASIO, *Credito commerciale e politica monetaria: una verifica basata sull'investimento in scorte,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 498 (June 2004)**.

G. DE BLASIO, *Does trade credit substitute bank credit? Evidence from firm-level data.* Economic notes, Vol. 34, 1, pp. 85-112, **TD No. 498 (June 2004).**

A. DI CESARE, *Estimating expectations of shocks using option prices,* The ICFAI Journal of Derivatives Markets, Vol. 2, 1, pp. 42-53, **TD No. 506 (July 2004).**

M. BENVENUTI and M. GALLO, *Il ricorso al "factoring" da parte delle imprese italiane,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 518 (October 2004)**.

L. CASOLARO and L. GAMBACORTA, *Redditività bancaria e ciclo economico,* Bancaria, Vol. 61, 3, pp. 19-27, **TD No. 519 (October 2004)**.

F. PANETTA, F. SCHIVARDI and M. SHUM, *Do mergers improve information? Evidence from the loan market*, CEPR Discussion Paper, 4961, **TD No. 521 (October 2004)**.

P. DEL GIOVANE and R. SABBATINI, *La divergenza tra inflazione rilevata e percepita in Italia,* in P. Del Giovane, F. Lippi e R. Sabbatini (eds.), *L'euro e l'inflazione: percezioni, fatti e analisi,* Bologna, Il Mulino, **TD No. 532 (December 2004).**

R. TORRINI, *Quota dei profitti e redditività del capitale in Italia: un tentativo di interpretazione*, Politica economica, Vol. 21, 1, pp. 7-41, **TD No. 551 (June 2005)**.

M. OMICCIOLI, *Il credito commerciale come "collateral"*, in L. Cannari, S. Chiri, M. Omiccioli (eds.), Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia, Bologna, il Mulino, **TD No. 553 (June 2005)**.

L. CASOLARO, L. GAMBACORTA and L. GUISO, *Regulation, formal and informal enforcement and the development of the household loan market. Lessons from Italy*, in Bertola G., Grant C. and Disney R. (eds.) The Economics of Consumer Credit: European Experience and Lessons from the US, Boston, MIT Press, **TD No. 560 (September 2005)**.

S. DI ADDARIO and E. PATACCHINI, *Lavorare in una grande città paga, ma poco*, in Brucchi Luchino (ed.), *Per un'analisi critica del mercato del lavoro*, Bologna , Il Mulino, **TD No. 570 (January 2006)**.

P. ANGELINI and F. LIPPI, *Did inflation really soar after the euro changeover? Indirect evidence from ATM withdrawals,* CEPR Discussion Paper, 4950, **TD No. 581 (March 2006)**.

S. FEDERICO, *Internazionalizzazione produttiva, distretti industriali e investimenti diretti all'estero,* in L. F. Signorini, M. Omiccioli (eds.), *Economie locali e competizione globale: il localismo industriale italiano di fronte a nuove sfide*, Bologna, il Mulino, **TD No. 592 (October 2002).**

S. DI ADDARIO, *Job search in thick markets: Evidence from Italy*, Oxford Discussion Paper 235, Department of Economics Series, **TD No. 605 (December 2006)**.

2006

F. BUSETTI, *Tests of seasonal integration and cointegration in multivariate unobserved component models,* Journal of Applied Econometrics, Vol. 21, 4, pp. 419-438, **TD No. 476 (June 2003).**

C. BIANCOTTI, *A polarization of inequality? The distribution of national Gini coefficients 1970-1996*, Journal of Economic Inequality, Vol. 4, 1, pp. 1-32, **TD No. 487 (March 2004)**.

L. CANNARI and S. CHIRI, *La bilancia dei pagamenti di parte corrente Nord-Sud (1998-2000),* in L. Cannari, F. Panetta (a cura di), Il sistema finanziario e il Mezzogiorno: squilibri strutturali e divari finanziari, Bari, Cacucci, **TD No. 490 (March 2004).**

M. BOFONDI and G. GOBBI, *Information barriers to entry into credit markets,* Review of Finance, Vol. 10, 1, pp. 39-67, **TD No. 509 (July 2004).**

FUCHS W. and LIPPI F., *Monetary union with voluntary participation,* Review of Economic Studies, Vol. 73, pp. 437-457 **TD No. 512 (July 2004).**

GAIOTTI E. and A. SECCHI, *Is there a cost channel of monetary transmission? An investigation into the pricing behaviour of 2000 firms*, Journal of Money, Credit and Banking, Vol. 38, 8, pp. 2013-2038 **TD No. 525 (December 2004)**.

A. BRANDOLINI, P. CIPOLLONE and E. VIVIANO, *Does the ILO definition capture all unemployment?,* Journal of the European Economic Association, Vol. 4, 1, pp. 153-179, **TD No. 529 (December 2004)**.

A. BRANDOLINI, L. CANNARI, G. D'ALESSIO and I. FAIELLA, *Household wealth distribution in Italy in the 1990s,* in E. N. Wolff (ed.) International Perspectives on Household Wealth, Cheltenham, Edward Elgar, **TD No. 530 (December 2004)**.

P. DEL GIOVANE and R. SABBATINI, *Perceived and measured inflation after the launch of the Euro: Explaining the gap in Italy,* Giornale degli economisti e annali di economia, Vol. 65, 2 , pp. 155-192, **TD No. 532 (December 2004).**

M. CARUSO, *Monetary policy impulses, local output and the transmission mechanism*, Giornale degli economisti e annali di economia, Vol. 65, 1, pp. 1-30, **TD No. 537 (December 2004).**

A. NOBILI, *Assessing the predictive power of financial spreads in the euro area: does parameters instability matter?,* Empirical Economics, Vol. 31, 1, pp. 177-195, **TD No. 544 (February 2005)**.

L. GUISO and M. PAIELLA, *The role of risk aversion in predicting individual behavior, In P. A. Chiappori e C. Gollier (eds.) Competitive Failures in Insurance Markets: Theory and Policy Implications, Monaco, CESifo*, **TD No. 546 (February 2005).**

G. M. TOMAT, *Prices product differentiation and quality measurement: A comparison between hedonic and matched model methods*, Research in Economics, Vol. 60, 1, pp. 54-68, **TD No. 547 (February 2005)**.

F. LOTTI, E. SANTARELLI and M. VIVARELLI, *Gibrat's law in a medium-technology industry: Empirical evidence for Italy*, in E. Santarelli (ed.), Entrepreneurship, Growth, and Innovation: the Dynamics of Firms and Industries, New York, Springer, **TD No. 555 (June 2005).**

F. BUSETTI, S. FABIANI and A. HARVEY, *Convergence of prices and rates of inflation*, Oxford Bulletin of Economics and Statistics, Vol. 68, 1, pp. 863-878, **TD No. 575 (February 2006).**

M. CARUSO, *Stock market fluctuations and money demand in Italy, 1913 - 2003*, Economic Notes, Vol. 35, 1, pp. 1-47, **TD No. 576 (February 2006)**.

S. IRANZO, F. SCHIVARDI and E. TOSETTI, *Skill dispersion and productivity: An analysis with matched data*, CEPR Discussion Paper, 5539, **TD No. 577 (February 2006).**

R. BRONZINI and G. DE BLASIO, *Evaluating the impact of investment incentives: The case of Italy's Law 488/92.* Journal of Urban Economics, Vol. 60, 2, pp. 327-349, **TD No. 582 (March 2006).**

R. BRONZINI and G. DE BLASIO, *Una valutazione degli incentivi pubblici agli investimenti*, Rivista Italiana degli Economisti , Vol. 11, 3, pp. 331-362, **TD No. 582 (March 2006).**

A. DI CESARE, *Do market-based indicators anticipate rating agencies? Evidence for international banks*, Economic Notes, Vol. 35, pp. 121-150, **TD No. 593 (May 2006).**

L. DEDOLA and S. NERI, *What does a technology shock do? A VAR analysis with model-based sign restrictions*, Journal of Monetary Economics, Vol. 54, 2, pp. 512-549, **TD No. 607 (December 2006)**.

R. GOLINELLI and S. MOMIGLIANO, *Real-time determinants of fiscal policies in the euro area*, Journal of Policy Modeling, Vol. 28, 9, pp. 943-964, **TD No. 609 (December 2006).**

P. ANGELINI, S. GERLACH, G. GRANDE, A. LEVY, F. PANETTA, R. PERLI,S. RAMASWAMY, M. SCATIGNA and P. YESIN, *The recent behaviour of financial market volatility*, BIS Papers, 29, **QEF No. 2 (August 2006)**.

2007

L. CASOLARO. and G. GOBBI, *Information technology and productivity changes in the banking industry*, Economic Notes, Vol. 36, 1, pp. 43-76, **TD No. 489 (March 2004)**.

M. PAIELLA, *Does wealth affect consumption? Evidence for Italy*, Journal of Macroeconomics, Vol. 29, 1, pp. 189-205, **TD No. 510 (July 2004).**

F. LIPPI. and S. NERI, *Information variables for monetary policy in a small structural model of the euro area*, Journal of Monetary Economics, Vol. 54, 4, pp. 1256-1270, **TD No. 511 (July 2004)**.

A. ANZUINI and A. LEVY, *Monetary policy shocks in the new EU members: A VAR approach*, Applied Economics, Vol. 39, 9, pp. 1147-1161, **TD No. 514 (July 2004)**.

R. BRONZINI, *FDI Inflows, agglomeration and host country firms' size: Evidence from Italy*, Regional Studies, Vol. 41, 7, pp. 963-978, **TD No. 526 (December 2004).**

L. MONTEFORTE, *Aggregation bias in macro models: Does it matter for the euro area?*, Economic Modelling, 24, pp. 236-261, **TD No. 534 (December 2004)**.

A. DALMAZZO and G. DE BLASIO, *Production and consumption externalities of human capital: An empirical study for Italy*, Journal of Population Economics, Vol. 20, 2, pp. 359-382, **TD No. 554 (June 2005).**

M. BUGAMELLI and R. TEDESCHI, *Le strategie di prezzo delle imprese esportatrici italiane*, Politica Economica, v. 3, pp. 321-350, **TD No. 563 (November 2005)**.

L. GAMBACORTA and S. IANNOTTI, *Are there asymmetries in the response of bank interest rates to monetary shocks?*, Applied Economics, v. 39, 19, pp. 2503-2517, **TD No. 566 (November 2005).**

S. DI ADDARIO and E. PATACCHINI, *Wages and the city. Evidence from Italy*, Development Studies Working Papers 231, Centro Studi Luca d'Agliano, **TD No. 570 (January 2006)**.

P. ANGELINI and F. LIPPI, *Did prices really soar after the euro cash changeover? Evidence from ATM withdrawals*, International Journal of Central Banking, Vol. 3, 4, pp. 1-22, **TD No. 581 (March 2006)**.

A. LOCARNO, *Imperfect knowledge, adaptive learning and the bias against activist monetary policies*, International Journal of Central Banking, v. 3, 3, pp. 47-85, **TD No. 590 (May 2006)**.

F. LOTTI and J. MARCUCCI, *Revisiting the empirical evidence on firms' money demand*, Journal of Economics and Business, Vol. 59, 1, pp. 51-73, **TD No. 595 (May 2006)**.

P. CIPOLLONE and A. ROSOLIA, *Social interactions in high school: Lessons from an earthquake*, American Economic Review, Vol. 97, 3, pp. 948-965, **TD No. 596 (September 2006).**

A. BRANDOLINI, *Measurement of income distribution in supranational entities: The case of the European Union*, in S. P. Jenkins e J. Micklewright (eds.), Inequality and Poverty Re-examined, Oxford, Oxford University Press, **TD No. 623 (April 2007).**

M. PAIELLA, *The foregone gains of incomplete portfolios,* Review of Financial Studies, Vol. 20, 5, pp. 1623-1646, **TD No. 625 (April 2007).**

K. BEHRENS, A. R. LAMORGESE, G.I.P. OTTAVIANO and T. TABUCHI, *Changes in transport and non transport costs: local vs. global impacts in a spatial network*, Regional Science and Urban Economics, Vol. 37, 6, pp. 625-648, **TD No. 628 (April 2007).**

G. ASCARI and T. ROPELE, *Optimal monetary policy under low trend inflation*, Journal of Monetary Economics, v. 54, 8, pp. 2568-2583, **TD No. 647 (November 2007).**

R. GIORDANO, S. MOMIGLIANO, S. NERI and R. PEROTTI, *The Effects of Fiscal Policy in Italy: Evidence from a VAR Model*, European Journal of Political Economy, Vol. 23, 3, pp. 707-733, **TD No. 656 (December 2007).**

2008

S. MOMIGLIANO, J. Henry and P. Hernández de Cos, *The impact of government budget on prices: Evidence from macroeconometric models*, Journal of Policy Modelling, v. 30, 1, pp. 123-143 **TD No. 523 (October 2004).**

P. ANGELINI and A. Generale, *On the evolution of firm size distributions*, American Economic Review, v. 98, 1, pp. 426-438, **TD No. 549 (June 2005).**

P. DEL GIOVANE, S. FABIANI and R. SABATINI, *What's behind "inflation perceptions"? A survey-based analysis of Italian consumers, i*n P. Del Giovane e R. Sabbatini (eds.), *The Euro Inflation and Consumers' Perceptions. Lessons from Italy*, Berlin-Heidelberg, Springer, **TD No. 655 (January 2008).**

FORTHCOMING

S. SIVIERO and D. TERLIZZESE, *Macroeconomic forecasting: Debunking a few old wives' tales,* Journal of Business Cycle Measurement and Analysis, **TD No. 395 (February 2001)**.

P. ANGELINI, *Liquidity and announcement effects in the euro area*, Giornale degli economisti e annali di economia, **TD No. 451 (October 2002).**

S. MAGRI, *Italian households' debt: The participation to the debt market and the size of the loan*, Empirical Economics, **TD No. 454 (October 2002)**.

P. ANGELINI, P. DEL GIOVANE, S. SIVIERO and D. TERLIZZESE, *Monetary policy in a monetary union: What role for regional information?*, International Journal of Central Banking, **TD No. 457 (December 2002)**.

L. MONTEFORTE and S. SIVIERO, *The Economic Consequences of Euro Area Modelling Shortcuts,* Applied Economics, **TD No. 458 (December 2002).**

L. GUISO and M. PAIELLA,, *Risk aversion, wealth and background risk*, Journal of the European Economic Association, **TD No. 483 (September 2003).**

G. FERRERO, *Monetary policy, learning and the speed of convergence*, Journal of Economic Dynamics and Control, **TD No. 499 (June 2004).**

F. SCHIVARDI e R. TORRINI, *Identifying the effects of firing restrictions through size-contingent Differences in regulation*, Labour Economics, **TD No. 504 (giugno 2004)**.

C. BIANCOTTI, G. D'ALESSIO and A. NERI, *Measurement errors in the Bank of Italy's survey of household income and wealth*, Review of Income and Wealth, **TD No. 520 (October 2004)**.

D. Jr. MARCHETTI and F. Nucci, *Pricing behavior and the response of hours to productivity shocks,* Journal of Money Credit and Banking, **TD No. 524 (December 2004).**

L. GAMBACORTA, *How do banks set interest rates?,* European Economic Review, **TD No. 542 (February 2005).**

R. FELICI and M. PAGNINI,, *Distance, bank heterogeneity and entry in local banking markets*, The Journal of Industrial Economics, **TD No. 557 (June 2005).**

M. BUGAMELLI and R. TEDESCHI, *Le strategie di prezzo delle imprese esportatrici italiane*, Politica Economica, **TD No. 563 (November 2005).**

S. DI ADDARIO and E. PATACCHINI, *Wages and the city. Evidence from Italy*, Labour Economics, **TD No. 570 (January 2006)**.

M. BUGAMELLI and A. ROSOLIA, *Produttività e concorrenza estera*, Rivista di politica economica, **TD No. 578 (February 2006).**

PERICOLI M. and M. TABOGA, *Canonical term-structure models with observable factors and the dynamics of bond risk premia*, **TD No. 580 (February 2006).**

E. VIVIANO, *Entry regulations and labour market outcomes. Evidence from the Italian retail trade sector*, Labour Economics, **TD No. 594 (May 2006)**.

S. FEDERICO and G. A. MINERVA, *Outward FDI and local employment growth in Italy*, Review of World Economics, Journal of Money, Credit and Banking, **TD No. 613 (February 2007).**

F. BUSETTI and A. HARVEY, *Testing for trend*, Econometric Theory **TD No. 614 (February 2007).**

V. CESTARI, P. DEL GIOVANE and C. ROSSI-ARNAUD, *Memory for Prices and the Euro Cash Changeover: An Analysis for Cinema Prices in Italy*, In P. Del Giovane e R. Sabbatini (eds.), The Euro Inflation and Consumers' Perceptions. Lessons from Italy, Berlin-Heidelberg, Springer, **TD No. 619 (February 2007).**

B. ROFFIA and A. ZAGHINI, *Excess money growth and inflation dynamics*, International Finance, **TD No. 629 (June 2007)**.

M. DEL GATTO, GIANMARCO I. P. OTTAVIANO and M. PAGNINI, *Openness to trade and industry cost dispersion: Evidence from a panel of Italian firms*, Journal of Regional Science, **TD No. 635 (June 2007)**.

A. CIARLONE, P. PISELLI and G. TREBESCHI, *Emerging Markets' Spreads and Global Financial Conditions*, Journal of International Financial Markets, Institutions & Money, **TD No. 637 (June 2007).**

S. MAGRI, *The financing of small innovative firms: The Italian case*, Economics of Innovation and New Technology, **TD No. 640 (September 2007)**.