THE ROLE OF INSTITUTIONAL FACTORS IN FOSTERING THE DEVELOPMENT OF INDUSTRIAL DISTRICTS IN ITALY

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1. Introduction

This paper aims to analyze the role of institutional factors in fostering the agglomeration of SMEs in Italian industrial districts. We follow Weber (1922) in distinguishing between institutions created for economic reasons and 'economically relevant' institutions. By 'institutions' here we mean both regulatory sources of economic activities and moral attitudes of economic actors.

The focus of our analysis is the role of 'economically relevant' institutional factors in promoting the growth of industrial districts, since the latter's characteristics of productive specialization and spatial agglomeration do not find adequate explanations in economic theory (both exogenous and endogenous growth theories; see, for example, Pellegrini, 2002).

In addition, while New geography models (Krugman, 1991) explain agglomeration with pecuniary external economies, in Italy a vast literature has highlighted the role of institutional factors in explaining the development of industrial districts (e.g. Becattini and Rullani, 1993). This literature, despite some efforts (Signorini, 2000), is mainly sociological and qualitative, focusing principally on case studies of some industrial districts.

The aim of this work is to improve the empirical analysis of the phenomenon by setting up a regression study of Italian provincial data. The paper is structured as follows. Section 2 overviews the relevant socio-economic literature in order to single out the institutional factors that have been proposed as potential candidates for explaining the development of industrial districts in Italy. Section 3 describes the features of the empirical

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study, with reference to both measurement issues and model specification. Details of the econometric methodology and regression results are given in Section 4, while Section 5 concludes with a brief summary of the main findings.

2. Review of the literature

In this paper, the institution taxonomy proposed by Parri (2002) is adopted. The author identifies three main institutional dimensions that foster economic development: economic institutions, collective action institutions established for extra-economic goals, and political institutions. The author also stresses the importance of the local social framework, called ‘economically relevant’ institutions of the community, which are important for determining the framework in which economic activity takes place.

According to the socio-economic literature on industrial districts, the family has a significant role among the ‘economically relevant’ institutions of the community. It is considered a possible source of internal savings and of free labour that could have fostered the proliferation of family-run businesses typical of industrial districts (Bordogna, 2002; Belfanti and Onger, 2002). In addition, a larger family size encourages the construction of interpersonal trust relationships and networks that facilitate the reduction of transaction costs and the circulation of information (Arrighetti and Seravalli, 1999). Moreover, some historical studies show that sharecropping (*métayers*) families experienced the transformation of their agriculture activities, based on family-run work, into small industrial firms (Belfanti and Onger, 2002).

Other forms of economically relevant institutions of the community are socio-cultural variables, such as trust and moral attitudes. Better endowments in this field can help to reduce transaction costs and increase circulation of information and cooperation, by trimming down the probability of encountering a free-riding behaviour.[1] The socio-economic literature claims that the network of relationships between economic actors

[1] According to Kreps (1990) the proximity of actors, characteristic of industrial districts area, helps a cooperative solution to the collective actions dilemma in game theory economic models.

in the community works as a key comparative advantage for industrial district areas.

A vast literature views collective action institutions as the explanation of how small firms in districts overcome the problems related to their small size. This operates also through associations formed by economic actors. Such associations increase individuals' contractual power vis-à-vis public institutions, which can in turn result in the provision of more and better collective goods. Forms of collective institutions are credit cooperatives and commercial promotion consortiums (Parri, 2002).

In the public sector it is possible to distinguish between central and local government institutions. While central institutions mainly supply universal goods, local ones provide selective public goods, i.e. public goods available to specific areas and individuals (Arrighetti and Seravalli,1999).[2] Local public institutions have a comparative advantage in providing public goods, since they have both access to economies of scale and the ability to adapt to local and specific concerns (Lanzalaco, 1999). In addition, some historical analyses (Belfanti, 1999) have used case studies on industrial districts to show[3] the importance of coordination activities provided by local public institutions.

Pyke *et al.* (1990) and Leonardi (1995) also stress the importance, for the development of industrial districts, of the external economies of scale granted by local governments as providers of specific services (e.g. public infrastructure, public social services, subsidized access to credit, access to public development funds, etc.), not afforded by private bodies.

Based on these considerations only local government activity is considered in the empirical analysis.

Obviously, the determinants of the development of industrial districts are not confined to institutional factors, on which this paper focuses, and in the empirical study due allowance is made for the existence of such alternative factors.

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2 Arrighetti and Seravalli (1999) use the term 'intermediate institutions' to refer not only to local public institutions but also to collective action institutions.

3 There is actually no consensus on this idea. Parri (2002) underlines a marginal role for these institutions in fostering local development, while showing the inefficiency of some recent experiences of local public institution activities such as the creation of centres for providing services to firms in industrial districts.

3. The empirical assessment strategy

The empirical assessment strategy adopted here aims to quantify the differential effects of institutional factor endowments on the local diffusion of industrial districts (LDIID) throughout Italy. To this purpose LDIID is regressed on a set of selected proxies of the above institutional factors and some control variables.

The units of analysis are the administrative provinces, the smallest geographical units for which data on institutional factors are available.

3.1 *Dependent variable specifications*

There is no consensus in the literature on how to define industrial districts. Some sociologists deny even the possibility of measurement. While the empirical approach adopted here needs an operationalization of our dependent variable, for the sake of robustness the analysis will be carried out by means of four alternative measures of LDIID.

Two quantitative approaches have been proposed in the literature. The first is based on the mapping proposed by the Italian National Institute of Statistics (Istat) of the local labour market areas (LLMAs). LLMAs are self-contained geographical units which are capable of offering employment to the majority of their resident population.

The Sforzi-Istat algorithm (Istat, 1997) selects industrial districts from LLMAs by choosing those LLMAs that fulfil certain criteria, based on the share of manufacturing employment, industrial specialization and prevalence of small and medium-sized firms. Also taking LLMAs as the spatial unit of analysis, Cannari and Signorini (2000) propose two models for defining industrial districts, one providing a continuous measure of "district intensity" for each LLMA, and the other singling out industrial districts from LLMAs as in Sforzi-Istat (Istat, 1997), but applying more restrictive classification criteria.

An alternative quantitative approach to the identification of industrial districts has been recently proposed by Iuzzolino (in this volume). The author does not use LLMAs as the spatial unit of analysis, but looks for sectoral-spatial agglomeration of firms in nearby municipalities.

To summarize, we consider the following four alternative measures of LDIID:

- DIST1 is an indicator of the provincial diffusion of industrial districts, calculated as the share of manufacturing employment in municipalities located within industrial districts by Sforzi's algorithm and total manufacturing employment in the province.[4]
- DIST2 is an indicator calculated as DIST1, but making use of the more restrictive district identification algorithm proposed by Cannari and Signorini (2000).
- DIST3 is calculated as follows. First, we multiply, for each municipality in the province, the number of manufacturing employees by the continuous degree of industrial district diffusion of the LLMA to which the municipality belongs, as provided by Cannari and Signorini (2000). Then we sum the figures of all the municipalities in the province and divide this aggregate by the total number of manufacturing employees.
- DIST4 is calculated as the percentage share of employees that work in municipalities which, according to Iuzzolino, belong to a district area.

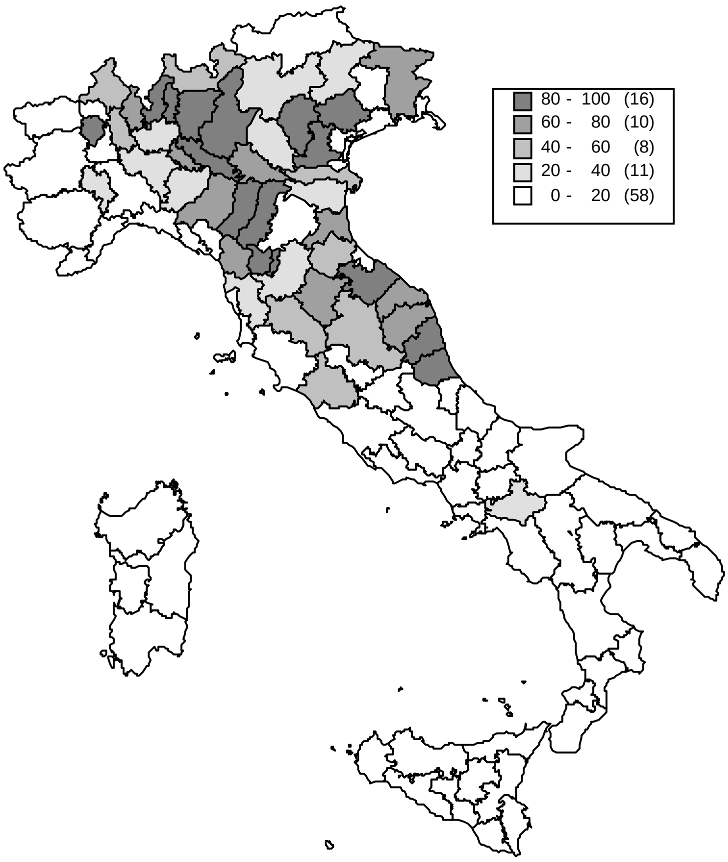
For all four indicators employment figures are taken from the 1996 Census of Industry and Services database.

Maps portraying the geographical distribution of industrial districts for our four measures of LDIID are displayed in Figures 1 to 4. A simple visual inspection shows that quite similar pictures obtain when using the DIST1 and DIST3 measures, while adopting more restrictive criteria (DIST2) to define an LLMA as an industrial district, the phenomenon tends to concentrate into a smaller area. Last, DIST4, the measure of LDIID not based on LLMA, shows fewer geographical disparities and a more significant presence of industrial districts in peripheral regions.

[4] We consider all manufacturing employment in a district and not only the employment share of the sector in which the district is specialized.

Figure 1

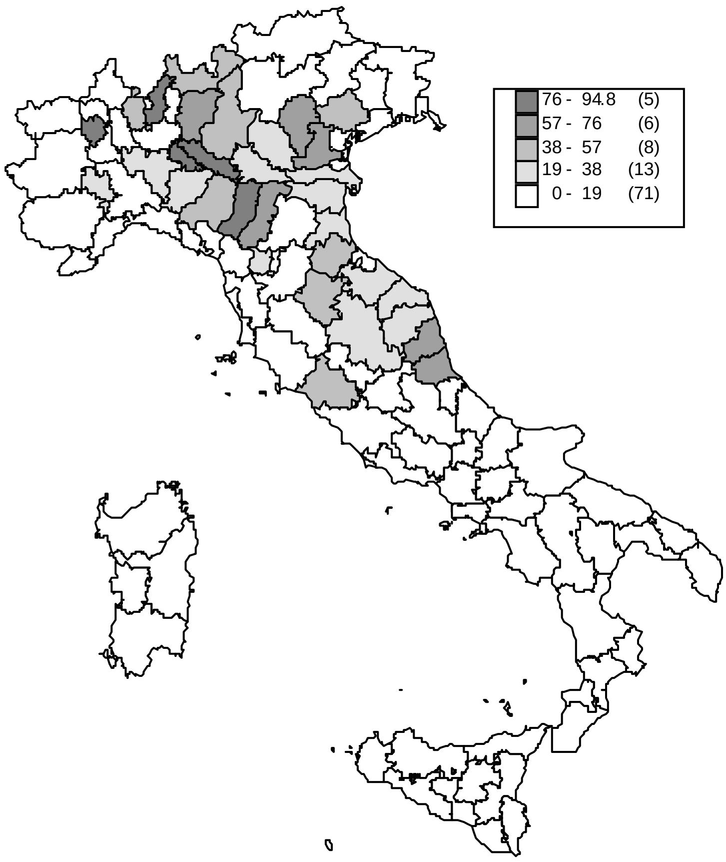
Local diffusion of industrial districts in Italian provinces
Measure DIST1



Source: Based on Isat data.

Figure 2

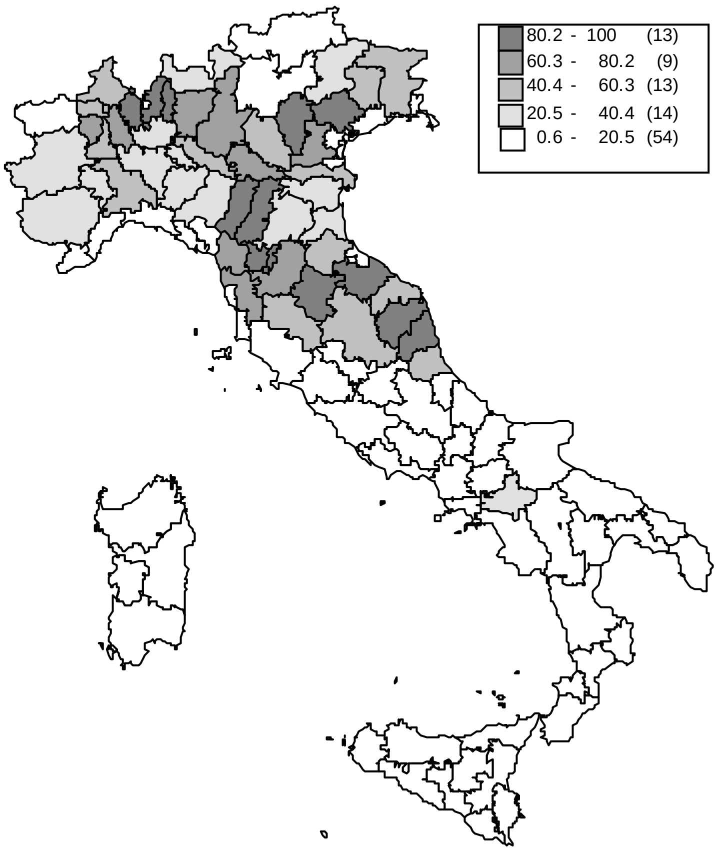
Local diffusion of industrial districts in Italian provinces
Measure DIST2



Source: Based on Istat data.

Figure 3

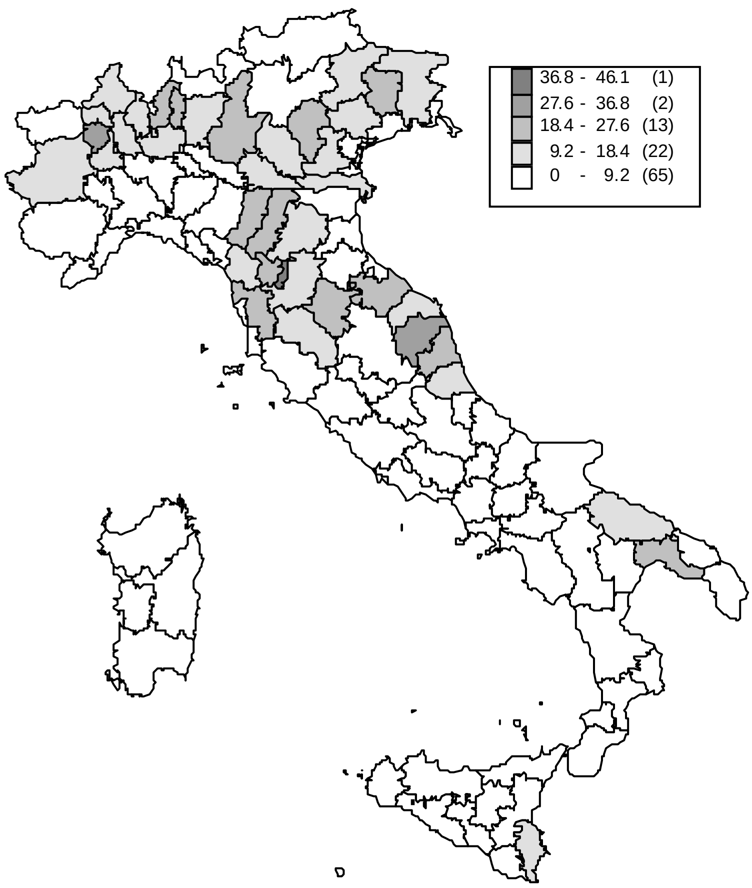
Local diffusion of industrial districts in Italian provinces
Measure DIST3



Source: Based on Istat data.

Figure 4

Local diffusion of industrial districts in Italian provinces
Measure DIST4



Source: Based on Istat data.

3.2 *Specification of the model*

The specification of the model takes into account the actual availability of data that can be used as proxies for the institutional factors underlined by the socio-economic literature as determinants of the development of industrial districts. In addition, we considered the simultaneity and reverse causation problems that these indicators might entail.

The simultaneity problem is mainly tackled by using pre-determined regressors. We decided to refer mainly to the 1950s or 1960s, since the literature on the origin of industrial districts demonstrates that in the 1970s this development model experienced a decisive boost.

According to Brusco and Paba (1997), based on census data there were 149 districts in 1951, employing 360,000 workers, quite homogeneously distributed over the country (with the exception of Sicily and Sardinia). In 1971, the number of districts had only slightly increased to 166, with a million employees. The real take-off took place in the two following decades, with the number of districts reaching 238, with 1,800,000 workers, in 1991.

As explained above, the taxonomy of institutions proposed by Parri (2002) is adopted, and the closest proxies were chosen for which predetermined data at the provincial level were available. A brief description of the indicators used follows in this section, while more details can be found in the Appendix.

Economically relevant institutions are represented by two variables: one which refers to family organization and the other to free-riding behaviour.

For the first variable the indicator used is the share of ‘extended’ households in the total number of households in 1951.[5] Our aim was to discover whether strong family institutions fostered the birth and development of SMEs within given areas, mainly by providing human resources, pooled savings and relational capital (the indicator is, moreover, strongly positively correlated to a proxy of the diffusion of sharecropping

[5] According to the Italian National Institute of Statistics, an household is ‘extended’ when it consists of the head of the household, his wife, their children, their parents or other relatives.

in 1950,[6] a variable that, in turn, received some attention in the literature on the development of industrial districts).

Individual propensity to free-riding is captured by an indicator (ISTOPPO) resulting from a principal component analysis of a set of variables related to the number of crimes affecting economic property and the rate of insolvency in contracts in the late 1950s.

Collective action institutions are proxied by a latent factor provided by a principal component analysis of various variables relating to economic associationism in the late 1960s (ISTAZCO). We excluded variables relating to the economic association of firms in the manufacturing sector, since this could have been endogenous with respect to our dependent variable, while we focused on variables relating to handicrafts and commerce. The use of these variables is designed to capture the propensity of economic actors to cooperate.

Table 1

Summary statistics

Variables	North		Centre		South	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
DIST1	42.7	35.8	40.2	37.7	6.4	18.5
DIST2	28.9	27.7	15.9	20.4	4.3	12.2
DIST3	43.1	27.8	42.2	38.1	4.7	12.4
DIST4	10.0	8.0	11.8	9.6	4.1	5.0
ISTFAMI	25.4	0.4	29.9	5.8	17.0	5.5
ISTOPPO	-0.6	0.56	0.7	1.8	0.2	0.8
ISTAZCO	0.6	0.9	0.1	0.9	-0.8	0.5
ISTELOC	0.6	0.1	0.5	0.1	0.4	0.1
VAFIT	20,960	1,455	17,901	1,484	12,160	1,175

6 The indicator is the percentage share of agricultural productive surface cultivated by farmers employed with quasi-autonomous contracts (*colonia parziale* or *compartecipazione*) in total agricultural productive surface. Source: ISL-University of Parma database.

The role of local government institutions is analyzed by means of an indicator of the level of effort made by local institutions to encourage economic development by supplying selective public goods such as infrastructure and education. It is calculated as the ratio of expenditure on education and public works to total local government expenditure in the early 1960s.

Table 1 provides some descriptive statistics for the set of indicators considered in the model.

3.3 *The identification problem*

Let D and F respectively denote the local diffusion of industrial districts and the chosen proxies of the above institutional factors. An empirical assessment of the relevance of the latter in explaining spatial variation in the former could, in principle, be based on the estimation of the following simple regression model:

$$D = F\lambda + \varepsilon. \tag{1}$$

This simple specification, however, is likely to suffer from a missing variable problem due to the omission from the set of regressors of an indicator of the degree of economic development in the area where the industrial district is located. An important result in new economic geography theory states, in fact, that local economic growth may trigger further agglomeration of industrial activity through the location of new firms attracted by increased local demand (Fujita *et al.*, 1999; Fujita and Thisse, 1996), and a visual inspection of the maps (Figures 1-4) portraying the geographical distribution of industrial districts over Italy shows how the former tend to concentrate in more developed areas.[7]

Given the plausible correlation between the explanatory variables in (1) and the omitted variable,[8] the OLS estimation of model parameters would result in biased measures of the influence of institutional factors on the formation and growth of industrial districts.

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[7] The correlation between per capita value added in 2000, as estimated by Istituto G. Tagliacarne, and the selected indicators of the diffusion of industrial districts is equal to 0.47, 0.39, 0.59 and 0.40 for DIST1, DIST2, DIST3 and DIST4, respectively.

[8] Hall and Jones (1999) and Guiso *et al.* (2004) are two recent studies providing evidence of the existence of a significant influence of institutional variables on local economic and financial development.

Let Y denote a proxy of local economic development, such as per capita GDP. To remove this bias one could base the empirical analysis on the following augmented specification:

$$D = F\lambda + \alpha Y + U. \tag{2}$$

In proceeding to estimate model (2) one has to take into account the issue of simultaneity, deriving from possible reverse causation links from D to Y [9] (agglomeration economies originating within districts are expected to foster, *ceteris paribus*, productivity and income growth in the areas where districts are located).

If, as found by a number of empirical studies, institutional factors exert a direct influence on both the formation of industrial districts and local economic development in general, the simultaneity issue is associated with an identification problem, illustrated by following system of structural equations:

$$\begin{cases} D = F\lambda + \alpha Y + U_1 & \text{(3.a)} \\ Y = F\psi + \delta D + U_2 & \text{(3.b)} \end{cases}$$

where U_1 and U_2 are two, possibly correlated, stochastic error terms with the usual properties and F are exogenous variables ($F \perp U_1, U_2$). As is well known, in the absence of sufficient restrictions on the λ and ψ coefficients or on the disturbances covariance matrix, neither equation in (3) is identified.

To identify the parameters of equation (3.a), we assume that a second set of exogenous variables exists, denoted as X, directly impacting on Y but having no direct effect on D, once the local level of Y is controlled for. In formal terms this amounts to assuming that the following set of simultaneous equations holds:

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9 The existence of causal feedback from D to Y was empirically verified by means of the Wu-Hausman exogeneity test, which did not allow the null hypothesis of simultaneity between the two variables (using DIST3 as a proxy of D) to be rejected at conventional significance levels.

$$\begin{cases} D = F\lambda + \alpha Y + U_1 & (4.a) \\ Y = F\psi + X\tau + \delta D + U_2. & (4.b) \end{cases}$$

As a preliminary specification, we included among the *X* variables measures of labour productivity (computed as the ratio of value added to labour input in year 2000), in the agriculture, construction and services sectors. At the root of this choice lies the *a priori* assumption that relatively high productivity levels in sectors other than manufacturing, while fostering income growth, do not exercise sizeable direct effects on the process of spatial economic agglomeration once the overall local level of development is properly controlled for.

The number of instrumental variables available is larger than would be strictly necessary for exact identification. This makes it possible to use the Sargan test to check for over-identifying restrictions. Since the test is not statistically significant (*p-value*= 0.62) it appears to support the validity of the chosen instruments.

Based on the identification equation (4.a), coefficients can be consistently estimated using standard instrumental variable techniques. However, the issue of parameter estimation becomes more complex once possible spatial interactions among the variables are allowed for; this point is treated in the following section.

4. The econometric analysis

In selecting the appropriate econometric estimation techniques for the different measures of the degree of district diffusion, one must take into account the fact that DIST1, DIST2 and DIST4 are left censored at zero, since for a substantial fraction of the provinces in the sample (respectively equal to 27, 21 and 16 units out of 85) the algorithms do not identify the presence of any industrial district. To cope with the censored nature of the dependent variable a Tobit specification was chosen in this case. DIST3 is not censored and, hence, a simple linear specification was used. Explanatory factors were the same in all the four specifications considered.

We allowed for possible spatial interaction effects among the observed locations. Indeed, the spatial distribution of industrial districts across the country does not appear to be random, with a clear tendency of the districts to concentrate in specific regions.

The observed association of high values in a given province with equally high values in nearby provinces implies a positive spatial autocorrelation in the data, originating from one or both of the following mechanisms:

- factors promoting the development of industrial districts are themselves spatially autocorrelated and transmit the feature to the dependent variable;
- geographically close areas interact with each other in a stronger way, for example due the fact that spatial proximity facilitates knowledge transmission and fosters imitative behaviour.

The second type of mechanism is known in the literature as the spatial spillover or contagion effect.

A spatial autocorrelation measure that is widely used in empirical studies is Moran's I index (see e.g. Cliff and Ord, 1981). Letting y_i denote the value of y observed on location i, the expression for Moran's statistic is the following:

$$I = \frac{N}{S} \sum_{i=1}^{N} \sum_{j=1}^{N} w_{ij} (y_i - \bar{y})(y_j - \bar{y}) \tag{5}$$

where N is the number of observations, $\bar{y}$ is the sample mean of y, $S = \sum_{ij} w_{ij}$ and w_{ij} the i-th row and j-th column element of the spatial contiguity, or weights, matrix W[10], with the usual properties

$$\begin{cases} w_{ij} > 0 & \text{if locations } i \text{ and } j \text{ are contiguous in space} \\ w_{ij} = 0 & \text{otherwise.} \end{cases}$$

The spatial weights matrix is said to be row normalized if the elements satisfy the restriction $\sum_j w_{ij} = 1$.

[10] More details of spatial contiguity matrices are given in Anselin (1988), Chapter 3.

Table 2 displays the estimated spatial autocorrelation indices for DIST3 and the set of explanatory variables. The spatial weights matrix used in the computations was obtained by row-normalizing a binary contiguity matrix, i.e. a matrix whose elements are equal to 1 if the two provinces share a common border, and 0 otherwise.

Table 2

Spatial autocorrelation estimates for LDIID measures and explanatory variables

Variables	*Moran's I*	*t test*	*p-value*
DIST3	0.59	8.01	0.000
ISTFAMI	0.83	11.16	0.000
ISTOPPO	0.28	3.80	0.000
ISTAZCO	0.64	8.09	0.000
ISTELOC	0.55	7.39	0.000
VAFIT	0.89	11.89	0.000

As is already apparent from the visual inspection of Figures 1-4, the geographical distribution of LDIIDs in Italy is denoted by a marked positive autocorrelation (neighbouring provinces tend to exhibit similar values). The value of the I index is equal to 0.59 (compared to a theoretical maximum that, given row normalization of *W*, is equal to 1) and is highly statistically significant.

A similar result is observed for the set of explanatory factors as well, with a particularly high value recorded for the family institutions indicator (ISTFAMI).

Substantial spatial autocorrelation in explanatory factors could therefore provide an explanation for the observed tendency of industrial districts to localize in specific areas of the country.

However, agglomerating effects induced by the spatial pattern of factors promoting LDIID can be strengthened by a spatial contagion mechanism analogous to the one set forth by Brusco and Paba.

To obtain inferences that are robust to the presence of spatial autocorrelation in the data, and with the objective (secondary to the purpose of the study but not lacking interest *per se*) of identifying and measuring possible spatial spillover effects amongst geographically close areas, the empirical analysis makes use of spatial econometrics techniques.

To this purpose, model (4.a) was extended by adopting the following spatially autoregressive specification:

$$D^* = \rho WD^* + F\lambda + \alpha Y + U_1^* \tag{6}$$

where D^* is the spatial series of the N observations on the dependent variable that may or may not be left censored; W, as above, is the spatial weights matrix; ρ is the spatial autoregressive coefficient measuring the degree of interdependence amongst contiguous areas; U_1^* is an error term, spatially uncorrelated and orthogonal to F but possibly heteroskedastic and, due to simultaneity, correlated with Y.

In the following we will adopt the same W matrix as the one utilised to compute the Moran statistics. With this specification for the weights, the spatial lag of the dependent variable (i.e., the product WD^*) turns out to be simply the mean of the observations on contiguous locations.

If significant spatial spillover effects among spatial units exist, the omission of the lagged dependent variable from the right-hand side of (6), given the likely non-zero correlation between the omitted variable and remaining regressors, would result in biased and inconsistent estimates of the λ and α coefficients. Even if the spatial lag of the endogenous variable is uncorrelated with regressors, its omission would cause the model's residuals to be correlated as a consequence of the fact that D^* is spatially autocorrelated. This, in turn, would imply that estimated standard errors of regression coefficients are biased, thus possibly leading to wrong inferences.

On the basis of proper assumptions about the relationship linking the observed variable D_i to the latent variable D_i^*, the two following specifications can be derived from expression (4):

1) linear : $D_i = D_i^*$

2) Tobit: $D_i = \begin{cases} D_i^* & \text{if } D_i^* \geq 0 \\ 0 & \text{otherwise.} \end{cases}$

While linear regression models including spatial lags of the dependent variable have received considerable attention in the statistical and econometric literature (Cliff and Ord, 1981 and Anselin, 1988 are two classic references in such contexts), Tobit models augmented to consider spatial interaction effects have only recently been introduced by LeSage (2000).

Both in the linear and the Tobit case the methodological approach follows the Bayesian specification proposed by LeSage (1997; 2000), which allows for the possible presence of outliers or heteroskedastic distrubances. Based on LeSage (2000), we complete the Bayesian specification of the model by augmenting (6) with the following assumptions regarding the model parameters and the error term:

$$U_1^* \sim N(0, \sigma^2 V), \qquad \mathrm{V} = diag\{v_1, v_2, ..., v_N\}$$

$$\pi(\rho) \propto \text{constant}$$

$$\pi(v_i^{-1}) \sim \text{ID } \chi^2(q)/q \qquad (i = 1,2,...,N)$$

$$\pi(\sigma^2) \propto 1/\sigma$$

$$q \sim \text{constant}.$$

The positive hyper-parameter q controls the amount of cross-sectional dispersion in error variances. As q diverges the model tends to become homoskedastic. According to some preliminary evidence of residual heteroskedasticity, a value of q=4 was specified, a level that allows the model to account for possible heteroskedasticity or outlying observations.

Having to cope with the simultaneity of *Y*, the parameter estimation was based on a two-step procedure analogous to two-stage least squares, the only difference being that parameter estimation in the second stage regression is not carried out by OLS, which is inconsistent due to the inclusion of the lagged dependent variable.

More specifically, in the first stage the reduced form equation for *Y* is estimated by OLS:

$$Y = F\gamma_1 + X\gamma_2 + E \tag{7}$$

and predicted values are computed as $\hat{Y} = F\hat{\gamma}_1 + X\hat{\gamma}_2$. In the second stage $\hat{Y}$ is substituted for *Y* in (6), yielding:

$$D^* = \rho W D^* + F\lambda + \alpha\hat{Y} + \tilde{U}_1^* \tag{8}$$

where $\tilde{U}_1^* = U_1^* + \alpha(Y - \hat{Y})$. The $\hat{Y}$ variable, having been obtained as the projection of *Y* on the set of the exogenous variables in the model, is orthogonal by construction to $\tilde{U}_1^*$, making it possible to estimate parameters in (8) using standard spatial econometric techniques. In this case, following the chosen Bayesian approach, estimates were obtained by means of the MCMC (Monte Carlo Markov Chain) method, as implemented by LeSage (1997; 2000).

Regression results, reported in Table 3, appear to provide some support for the hypothesis that the strength of family institutions (ISTFAMI) has exerted a positive influence on manufacturing business agglomeration. The estimated coefficient is always significant and has the expected sign with respect to all four LDIID proxies considered. At the same time provinces denoted by a higher level of social opportunism (ISTOPPO) have experienced a lower level of LDIID, the effect being statistically significant according to the results obtained with the first three specifications of the dependent variable, with a slight increase in the *p-value* of the coefficient when a measure closer to the simple notion of spatial concentration of industrial activity (DIST4) is used as a proxy for LDIID.

The influence of local institutions (ISTELOC) is never significant in any of the four regressions, while collective action (ISTAZCO) is significant, with the expected sign, only when LDIID is proxied by DIST2, i.e. when a more restrictive definition is adopted.

Geographical dummy variables are mostly significant, pointing to the existence of some remaining environmental factors, not explicitly considered in the analysis and controlled for in this way, that affect industrial agglomeration in the observed areas.[11]

Table 3

Cross-section regression results including spatial spillover effects

	Functional form			
	Tobit	Tobit	Linear	Tobit
Regressors	DIST1	DIST2	DIST3	DIST4
Constant	-187.94	-83.02	-84.99	-23.87
ISTFAMI	1.97**	0.25**	1.63**	0.25**
	(0.011)	(0.047)	(0.018)	(0.047)
ISTOPPO	-10.00**	-11.65***	-7.06***	-1.11
	(0.021)	(0.000)	(0.008)	(0.157)
ISTAZCO	1.94	6.76**	-1.48	0.129
	(0.351)	(0.018)	(0.330)	(0.444)
ISTELOC	48.48	-2.61	-17.80	-2.61
	(0.131)	(0.377)	(0.482)	(0.377)
VAFIT	0.006**	0.004***	0.003***	0.001*
	(0.037)	(0.047)	(0.055)	(0.069)
DUMSUD	61.31**	33.65**	26.48*	9.45*
	(0.022)	(0.019)	(0.067)	(0.058)
DUMCEN	28.37**	16.08**	10.46	5.82**
	(0.029)	(0.015)	(0.143)	(0.021)
ρ	0.27***	0.27***	0.47***	0.36***
	(0.000)	(0.007)	(0.000)	(0.000)
R^2	0.50^	0.33^	0.62	0.25^
Number of observations	85	85	85	85
Number of censored obs.	27	31	0	16

Figures in brackets are *p-values*. *, **, and *** respectively denote significance at the 10, 5 and 1 per cent levels.
^ A pseudo-R^2 is displayed for Tobit specifications, computed as the squared correlation between observed and fitted values of the dependent variable:

[11] In evaluating the sign and magnitude of the coefficients estimated for the geographical dummies it has to be considered that these variables interact and are, indeed, highly co-linear with the spatially lagged values of the endogenous variable.

Per capita GDP, instrumented as explained above, always has the expected positive coefficient, which is also statistically significant; the same applies to the coefficient of the lagged endogenous variable.

The latter evidence is consistent with the existence of a contagion process, spreading the impact of local factors across different locations through a so-called spatial multiplier mechanism (Anselin, 2003). A positive shock to factor endowments of a given area will thus trigger an increase in the dependent variable not only on the same location but on the other ones as well, albeit with a magnitude that is decreasing as the distance between locations increases.

5. Conclusions

While being well aware that economic development is not a deterministic process, that each local community has experienced a somewhat specific path to growth, and that joint availability of individual factors can have an influence on the local diffusion of industrial districts over and above that produced by the single variables themselves (Viesti, 2000), we think it useful to attempt some generalizations based on the results of the econometric analysis.

The main findings appear to be the confirmed roles of strong family institutions and of a social environment that reduces individual opportunism. in fostering local development through industrial agglomeration. There is no clear evidence, except when a more restrictive definition of industrial district is adopted, of a positive influence of a larger endowment in relational capital. Local government intervention, as measured by the proxy utilised in the study, does not appear to have played an important role in fostering the development of industrial districts.

Apart from providing some empirical support to the thesis of the relevance of social institutions in promoting industrial agglomeration in Italy, the spatial econometric analysis revealed positive and significant spillover effects across neighbouring areas. This evidence is consistent with the existence of a contagion process of the kind set forth, for example, in Brusco and Paba (1997). The existence of such a spatial interaction mechanism implies that, to promote the growth of industrial districts in a given area, not only the local degree of institutional development is important, but also that recorded in surrounding areas.

APPENDIX – DESCRIPTION OF THE INDICATORS[12]

ISTFAMI – Source: ISL-University of Parma database.

Percentage share of 'extended' families out of total families in 1951 (Istat,1951). A family is considered 'extended' when it is composed of the head of the household, his wife, their children, their parents or other relatives.

ISTOPPO – Source: Arrighetti *et al.* (2001).

The indicator is the first component, explaining 62 per cent of total variance in a principal component analysis carried out on the following variables, provided by the ISL-University of Parma database:

- the number of protests on ordinary bills and bank cheques and the number of bills noted for non-acceptance per 1,000 inhabitants in 1958 (Istat, 1958; 1960);
- the number of crimes against patrimony, public goods, industry and commerce denounced, for which the judicial authority has began criminal proceeding, per 1,000 inhabitants in 1958 (Istat 1958; 1960).

ISTAZCO – Source: Source: Arrighetti *et al.* (2001).

The indicator is the first component, explaining 45 per cent of total variance in a principal component analysis carried out on the following variables, also provided by the ISL-University of Parma database:

- the rate of membership in craft associations in 1970, calculated as the ratio between the number of craft firms belonging to their trade associations and the number of artisan firms on the official register;
- the ratio of the total number of valid votes cast by artisans in the commission elections to the total of officially registered artisans in 1970. Source: "Giorgio Coppa" historical archive kept by Confederazione Nazionale dell'Artigianato;

[12] Our analysis does not use information from the regions which experienced changes in the provincial distribution in the 1950s (Friuli Venezia Giulia, Molise and Sardinia) and figures from the provinces created after 1995: Biella, Verbano-Cusio-Ossola, Lodi, Lecco, Rimini, Prato, Crotone and Vibo Valentia. Our sample represents 92.2 per cent of the whole Italian population and 92.5 per cent of the total value added in 2000, according to local area accounts provided by Istat.

- the percentage share of agricultural firms supplying products to agricultural cooperatives or similar entities out of the total number of agricultural firms existing in 1970 (Istat, 1974);
- the percentage share of members of collective purchase groups or voluntary commercial unions out of the total number of commercial licences in official registers in 1965 (Source: Minister of Industry, Commerce and Handicraft, 1966);
- a dummy variable that equals one if in the province there was a credit guarantee consortium belonging to Artigianfidi created before 1975 Source: Artigianfidi Research Unit.

ISTELOC – Source: ISL-University of Parma database

The variable is the mean in the period 1961-1963 of the share of money paid by the municipality to finance public works and education in the total amount of payments granted by the municipalities for capital charges, general expenses, public order, public health and justice (Source: Istat, 1962; 1963).

VAFIT – The predicted values of the regression of per capita provincial GDP in 2000 (Source: Istat) regressed on the exogenous variables of the model (dummies + ISTAMI, ISTOPPO, ISTAZCO) and three instrumental variables: value added per worker in the agricultural, services and construction sectors.

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