ISTITUTO DI STUDI E ANALISI ECONOMICA

Specialization and Risk Sharing: evidence from European Regions

by

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Working paper n. 122
December 2009

Stampato presso la sede dell’Istituto

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ABSTRACT

Economic theory emphasizes that risk sharing makes it possible to exploit benefits from comparative advantages and economies of scale. Unlike previous studies we reject the assumption of parameter homogeneity across geographical units in measuring risk sharing. The estimated regional-specific index of risk sharing is then used as a covariate in a model of industrial specialization for the EU15 regions. By estimating a number of nonparametric additive spatial autocovariance models, allowing for nonlinearities and spatial dependence, we show that industrial specialization is positively affected by risk sharing measures even controlling for other relevant regressors.

Keywords: Risk sharing, specialization, European regions, non-parametric methods, spatial econometrics

JEL classification: E21, F15, O40, C14, C31

NON-TECHNICAL SUMMARY

According to international trade models, higher specialization levels can be reached by exploiting comparative advantages in technology and endowments or through economies of scale and agglomeration effects. Most observers posit however that higher specialization levels create greater vulnerability to idiosyncratic shocks and are, thus, likely to foster asymmetric developments and differences in growth rates across economies. This is only part of the whole story, however. Effective inter-regional insurance mechanisms or “risk sharing” (through well-functioning redistributive fiscal transfers and developed credit and financial markets) may help “protect” the economic environment against idiosyncratic shocks, even in the presence of diverging economic structures. We may therefore expect higher specialization levels to be reached when more inter-regional insurance is achieved.

Although a number of studies have focused on explaining specialization patterns in the European regions, to date there has been no comprehensive study on the effects of cross-regional insurance mechanisms on the degree of specialization in the context of the EU15 regions. This paper is an attempt to fill this gap. A regional perspective in studying the effect of risk sharing on specialization is motivated by two main considerations: *i*) even in the presence of a well established empirical regularity of little international risk sharing, cross-regional insurance mechanisms may operate at a less aggregate level, since the degree of social and economic cohesion within countries is probably higher than between countries; *ii*) regional economies are more vulnerable to external shocks and the probability that sector-specific shocks are asymmetric is much higher at a regional level.

Using data on 144 European NUTS2 regions belonging to the EU15 countries, we estimate a model of regional specialization by adopting a three-step strategy. First, we construct the dependent variable as a measure of regional specialization using the median of Balassa indices. Second, unlike previous studies, we provide a regional-specific measure of the degree of insurance, which is the key explanatory variable. Third, we test whether the data evidence a positive effect exerted by the degree of risk sharing on the level of specialization within a innovative framework which enables us to jointly control for: *i*) further observable factors potentially of relevance in explaining specialization; *ii*) unobservable factors; *iii*) possible endogeneity bias; *iv*) nonlinearities; and *v*) spatial dependence. Our results corroborate the hypothesis that risk sharing positively affects the degree of specialization, even when controlling for a number of causative determinants suggested by the

relevant literature. In the case of European regions, however, this effect is strongly nonlinear.

SPECIALIZZAZIONE E *RISK SHARING* NELLE REGIONI EUROPEE

SINTESI

La teoria economica sottolinea come meccanismi di condivisione del rischio (*risk sharing*) permettano di sfruttare i benefici derivanti dai vantaggi comparati e dalle economie di scala. Contrariamente alla letteratura esistente, in questo lavoro si pone a verifica (e si rifiuta) l'ipotesi di omogeneità dei parametri tra unità geografiche (le regioni europee) nella misurazione del *risk sharing*. Il valore stimato di *risk sharing* regionale è successivamente utilizzato come variabile esplicativa in un modello di specializzazione industriale. Attraverso la stima di una serie di modelli non-parametrici con dipendenza spaziale, si mostra come il grado di specializzazione industriale sia influenzato positivamente dalla misura di *risk sharing* anche tenendo conto dell'effetto di altre variabili esplicative suggerite dalla letteratura.

Parole chiave: Risk sharing, specializzazione, regioni europee, metodi non-parametrici, econometria spaziale

Classificazione JEL: E21, F15, O40, C14, C31

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1 INTRODUCTION

Understanding specialization patterns is of crucial importance for academics and policy makers. According to the theoretical paradigm provided by international trade models, higher specialization levels can be reached by exploiting comparative advantages in technology and endowments or through economies of scale and agglomeration effects (Helpman and Krugman, 1985; Krugman, 1991; Fujita et al., 1999). Furthermore, as suggested by Greenwood and Jovanovic (1990), Acemoglu and Zilibotti (1997) and Feeney (1999), specialization may have major consequences in boosting economic growth. On the other hand, most observers posit that higher specialization levels create greater vulnerability to idiosyncratic shocks and are, thus, likely to foster asymmetric developments and differences in growth rates across economies. This is only part of the whole story, however. Effective inter-regional insurance mechanisms or “risk sharing” (through well-functioning redistributive fiscal transfers and developed credit and financial markets) may help “protect” the economic environment against idiosyncratic shocks, even in the presence of diverging economic structures.

Although a number of studies have focused on explaining specialization patterns in the European regions (Molle, 1997; Hallet, 2002), to date there has been no comprehensive study on the effects of cross-regional insurance mechanisms on the degree of specialization in the context of the EU15 regions. This paper is an attempt to fill this gap. Our investigation is also closely related to the strand in the literature which seeks to measure risk sharing (Asdrubali et al. 1996; Sørensen and Yosha, 1998; Mélitz, 2004; among others) and in particular to the paper by Kalemli-Ozcan et al. (2003), which represents, to the best of our knowledge, the sole empirical study on the relationship between industrial specialization and degree of cross-regional insurance produced to date.

A regional perspective in studying the effect of risk sharing on specialization is motivated by two main considerations: *i*) even in the presence of a well established empirical regularity of little international risk sharing (the so-called “home bias” phenomenon pointed out by French and Poterba, 1991), cross-regional insurance mechanisms may operate at a less aggregate level, since the degree of social and economic cohesion within countries is probably higher than between countries (Cochrane, 1991); *ii*) regional economies are more vulnerable to external shocks and the probability that sector-specific shocks are asymmetric is much higher at a regional level (De Nardis et al., 1996).

Using data on 144 European NUTS2 regions belonging to the EU15 countries, we estimate a model of regional specialization by adopting a three-step strategy. First, we construct the dependent variable as a measure of regional specialization using the median of Balassa indices. Second, unlike previous studies, we provide a regional-specific measure of the degree of insurance, which is the key explanatory variable. Third, we test whether the data evidence a positive effect exerted by the degree of risk sharing on the level of specialization within a innovative framework which enables us to jointly control for: *i*) further observable factors potentially of relevance in explaining specialization; *ii*) unobservable factors; *iii*) possible endogeneity bias; *iv*) nonlinearities; and *v*) spatial dependence. The econometric results corroborate the hypothesis that industrial specialization is positively affected by risk sharing measures even controlling for other relevant regressors.

The rest of the paper is structured as follows. Section 2 illustrates the conceptual framework and the assumptions on the determinants of industrial specialization. Then discussed are the data sources and details on variable construction. The econometric methodology and estimation results are presented in Section 4. Final remarks conclude.

2 SPECIALIZATION AND ITS DETERMINANTS

2.1 Measuring specialization

Regions are said to be specialized when a limited range of industries dominate their production activities. As discussed in Combes and Overman (2004), various indicators of the degree of specialization (such as, for example, the Herfindahl, Theil and Gini indices) have been proposed in the literature. However, none of these measures can be said to be optimal and they prove to be strongly correlated.[1] A more critical issue is instead the choice of the variable used to construct the specialization index. Several candidates (such as value added, export and employment) can be considered. While the use of employment shares entails observing specialization from the input standpoint, the choice of production quantities (value added or export) means taking an

1 In our empirical exercise, we have found pairwise correlation among various indices of specialization ranging from 0.6 and 0.9.

output-based perspective. Given the alternatives, employment data rather than production data should be preferred because they are less sensitive to valuation problems.

2.2 Main determinants

While the possible role of cross-regional insurance mechanisms in explaining industrial specialization patterns was neglected until the paper by Kalemli-Ozcan et al. (2003), other causative factors have been widely employed in this strand of the literature. In keeping with the most recent empirical contributions, we include measures of the level of I) *exposure to risk* (namely, risk sharing and volatility); II) *socio-economic conditions* (degree of development and the size of the economy); III) *structural characteristics* of the manufacturing sector (its share over total GVA) as candidate variables in explaining the sectoral concentration of productive structures in European regions. There follows discussion of these candidate explanatory factors.

Risk sharing. Economic theory maintains that pursuing risk sharing makes it possible to spread production risk among regions and to achieve a higher degree of specialization by exploiting otherwise idle comparative advantages (Helpman, 1981) or new economic opportunities (Obstfeld, 1994; Murdoch, 1995). When individuals are unable to borrow or insure, they tend to increasingly mitigate risk by choosing safer production techniques or by forsaking specialization for a more sectorally diversified range of productions (self-insurance). By contrast, well-functioning cross-regional insurance markets are central mechanisms with which to smooth idiosyncratic shocks without necessarily implying self-insurance. A first channel to achieve cross-regional insurance is an efficient system of redistributive fiscal transfers (such as unemployment benefits). The finance literature suggests portfolio diversification as another viable channel to buffer asymmetric shocks through inter-regional insurance mechanisms (Asdrubali et al., 1996).

GDP volatility. Under the hypothesis of no uncertainty, full risk sharing can be achieved and each region will specialize in a different technology so as fully to exploit the economies of scale in production. By contrast, facing uninsurable risks discourages agents from taking production risks (Ramey and Ramey, 1995; Acemoglu and Zilibotti, 1997), which suggests a negative effect of output volatility on regional specialization.

Stage of development. A recent body of literature posits that specialization is likely to change over the development path in a nonlinear way. In Imbs and Wacziarg (2003), the link between overall specialization and the level of income per capita follows a U-shaped pattern. At low levels of per capita income, economies are forced to specialize in natural resources. Subsequently, they diversify (reducing their degree of overall specialization), but re-specialize once a relatively high level of income per capita has been reached, in a way consistent with models featuring endogenous stages of specialization to both trade and economic growth (Saint-Paul, 1992). By contrast, Kim (1995) documents an inverted U-shaped relationship for the US regions: in the early stages of national growth, a steady increase in regional specialization is observed; thereafter, a decrease in specialization takes place. De Benedictis et al. (2009) reach similar conclusions. Thus, the shape of such nonlinearity is ambiguous and it should be subjected to empirical scrutiny.

Size of the economy. In accordance with the New Economic Geography paradigm (Krugman, 1991; Fujita et al., 1999), agglomeration economies (such as market size effects) may promote diversification through the attraction of industries to larger regions. Kalemli-Ozcan et al. (2003) explain a negative effect of the economy's size on the degree of specialization with demographic argumentations (for instance, heterogeneity of the population) as well as within-region geophysical characteristics (climate, landscape and natural resources). The empirical evidence, however, is mixed, with estimation findings dependent on the methodology adopted and on the dataset used.[2]

Share of the manufacturing sector. A further candidate explanatory variable is the share of manufacturing on total output. Economic intuition suggests that when manufacturing is relatively small with respect to other economic sectors, regions are likely to be specialized in a few sectors. By contrast, when the share of manufacturing on total GVA increases, regions may be able to afford a broader range of industrial productions. On this reasoning, we expect the manufacturing/total GVA ratio to have a negative effect on the degree of the industrial specialization.

2.3 The role of spatial externalities and nonlinearities

In order to demonstrate that risk sharing positively affects specialization, we use an empirical framework which allows for both spatial externality effects and nonlinearities. The New Economic Geography paradigm (Krugman, 1991, Fujita *et al*., 1999) posits that a number of agglomeration economies (such as forward/backward linkages and knowledge spillovers) affect decisions on

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2 For instance, Kalemli-Ozcan et al. (2003) and Ezcurra et al. (2004) find that the expected negative relationship is empirically confirmed; by contrast, Imbs and Wacziarg (2003) document that the size of the economy does not have a role in explaining specialization patterns.

industrial location and productive specialization. Obviously, there is no reason to expect these centripetal forces to operate only within the administrative boundaries of the regions. Admitting regional interactions and spillovers requires a modeling approach able to capture possible spatial externality effects (Midelfart-Knarvik et al., 2001).

Furthermore, we cannot assume that all economies follow a common linear pattern, i.e. that the effect of risk sharing and of the other determinants of regional specialization is globally linear.[3] We relax the assumption of linearity in order to avoid possible misspecification problems and to jointly model nonlinearities and interaction effects in a very flexible framework. Finally, our empirical design seeks to control for unobservable factors and to deal with endogeneity problems due to simultaneity and/or measurement errors.

3 DATA SOURCES AND VARIABLE CONSTRUCTION

3.1 Specialization measures

As explained in Section 2.1, we use sector employment shares to construct an indicator of overall specialization. Employment shares are computed using annual data at the NUTS2 level for 144 European regions belonging to the EU15 countries over the period 1995-2005. The Eurostat Regio dataset provides information on the number of employees at two-digit level of the classification of economic activity for the period 1995-2005.[4] Any other source of data (for instance, Cambridge Econometrics) provides employment or GVA regional data at a more aggregate sectoral level and/or over a shorter temporal window.

Our preferred specialization measure is an index of the position of the distribution of Balassa index. This proxy for regional specialization has the

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3 For instance, a quadratic term for per capita GDP is usually considered in the specification of a regional specialization model by Imbs and Wacziarg (2003) and by Kalemli-Ozcan et al. (2003).

4 Namely, sector DA (food products, beverages and tobacco), DB (textile and textile products), DC (leather and leather products), DD (wood and wood products), DE (pulp, paper and paper products; publishing and printing), DF (coke, refined petroleum products and nuclear fuel), DG (chemicals, chemicals products, and man-made fibres), DH (rubber and plastic products), DI (other non-metallic mineral products), DJ (basic metals and fabricated metal products), DK (machinery and equipment n.e.c.), DL (electrical and optical equipment), DM (transport equipment) and DN (manufacturing n.e.c.).

advantage of being directly derived from a measure of sectoral revealed comparative advantages (RCA) as documented by De Benedictis and Tamberi (2004). Formally:

$$y_i = -Me\left(\frac{E_{is}}{E_s} / \frac{E_i}{E}\right) \tag{1}$$

where E_{is} stands for average employment in the s-th sector for the i-th region over the period 1995-2005, E_i is the average overall employment in the i-th region; E_s indicates the employment in the s-th sector in Europe, while E is the overall European employment.

Since the RCA index follows an asymmetric distribution (with a fixed lower bound, 0, and a variable upper bound, E/E_i), its median, $Me(.)$, proves to be the most appropriate indicator of the distribution position. When $Me(.)$ is low, an economy shows a comparative advantage in a large share of sectors, so that its productive structure is diversified; *viceversa,* when $Me(.)$ is high, an economy is specialized. Hence we use the opposite median as a direct measure of specialization.

3.2 Risk sharing measures and other determinants of specialization patterns

Annual data over the period 1980-2003 for GDP values at 1995 euros, levels of population, manufacturing and total gross value added (GVA) for our sample of 144 European NUTS2 regions are taken from the Cambridge Econometrics database.[5]

Region-specific risk sharing index. Following Asdrubali et al. (1996), an index of cross-regional insurance can be computed by estimating:

$$\Delta gdp_{it} - \Delta c_{it} = \gamma_t + \beta \Delta gdp_{it} + \xi_{it} \;\; i = 1,\ldots,N(=144) \;,\; t = 1981,\ldots,2003 \tag{2}$$

where Δ is the first difference operator, gdp_{it} and c_{it} are the logarithms of real per capita GDP and consumption, respectively, at time t for region i, γ_t

5 As in Kalemli-Ozcan et al. (2003), specialization measures and candidate explanatory variables are computed on partially overlapping temporal windows. Moreover, in order to limit measurement error problems in the estimation of our proxy of risk sharing, we use the longest time span available from Cambridge Econometrics data sources.

is a vector of time fixed effects which allows us to control for European wide business cycle and trends in the average values and ξ_{it} is a vector of error terms. The values for the index of risk sharing, β, are expected to lie in the [0,1] interval. The comovement of per capita consumption with per capita GDP fluctuations, $\delta \equiv (1-\beta)$, measures the fraction of idiosyncratic GDP shocks that is not eliminated through insurance: if consumption is perfectly insured, it is not affected by fluctuations of GDP, implying $\delta = 0$. When $\delta = 1$, there is a perfect match between per capita consumption and per capita GDP dynamics, so that there is no inter-regional insurance at all. Finally, when consumption is only partially insured, $\delta > 0$.[6]

We estimate model (2) by Seemingly Unrelated Regression (SUR), which allows for heterogeneous slopes (β_i) across spatial units taking cross-section error correlation into account. In all entities of reference, the estimated parameters turn out to be statistically different from zero at the 5 percent significant level or better, with an estimated average value of $\hat{\beta}_i^{SUR}$ equal to 0.528. The average magnitude of $\hat{\beta}_i^{SUR}$'s is remarkably higher than the estimates documented in hitherto the literature and in a manner consistent with the idea that higher risk sharing takes place when smaller geographical units are taken into account (Cochrane, 1991).

Other determinants. Besides a regional risk-sharing index, our set of candidate explanatory variables includes a number of regressors, whose construction is detailed below. These variables are constructed as averages of yearly observations over the period 1980-2003. Following Kalemli-Ozcan et al. (2003), GDP volatility, vol_i, is measured by the standard deviation of the first differences of the logarithm of GDP. In keeping with previous studies (Kalemli-Ozcan et al., 2003; De Benedictis et al., 2009; Imbs and Wacziarg, 2003; Ezcurra et al., 2004), we employ gdp_i as a proxy for the degree of economic development. We also include (the logarithm of) regional population, pop_i, in order to measure regional size. Finally, (the logarithm of) the share of the manufacturing sector on total GVA, man_i, is considered.

6 Equation (2) is intended to capture overall insurance mechanisms as in Asdrubali et al. (1996). Although it would have been advisable to identify the (credit, financial and fiscal) channels of risk sharing, the temporal slightness of regional data on real per capita disposable income series (1995-2003) prevented us from computing the amount attributable to the fiscal channel and to the other channels.

For the ease of interpretation of the econometric results discussed below, all quantities are standardized by computing the deviations from the European average and dividing by their standard deviation.

4 ESTIMATION RESULTS

4.1 Econometric framework

As discussed in Section 2.3, modelling regional specialization requires a flexible approach which allows for nonlinearity and spatial dependence. Nonlinearities could be captured by a polynomial regression model, as in Imbs and Wacziarg (2003) and in Kalemli-Ozcan et al. (2003). We instead use a semiparametric methodology, since it is much more flexible than any parametric specification. By using a particular version of the semiparametric model that allows for additive components, we are able to obtain graphical representation of the relationship between regional specialization and regional characteristics. Additivity ensures that the effect of each of the model predictors can be interpreted net of the effect of the other regressors, as in linear multiple regression. A typical semiparametric additive model (AM) is specified as follows:

$$y_i = X_i^{*'}\alpha^* + f_1(x_{1i}) + f_2(x_{2i}) + f_3(x_{3i}, x_{4i}) + ... + \varepsilon_i \tag{3}$$

where ε_i is a vector of independently, identically and normally distributed errors, $\varepsilon_i \sim iidN(0, \sigma_\varepsilon^2)$, $f_j(\cdot)$ are unknown smooth functions of the covariates, X_i^* is a vector of strictly parametric components (also including country dummies so as to partially capture unobserved heterogeneity) and α^* is the corresponding parameter vector. For our analysis, we employ the methodology proposed by Wood (2006) to estimate AMs with spline based penalized regression smoothers which allow for automatic and integrated smoothing parameters selection via Generalized Cross Validation (GCV).

In order to control for spatial interaction effects, Model (3) has to be augmented. As pointed out by Anselin (2004), spatial externalities may occur either in unmodeled effects (when unmodeled variables subsumed in the error term jointly follow a spatial random process) or in modelled effects (when the exogenous terms affect the left hand side of the model through a '*global multiplier effect*'). In a parametric linear setting, such as $y = X'\alpha + \varepsilon$, global multiplier effects are modelled by replacing X and ε with $(I - \rho W)^{-1} X$ and $(I - \rho W)^{-1}\varepsilon$, respectively, where I is an identity matrix, ρ is the parameter of spatial externality and W is a spatial weights matrix.[7] In the present context, the inverse spatial transformation of X and ε suggests that the attractiveness of region i is affected not only by its own characteristics and random shocks, but also by the features and random shocks of all other regions. However, given the characteristics of the standardized spatial weights matrix, the strength of spatial dependence between observed regions declines with the distance between them. In other words, neighbouring units exhibit a higher degree of spatial contagion than do units located far apart ('*spatial diffusion with friction*'). The introduction of the spatial multiplier effect in the model yields a reduced form as $y = (I - \rho W)^{-1} X'\alpha + (I - \rho W)^{-1}\varepsilon$ and the structural form becomes the standard spatial autoregressive model (SAR) $y = \rho W y + X'\alpha + \varepsilon$. These arguments can be extended to the semiparametric AMs, with the obvious difference that the effect of spatial externalities may not be homogenous over space. Hence equation (3) can be extended by including the smooth term $f_4(\sum_{j \neq i} w_{ij} y_j)$ on the right hand side (SAR-AM):

$$y_i = X_i^{*'}\alpha^* + f_1(x_{1i}) + f_2(x_{2i}) + f_3(x_{3i}, x_{4i}) + f_4(\sum_{j \neq i} w_{ij} y_j) + \ldots + \varepsilon_i \qquad (4)$$

Because of the feedbacks between y_i and its spatial lag term $\sum_{j \neq i} w_{ij} y_j$, $f_4(\cdot)$ enters endogenously into equation (4), that is $f_4(\cdot)$ and ε_i are correlated. In order to deal with endogeneity problems in the estimation of nonparametric models, we use the procedure proposed by Blundell and Powell (2003), which consists of extending the "control function" method to semiparametric models through a two-step procedure. In the first one, an auxiliary semiparametric regression $\sum_{j \neq i} w_{ij} y_j = X_i^{*'}\alpha^* + f_1(x_{1i}) + f_2(x_{2i}) + f_3(x_{3i}, x_{4i}) + h(Z_i) + \ldots + \upsilon_i$ is

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7 The characteristic element of this matrix, w_{ij}, summarizes the interaction between regions i and j. Throughout the paper, the $W = \{w_{ij}\}_{i,j=1,\ldots,N}$ matrix is specified so that w_{ii} are set to zero whereas $w_{ij} = d_{ij}^{-2}$ if $d_{ij} < \bar{d}$ and $w_{ij} = 0$ if $d_{ij} > \bar{d}$, with d_{ij} being the great circle distance between the centroids of region i and region j and $\bar{d}$ the cut-off distance (equal to 424 km). The results are robust to the alternative choices of the spatial weights matrix.

considered, with Z_i being a set of conformable instruments and υ_i a sequence of random variables satisfying $E(\upsilon_i \mid Z_i) = 0$. Moreover, if Z_i and ε_i are independent, then it follows that $E(\varepsilon_i \mid \upsilon_i, Z_i) = E(\varepsilon_i \mid \upsilon_i)$ and, thus, $E(\varepsilon_i \mid \sum_{j \neq i} w_{ij} y_j) \neq 0$ when $E(\varepsilon_i \mid \upsilon_i) \neq 0$. The second step consists of estimating an AM of the form:

$$y_i = X_i^{*'} \alpha^* + f_1(x_{1i}) + f_2(x_{2i}) + f_3(x_{3i}, x_{4i}) + f_4(\sum_{j \neq i} w_{ij} y_j) + f_5(\hat{\upsilon}_i) ... + \varepsilon_i \qquad (5)$$

Another source of bias is the inclusion of variables measured with error, such as our proxy of risk sharing. Since this variable is an estimated coefficient, we cannot exclude the existence of a correlation between risk sharing and the error term. Again, the control function approach can be used to take account of this problem.

4.2 Model selection

Tables 1 and 2 report the estimation results and a battery of diagnostics tests for different parametric and semiparametric specifications of the model of regional specialization in Europe. In our econometric investigation all variables are weighted by the regional population so as to reduce the possible impact of small highly specialized regions. Moreover, country dummies are included in all models to control for residual spatial heterogeneity and for possible unobservable country-specific factors. Furthermore, the presence of Wy (in matrix form) and $\hat{\beta}^{SUR}$ in the set of covariates suggest including two additional terms, $\hat{\upsilon}_1$ and $\hat{\upsilon}_2$, which represent the estimated residuals from two distinct first step estimations.[8]

Model A in Table 1 resembles a specification widely used in previous studies, where the regression function is linear in all terms except gdp which is assumed to have a quadratic effect on y. A significant positive effect of risk sharing on regional specialization emerges, corroborating the findings of Kalemli-Ozcan et al. (2003). Note, however, that all other regressors are statistically not significant according to p-values associated with White-

8 The set of additional instruments for the two auxiliary regressions are the spatial lags of exogenous terms, an indicator of “financial depth”, measured by the share of the financial and real estate sectors on total GVA, and the investor protection index provided by the World Bank.

corrected robust standard errors, except for gdp as well as its square term and the manufacturing share. Furthermore, traces of endogeneity for the risk sharing parameter are apparent, while no clear spatial dependence is found.[9] As for diagnostic tests, we find not only some departures from the normality assumption, but also from linearity. This encourages us to adopt a semiparametric modelling strategy.

Table 1 **Estimation results: Model A**

	Coefficients	p-values
$\hat{\beta}^{SUR}$	0.913	0.000
vol	0.062	0.679
gdp	0.453	0.013
gdp^2	0.349	0.004
pop	0.013	0.967
man	-0.158	0.085
Wy	0.226	0.490
$\hat{v}_1$	-1.191	0.018
$\hat{v}_2$	-0.690	0.007
R^2-adj	0.324	
F test 1)	6.224	0.000
F test 2)	1.608	0.064
Normality	0.982	0.054
Heteroskedasticity	44.533	0.004
Linearity	1.812	0.024

Notes: The dependent variable, y, is the index of specialization. 'R^2-adj' is the determination coefficient adjusted for the degrees of freedom. 'F test 1) and 2)' indicate the test for the joint significance of additional instruments in the first step of the model. The 'Normality' test is based on Shapiro-Francia statistics. 'Heteroskedasticity' is Koenker's studentized version of the Breusch-Pagan test against heteroskedasticity. 'Linearity' is the statistics of the RESET test. All models include a full set of country dummies. p-values refers to White-corrected robust standard errors. The number of observations is 144.

9 The F test for the overall significance of the additional instruments confirm the validity of our set of instruments. Furthermore, the Sargan test gives a statistics of 20.934 with a p-value of 0.283, indicating no correlation between the instruments and the error term.

Table 2 shows the results and diagnostics for three different semiparametric specifications: in Model B all terms enter nonlinearly but additively, while Models C and D allow for some interactions between variables. When allowing for nonlinearity (Model B), the model fit improves significantly, with the adjusted R^2 increasing from 0.32 to 0.54. The F-tests for the overall significance of the smoothed terms in Model B have p-values lower than 0.05 in four out of eight cases, while the number of effective degrees of freedom (edf) suggests that the relationship between regional specialization and its determinants is far from being linear, except for the spatial lag term. This more flexible specification, however, does not allow recovery of significance for Wy, which turns out to be strongly endogenous. The lack of significance for population density as well as for volatility suggests possible collinearity or concurvity problems which call for interaction effects.

In order to address the issue at stake, we consider a number of alternative specifications which allow for smooth interaction terms. Adopting the same taxonomy as in Section 2 above, we test for the joint effect of measures of: I) *exposure to risk*, $f(\hat{\beta}^{SUR}, vol)$; II) *socio-economic conditions*; $f(gdp, pop)$, as causative determinants of the sector concentration of productive structures in European regions. After considerable experimentation, we opted for a specification (Model C) which admits the joint smooth effect of *exposure to risk* (risk sharing and volatility) and the joint smooth effect of *socio-economic conditions* variables (size of the region and its economic development) and the univariate smooth term $f(man)$. Furthermore, the spatial lag term enters linearly, while only significant country dummies are retained (namely, Italy, Greece and Austria).

The two interaction terms are significant at the 1 percent level and the edf clearly indicates nonlinear effects. The univariate smooth terms are significant and nonlinear as well, except for $f(\hat{\upsilon}_2)$, which is not statistically relevant.[10] The estimation results from Model D, where the term $f(\hat{\upsilon}_2)$ is removed from the set of covariates, indicate, indeed, an improvement in the goodness of fit as well as a decrease in the GCV score together with a more satisfactory performance with respect to diagnostic tests.

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10 This finding gives support for our SUR-based approach to the computation of regional-specific risk sharing indexes, thus ruling out any bias due to the measurement of $\hat{\beta}^{SUR}$'s; in turn, the lack of statistical significance of $f(\hat{\upsilon}_2)$ suggests its exclusion from the set of covariates.

Table 2 **Estimation results: Models B, C and D**

	Model B			Model C			Model D		
	F test	p-values	edf	F test	p-values	edf	F test	p-values	edf
$f(\hat{\beta}^{SUR})$	4.132	0.001	5.894	·	·	·	·	·	·
$f(vol)$	0.389	0.704	1.765	·	·	·	·	·	·
$f(gdp)$	4.153	0.008	2.560	·	·	·	·	·	·
$f(pop)$	0.580	0.578	1.702	·	·	·	·	·	·
$f(man)$	3.371	0.024	2.329	5.114	0.001	3.843	5.037	0.001	3.759
$f(\hat{\beta}^{SUR}, vol)$	·	·	·	2.055	0.030	10.557	2.045	0.033	10.065
$f(gdp, pop)$	·	·	·	4.564	0.000	16.025	4.733	0.000	16.231
$f(Wy)$	1.873	0.168	1.000	·	·	·	·	·	·
Wy	·	·	·	0.467a	0.023	.	0.422a	0.036	.
$f(\hat{v}_1)$	6.235	0.000	4.894	6.826	0.000	5.571	6.539	0.000	5.548
$f(\hat{v}_2)$	1.770	0.093	7.375	1.070	0.342	1.336	·	·	·
R^2-adj		0.540			0.641			0.643	
Deviance		67.4			74.5			76.3	
GCV score		3.586			2.792			2.692	
F test 1)	13.619	0.000		13.619	0.000		14.216	0.008	
F test 2)	1.976	0.011		1.976	0.011		.	.	
Normality		0.299			0.932			0.724	
Constant Variance	0.749	0.490	1.153	0.554	0.607	1.096	0.376	0.712	1.048

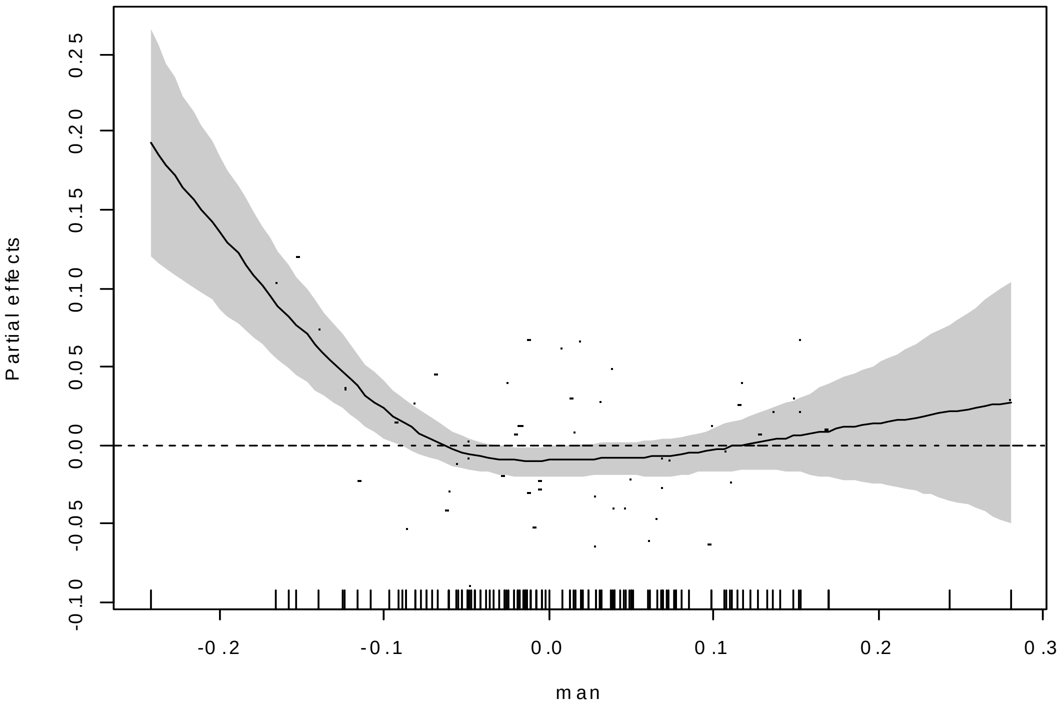
Notes: F tests are used to investigate the overall ("approximate") significance of smooth terms. edf (effective degrees of freedom) reflect the flexibility of the model. An edf equal to 1 suggests that the smooth term can be approximated by a linear term. Deviance is the percentage of explained deviance. The GCV score ($\times 1000$) provides a criterion for choosing the model specification among different possible alternatives; the model which minimizes the GCV is preferred. The test of constant variance of the residuals (Constant Variance) is based on the estimation of the model $|\hat{\varepsilon}| = \alpha + f(\hat{y}) + u$, where $|\hat{\varepsilon}|$ is the absolute value of the residuals of the model and $\hat{y}$ is the vector of fitted values. Under the null of constant variance, the smooth term $f(\hat{y})$ must be estimated with one degree of freedom and, according to a F test, should not have a significant effect on $|\hat{\varepsilon}|$. [a] The value is the coefficient of a linear term. See also Notes in Table 1.

4.3 Partial effects of the smooth terms

In this Section we discuss in some detail the partial effects of univariate and bivariate smooth terms estimated using our preferred specification (Model D).

Figure 1 shows the fitted univariate smooth function (solid line) $f(man)$, alongside Bayesian confidence intervals (shaded grey areas) at the 90 percent level of significance, computed as suggested by Wood (2004). In the plot, the vertical axis displays the scale of the expected values of regional specialization, while the horizontal one reports the scale of the manufacturing share. A nonlinear pattern for $f(man)$ emerges, with a clear downward pattern only up to a threshold corresponding to the European average: when manufacturing is a small fraction of total GVA, the regional producers necessarily concentrate in a few sectors; while as the share of manufacturing increases, the region becomes less specialized. Beyond the European average, the manufacturing share has no effect on the degree of specialization, since the confidence intervals become much larger.

Fig.1 **Partial effects of the univariate smooth term** $f(man)$



[Figures 2a) show the joint effect of risk sharing and GDP volatility - $f(\hat{\beta}^{SUR}, vol)$ -from two different perspectives. In each plot, the vertical axis reads as the previous graph, while the two axes of the horizontal plane report the scale of risk sharing and GDP volatility. Three main remarks ensue. First, the two variables interact negatively: the expected degree of specialization reaches its maximum at the highest level of risk sharing and at the lowest value for volatility; the opposite holds for high levels of vol and low values for $\hat{\beta}^{SUR}$.

Second, for high levels of risk sharing, GDP volatility negatively affects the degree of specialization only up to the European average; beyond that threshold, volatility does not influence specialization. Conversely, when the level of risk sharing is low, the effect of vol is hump-shaped. Third, for high levels of GDP volatility, risk sharing influences specialization almost linearly, while for low levels of vol the effect of risk sharing is nonlinear, because the slope of the curve increases with increasing levels of $\hat{\beta}^{SUR}$. Most importantly, we can safely conclude that the effect of risk sharing is quasi-monotonically positive, in a way fully consistent with our theoretical assumptions.

Figures 2b) display the effect of the interaction between population and per capita GDP, $f(pop, gdp)$. The well-established U-shaped relationship between the degree of specialization and the stage of economic development (e.g. Imbs and Wacziarg, 2003) turns out to be the result of differentiated behaviour between large and small regions. When gdp is below the European average, the expected level of specialization is actually high for small regions. This finding, however, does not hold when larger geographical entities are considered. In this case, indeed, the degree of sector concentration increases with per capita GDP only at later stages of development. As for the effect of population, we observe a negative relationship with the degree of specialization, but the slope of the curve becomes steeper at higher gdp levels. This suggests that if a region is large but relatively poor in terms of per capita GDP, the heterogeneous demand must be satisfied by means of diversified industrial production; when, instead, the region is relatively rich, consumers' needs are more likely fulfilled by interregional trade and comparative advantages can thus be better exploited.

Fig.2 **Partial effects of multivariate smooth terms**

Figure 2a) $f(\hat{\beta}^{SUR}, vol)$

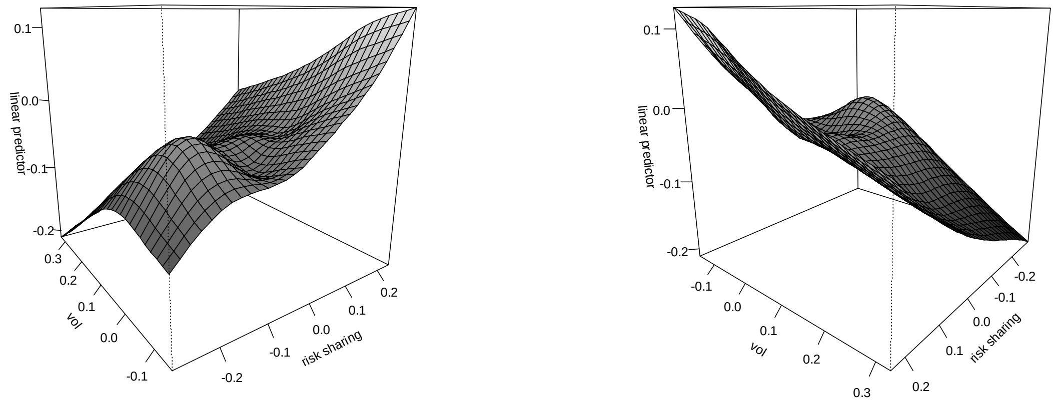
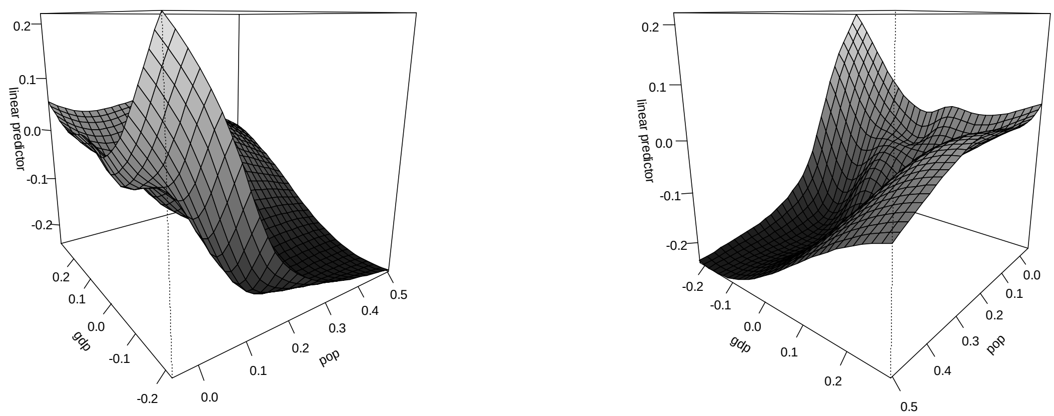


Figure 2b) $f(gdp, pop)$



5 CONCLUSIONS

Using data on 144 European NUTS2 regions belonging to the EU15 countries, we have estimated a regional-specific measure of risk-sharing. Unlike previous studies we have tested (and rejected) the assumption of parameter homogeneity across spatial units, although a clear country pattern has been found. Our estimates document that the overall degree of insurance is about 50 percent of production risk, which is well above the empirical evidence hitherto reported in the literature and in a way consistent with the idea put forward by Cochrane (1991) that substantial interregional risk sharing takes place when smaller geographical units are taken into account. We have then exploited the cross-section heterogeneity in the risk sharing parameters, using these as explanatory variables in a model of regional specialization. Our results corroborate the hypothesis that risk sharing positively affects the degree of specialization, even when controlling for a number of causative determinants suggested by the relevant literature. In the case of European regions, however, this effect is strongly nonlinear.

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