ECONOMIC GEOGRAPHY AND SPATIAL WAGE STRUCTURE IN SPAIN

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Abstract: In this paper, the nominal wage equation of a Core-Periphery model of New Economic Geography is derived. In order to show empirical evidence of the resulting hypothesis, the nominal wage equation is estimated for the Spanish provinces in the year 2003. Our results illustrate that the market access variable is statistically significant and quantitatively important in the explanation of the spatial wage structure observed in Spain. Moreover, we show that there are at least three channels through which market access might be affecting Spanish wages: human capital, productive capital and the size of R&D activities.

Key Words: New Economic Geography, Spatial Wage Structure, Market Access, Spanish Provinces

JEL Classification: R11, R12, R13, R14, F12, F23

1. Introduction

Economic activities tend to cluster at many geographical levels (Florence (1948), Hoover (1948), Fuchs (1962), Enright (1990), Ellison and Glaeser (1996), Dumais *et al*. (1997), Porter (1998, 2000). At a world wide level, there is the North-South dualism where NAFTA (North American Free Trade Agreement- US-Canada-Mexico) the EU-15 countries and East Asia account for 83% of the world's GDP in the year 2000. Furthermore, Hall and Jones (1999) observe that high-income nations are clustered in small cores in the Northern hemisphere and that productivity per capita steadily declines with increasing distance from the core regions (New York, Brussels and Tokyo). Moving to a smaller geographic scale, the so called "blue banana" in the European Union[1] is well known case: a large agglomeration area that takes in Greater Manchester, London, Paris, Northern Italy and the Ruhr Valley. This area represents roughly 20% of the former EU15,

[1]For a comprehensive analysis of the spatial structure of Europe see Faíña *et al.* (2001), Faíña and Lopez-Rodriguez (2006b). A similar case is the "US manufacturing belt" in the East coast of the United States.

but contains 40% of its GDP and 50% of its population (see Lopez-Rodriguez *et al*. 2007). At the country level, the case of the Île-de-France (the metropolitan area of Paris) which accounts for 2.2% of the area of the country and 18.9% of its population and produces 30% of its GDP is a clear example of the clustering of economic activities in a space.[2] In the case of Spain, and in terms of the New Economic Geography literature, a "core-periphery" structure can also be defined. There, the "core" (the triangle area comprising the axis Basque Country-Gerona, Gerona-Valencia and Valencia-Basque Country plus the capital, Madrid) represents a quarter of the total Spanish Peninsular area but concentrates almost 50% of its population and 60% of its GDP; whereas the rest of the country could be characterized in a New Economic Geography fashion as the Spanish "periphery" with less economic activity and a much lower level of per capita income (see Faiña and Lopez-Rodriguez 2006a). Thus, a number of papers have recently presented evidence about the tendency for Spanish economic activity to be geographically concentrated, where core regions are leaving behind the group of peripheral regions (Márquez and Hewings, 2003; Márquez *et al* 2003; and Márquez *et al* 2006).

<<Insert table 1 here>>
<<insert figure 1 here>>

Figure 1 shows the distribution of per capita GDP in Euros in the 47 peninsular Spanish provinces for the years 1995 and 2004 by quantiles. From this graph and from computations in table 1, it can be seen that disparities in terms of per capita GDP among the Spanish provinces are still quite high and there has been little tendency for them to change.

Looking at the ratio between the capital of Spain (Madrid) and the average per capita GDP, the figures for the year 2004 show that the income in Madrid is almost 40% higher than in the average Spanish province. A similar ratio can also be found when the so-called Spanish "core" is compared with the Spanish "periphery." Comparing these computed ratios for the year 2004 with the same computations for the year 1995 (see also figure 1), it is possible to observe that per capita GDP disparities either between Madrid and the average provincial per capita GDP of the "core" and the "periphery" have not changed. On the contrary, there was a slight increase. In addition, these disparities show a well defined income gradient, in the sense that provincial per capita GDP is a decreasing function of the distance from Zaragoza (a proxy for the geographical

[2]Sao Paulo with 18 million population but 15% of Brazilian GDP is another example of a high concentration of economic activities in a space.

center of the so-called Spanish “core”). Figure 2 illustrates the relationship between provincial per capita GDP and distance from Zaragoza for the year 2004.

<<insert figures 2, 3 here>>

Another way of looking at the core-periphery structure is by computing indexes of market potential. Figure 3 represents the territorial structure of Spain based on market potential computations. In this figure, it can also be appreciated that the highest values of market potential are found in the Mediterranean façade along the triangle defined by the Basque Country-Gerona-Valencia plus the capital Madrid, whereas the lowest market potential values are observed on the Atlantic façade.

There are many theories that explain the lack of convergence among countries or regions. From the point of view of growth theories, Barro and Sala-i-Martin, (1991, 1995) show that differences in saving rates, investment rates, human capital levels, sluggish technological diffusion, etc. may prevent income levels from moving closer together. Traditional theories of economic development emphasize the role of *first nature geography (*access to waterways, ports, airports, hydrocarbons, climate conditions) in determining income levels (see Hall and Jones, 1999). In the early 1990s, a new branch of research within the spatial economics began with the pioneering works of Krugman (1991a, b). The so-called New Economic Geography added new insights and provided micro foundations to the explanation of why economic activities are clustered in space. This new line of research which the building blocks are increasing returns to scale at the firm level, transportation costs and imperfect competition emphasize the role of the so-called *second nature geography* (distance to consumer markets and distance to input suppliers) as opposed to *first nature geography* as a way of explaining differences in income levels among regions or countries. Krugman’s (1991a, b) contributions have generated a plethora of theoretical contributions. However, empirical research is still lagging behind. The first empirical attempt to validate the forces at work in the New Economy Geography models at the country level was made by Hanson (1998, 2005) for the United States. Since Hanson’s contributions, many other scholars have tried to test the theoretical predictions of New Economic Geography models for different geographical settings. For instance, Redding and Venables (2004) tested a NEG model for a sample of world countries, Breinlich (2006), Head and Mayer (2006) and Lopez-Rodriguez and Faiña (2007) tested it for different samples of European Union regions. Brackman *et al*.

(2004), Combes and Lafourcale (2004), Roos (2001) and Pires (2006)[3] center their analysis on single countries.

The main goal of this paper is to contrast the theoretical predictions of core-periphery New Economic Geography models about the important role played by *second nature geography* in explaining per capita GDP disparities. The empirical evidence is presented by estimating a derived nominal *wage equation* for the set of the 47 Spanish peninsular provinces in the year 2003. Our results prove to be robust with the established hypothesis, contributing to the empirical literature on New Economic Geography. Moreover, the obtained results show that there are at least three channels through which market access might be affecting Spanish wages: human capital levels, productive capital and the size of R&D activities.

The remaining part of the paper is structured as follows. Section 2 presents the theoretical framework. Section 3 deals with the econometric specifications, data base and variables used in the empirical analysis. In addition, the results and discussions of the econometric estimations are presented. Finally, section 5 offers the main conclusions.

2. Theoretical Background: New Economic Geography and Market Access

The theoretical framework is a reduced version of a standard New Economic Geography model (multi-regional version of Krugman, 1991b) that incorporates the key elements to derive the so-called wage equation and market access. The wage equation will form the basis of the empirical estimations.

Consider a regional setting composed of R locations (j =1, 2………R), focusing on the analysis of the manufacturing sector. In this sector, firms produce a great number of varieties of a homogenous differentiated good (D) under increasing returns to scale and monopolistic competition. Firms face transport costs in an iceberg[4] form in order to receive one unit of the differentiated good at location j from location i , $T_{i,j} > 1$ units must be shipped from i, so $T_{i,j} - 1$ measures the fraction of good that is melted in transit from i to j . The manufacturing sector can

[3] A recent and extensive survey of the empirical literature on New Economic Geography models can be seen in López-Rodríguez and Faiña (2008). Other surveys are those of Overman *et al.* (2003), Combes and Overman (2004) and Head and Mayer (2004).

[4] See McCann (2005) for a discussion on some problems with the iceberg assumption.

produce the differentiated good in different locations. On the demand side, the final demand in location j can be obtained via utility maximization of the corresponding CES utility function:

$$\max_{m_{i,j}(z)} D_j \tag{1}$$

where D_j represents the consumption of the differentiated good in location j. D is an aggregate of industrial varieties defined by a CES function a la Dixit and Stiglitz (1977):

$$D_j = \left[\sum_{i=1}^{R} \int_0^{n_i} m_{i,j}(z)^{\sigma-1/\sigma} dz \right]^{\sigma/\sigma-1} \tag{2}$$

where $m_{i,j}(z)$ is the consumption of the each available variety z in location j that is produced in location i and n_i is the number of varieties produced in location i. σ represents the elasticity of substitution among the varieties of the differentiated good where $\sigma > 1$. Products are homogeneous if σ tends to infinity and varieties are very differentiated if σ is close to one. Consumers maximize their utility (function 1) bearing in mind the following budget constraint:

$$\sum_{i=1}^{R} n_i x_{ij}^{D} p_{ij} = Y_j \tag{3}$$

The consumer's problem solution gives the final demand in location j for each variety produce in location i.

$$x_{ij}^{D} = p_{ij}^{-\sigma} \left[\sum_{n=1}^{R} n_n p_{nj}^{1-\sigma} \right]^{-1} Y_j \tag{4}$$

where p_{ij} ($p_{ij} = p_i T_{ij}$), is the price of varieties produced in location i and sold in j and Y_j represents the total income in location j.

If we define a price index for the differentiated goods[5] $P_j = \left[\sum_{n=1}^{R} n_n p_{nj}^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$ and rewrite the consumption expenditure as $E_j = Y_j$, final demand in location j can be written as $x_{ij}^{consD} = p_{ij}^{-\sigma} P_j^{\sigma-1} E_j$. However, in order for x_{ij}^{consD} units of consumption to arrive at location j,

[5] This Industrial Price Index in location *j* measures the minimum costs of purchasing a unit of the composed index of manufacturing goods *D* so it can be interpreted as an expenditure function.

$T_{i,j}x_{ij}^{consD}$ must be shipped. So the effective demand a firm in location *i* faces from a consumer in location *j* is given by:

$$x_{ij}^{D} = T_{ij}\,p_{ij}^{-\sigma}P_{j}^{\sigma-1}E_{j} = p_{i}^{-\sigma}T_{ij}^{1-\sigma}P_{j}^{\sigma-1}E_{j} \tag{5}$$

On the supply side a typical firm in location i maximizes the following profit function:

$$\Pi_{i} = \sum_{j=1}^{R}\frac{p_{ij}x_{ij}^{D}}{T_{i,j}} - w_{i}^{D}(F + cx_{i}^{D}) \tag{6}$$

Technology in the increasing returns to scale manufacturing sector is given by the usual linear cost function: $l_{Dij} = F + cx_{ij}^{D}$, where l_{Dij}, represents the industrial workers used for the production of a variety in location i and sold in location j, F, represents a fixed cost of production, c, is the variable unit cost and x_{ij}^{D} is the amount of the differentiated good demanded in location j and produced in location i ($x_{i}^{D} \equiv \sum_{j} x_{ij}^{D}$ represents the total amount of output produced by the firm in location i and sold in the different j locations) and w_{i}^{D} is the nominal wage paid to the manufacturing workers in location i. The assumptions of increasing returns to scale, preference for variety by consumers, and the existence of an infinite number of varieties of the differentiated good means that each variety is going to be produced by a single specialized firm in only one location. In this way, the number of the manufacturing firms is exactly the same as the number of available varieties. Each firm maximizes its profit behaving as a monopoly of its own variety of the differentiated good. First order conditions for profit maximization yield the standard result that prices are set as a constant mark-up over marginal costs.

$$p_{i} = \frac{\sigma}{\sigma-1}w_{i}^{D}c \tag{7}$$

where $\frac{\sigma}{\sigma-1}$ represents the Marshall-Lerner price-cost mark-up. The higher this ratio, the higher the degree of monopoly power by a firm. As a result, Krugman (1991b) interprets σ as an inverse measure of scale economies since it can be thought as a direct measure of price distortion and as an indirect measure of market distortion due to monopolistic power. Given that

$\frac{\sigma}{\sigma - 1}$ is greater than one, Krugman (1991b) interprets this result as a way of justifying the existence of increasing returns to scale. If this pricing rule is substituted into the profit function, the following expression for the equilibrium profit function can be obtained:

$$\Pi_i = \left(w_i^D\right)\left[\frac{cx_i^D}{\sigma - 1} - F\right] \tag{8}$$

Free of entry assures that in the long run firms break even. So, the incentives for a firm to relocate in a different location have vanished. This implies that the equilibrium output is the following:

$$x_i^D = \bar{x} = \frac{F(\sigma - 1)}{c} \tag{9}$$

The price that is needed to sell this amount of output is $P_i^\sigma = \frac{1}{\bar{x}} \sum_{j=1}^{R} E_j P_j^{\sigma-1} T_{i,j}^{1-\sigma}$. This expression is combined with the fact that in equilibrium prices are a constant mark-up over marginal costs, the following zero-profit condition can be obtained:

$$w_i^D = \left(\frac{\sigma - 1}{\sigma c}\right)\left[\frac{1}{\bar{x}} \sum_{j=1}^{R} E_j P_j^{\sigma-1} T_{i,j}^{1-\sigma}\right]^{1/\sigma} \tag{10}$$

This equation is the so-called *nominal wage equation* in the literature of New Economic Geography, and constitutes the key relationship that is going to be empirically tested. Equation (10) shows that the nominal wage level at location *i* depends on a weighted sum of the purchasing power of the surrounding locations where the weighted scheme is a distance function that decreases as the distance between *i* and *j* increases. In the New Economic Geography Literature, the right hand side of the expression (10) has different names; the most common are *market access* (see Redding and Venables, 2001, 2004) and *real market potential* (see Head and Mayer, 2004). Here, this expression will be referred as market access, and it will be denoted by MA. The meaning of this equation is that those firms in locations that have a good access to big markets (high market access) will tend to remunerate their local factors of production (workers) with better salaries due to their savings in transportation costs.

If we normalize output production choosing our units in such a way that $c = \frac{(\sigma - 1)}{\sigma}$, and we set the fixed input requirement as $F = \frac{1}{\sigma}$, and define market access in location i as $MA_i = \sum_{j=1}^{R} E_j G_j^{\sigma-1} T_{i,j}^{1-\sigma}$, we can rewrite the *nominal wage equation* as:

$$w_i^D = \left[MA_i \right]^{1/\sigma} \tag{11}$$

This simplification in the *nominal wage equation* is very similar to the Harris (1954) market potential function in the sense that the economic activity is higher in those regions that are closer to big markets. So, New Economic Geography gives the micro-foundations for the ad-hoc formulation of the Harris (1954) market potential formulation.

3. Data and results

In this section, in order to show empirical evidence of the resulting hypothesis raised in section 2, the nominal wage equation for the Spanish provinces in the year 2003 is estimated. The research strategy will be the following. As a starting point for the regional investigation, a basic relationship is estimated. As this basic regression could be merely informative, a conditioning scheme should be undertaken, the unconditional baseline regression is transformed to a conditional one, informing about the relevance of the market access variable under the inclusion of control variables that might be affecting Spanish wages.

3.1. Econometric specification: baseline regression

Taking logs in the expression (11), the estimation of the nominal wage equation is based on the estimation of the following expression:

$$\log(w_i) = \theta + \sigma^{-1} \log\left[MA_i \right] + \eta_i \tag{12}$$

where η_i represents the error term and the other variables were defined in the previous section. This equation relates nominal wages in location *i* with GDP in the surrounding locations weighted by distance and prices. In accordance with the theoretical predictions of the model, the higher the prices and GDP in the surrounding locations and the shorter the distance between the different locations, the higher will be the local wage. This specification captures the notion of a spatial wage structure and allows testing for a direct relationship between nominal wages in a particular location and its market access. This also constitutes an important condition to reveal agglomeration dynamics.

With respect to the data for the empirical illustration, we proxy the “wage” variable of the model by the provincial per capita GDP expressed in Euros for the corresponding year of analysis, 2003. With respect to the “market access” variable it is computed as a weighted sum of the GDP of the surrounding provinces, where the weighted scheme is the distance measured in km between the capital cities of each province. The internal distance for each province is computed as proportional to the square root of the provinces’ area. The expression used to compute it is $0.66\sqrt{\frac{Area}{\pi}}$, where “Area” represents the size of the province in km^2. This expression generates the average distance between two points in a circular location (see Head and Mayer, 2000, Nitsch 2000 and Crozet 2004 for a discussion of this internal distance). Data on per capita GDP and GDP for the Spanish peninsular provinces in 2003 is taken from the Spanish National Statistical Institute (INE).[6]

<<insert figure 4 here>>

<<insert table 2 here>>

Presenting an exploratory data analysis, figure 4 provides a clearer view of the relationship between provincial per capita GDP and Market Access in year 2003 by means a simple graph. From figure 4, it seems that there is a positive relationship between the provincial per capita GDP for year 2003 and the provincial market access. Table 2 summarizes the results of the estimation of equation (12) for the sample of 47 Spanish provinces for the year 2003. Ordinary Least Squares (OLS) are used in this baseline estimation (Column 1 of table 2).

From table 2, itg can be seen that there is a significant relationship between both variables. On average, if market access increases by 1%, per capita GDP will increase by 1.26%. As a consequence, market access would be a relevant variable in explaining the wage structure in this Spanish system. Nevertheless, a problem we face with our baseline regression is that the

[6] It is necessary to emphasize that the availability of the instrumental variables that are used in the estimation is limited to year 2003. This is the reason why year 2003 is taken as the year of analysis.

independent variable, market access, is endogenous and simultaneously determined with GDP. This could lead to the well-known simultaneity bias in the regressions violating the necessary conditions to obtain estimates with good properties. The standard approach to overcome the consequences of simultaneity (biasness, inefficiency and inconsistency on OLS-estimators) is the instrumental variables (IV) estimation. IV estimation is based on the existence of a set of instruments that are strongly correlated to the original endogenous variables but asymptotically uncorrelated with the error term. Once these instruments are identified, they are used to construct a proxy for the explanatory endogenous variables which consists of their predicted values in a regression on both the instruments and the exogenous variables. However, it is difficult to find such instruments because most socioeconomic variables are endogenous as well. In this work, accessibility variables have been suggest as instruments, since they are highly correlated with the market access variable, but also noncontemporareously correlated with the errors. Consequently, three variables are proposed as instruments. A gravity indicator of efficiency and a location indicator, which are taken from Monzón *et al*. (2005)[7], and a third one, the mean time of access to a commercial airport, which is based on our own elaboration. The gravity indicator of efficiency is an adimensional variable, and shows the role played by infrastructures in the territorial distribution of the levels of accessibilities. It is an indicator of relative accessibility, showing the quality of the infrastructures in the relationships between nodes within a territory. When the value of this indicator decreases, accessibility increases. The other accessibility indicator, location indicator, is measured in minutes, and it shows how infrastructures enables access to the places where population is concentrated. If the value of the location indicator decreases, accessibility increases. Both, the gravity indicator of efficiency and the location indicator are available at the provincial level. Finally, market access is instrumented with the variable mean time of access to a commercial airport. Data on this instrument are available only at the regional level. Thus it was assumed that mean time of access to a commercial airport is identical in all provinces within the same region.

IV estimation (see column 2 of table 2) again finds a positive and highly statistically significant effects of market access, although there is a strong correction of the coefficient, changing from 1.26 to 0.16. Now, it is necessary to take into account other variables that might be affecting

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[7]These authors base their computations on the works of Schürmann et al. (1997) and Geurs and Ritsema Van Eck (2001).

Spanish wages through the market access variable. From an econometric point of view, the inclusion of these variables may avoid problems derived from misspecification that could be biasing the coefficient of interest.

3.2. The conditioned econometric specification

As already noted, equation (12) is a restricted specification to analyze the effects of market access on nominal wages. The reason is that when running this bivariate regression it cannot be assured that the relationship is a causality relationship or simply capture correlations with omitted variables, such as infrastructure, human capital, innovation, etc. In order to deal with these issues and control for the existence of other shocks that might be affecting the dependent variable and are correlated with market access, an alternative specification that explicitly takes into account the aforementioned considerations is also estimated. The specification of the extended nominal wage equation takes the following form:

$$Lnw_{\mathrm{i}} = \theta + \sigma^{-1} \ln MA_i + \sum_{n=1}^{N} \gamma_n X_{i,n} + \eta_i \tag{13}$$

where X_{in} is a vector of control variables and γ_{in} the correspondent coefficient.

In the introductory section of this paper, it was proposed three channels through which market access might influence per capita GDP levels in the Spanish provinces. Besides the direct trade cost saving that accrues to central locations, stocks of medium and high educational levels are highly correlated with market access. The theoretical foundations for the relationship between market access and educational levels have been put forward by Redding and Schott (2003). They proved that high market access provides log-run incentives for human capital accumulation by increasing the premium of skilled labor. Empirical work carried out at international and European level have confirmed this relationship (see Lopez-Rodriguez *et al*. 2007; Redding and Schott, 2003). Research and Development expenditures are also affected by spatial proximity and geography. For instance, at the European level, the regional dimension is very relevant due to the presence of border effects. The interaction of high market access in dense and central European regions which makes them large and profitable markets for innovation, together with increasing returns to innovation and localization of the knowledge spillovers, seem to explain the pattern of high concentration[8] of innovative activities in the centre of Europe. Following Breinlich (2006), the stocks of productive capital is also incorporated as a control variable.

So taking into account that an important part of advantages of centrality works through accumulation incentives, a straightforward way of disentangling the importance of the direct trade cost advantage to central locations is by including the percentage of the workforce with secondary and/or tertiary education, (*lab*), the provincial per capita productive capital stock, *K*, and the per capita R&D expenditures, *R&D*, as additional regressors in the baseline specification estimated earlier.

Data on human capital and per capita productive capital stock were taken from the Instituto Valenciano de Investigaciones Económicas (IVIE) and refer to the year 2003. Data on per capita R&D expenditures were taken from the National Institute of Statistics of Spain (INE). While data on human capital, per capita productive capital stock and market access are available at provincial level, data on R&D expenditures are available only at regional level. Thus, it was assumed that per capita R&D expenditures are identical in all provinces within the same region. As consequence, the conditioning approach involves testing the following equation (14).

$$Ln\, w_i = \theta + \sigma^{-1} \ln MA_i + \gamma_1 \ln lab_i + \gamma_2 \ln K_i + \gamma_3 \ln R\&D_i + \eta_i \qquad (14)$$

In addition, another goal of this section is to shed further light on the analysis derived from equation (14) by broadening the empirical analysis by means of the consideration of the spatial dimension. In this sense, the geographic dimension of the dependent variable is explored by using an exploratory spatial data analysis (ESDA) approach. This analysis will help with the identification of the type of spatial pattern present in the per capita GDP provincial data. All computations were carried out by using SpaceStat 1.91 (Anselin, 2002), GeoDA (Anselin, 2003) and ArcView GIS 3.2 (ESRI, 1999) software packages. First, we test global spatial autocorrelation for the initial per capita income by using Moran's I statistic (Cliff and Ord, 1981), $I = \frac{N}{S_0}\frac{z'Wz}{z'z}$, where N is the number of provinces, $S_0 = \sum_i \sum_j w_{ij}$, z_{it} is the log of GDPpc in province i at time t=2003 in deviation from the mean, W was defined expressing for each province (row) those provinces (columns) that belong to its neighborhood. Formally, w_{ij}=1 if

[8] For comprehensive analysis of innovation activity in Europe see Bilbao-Osorio and Rodriguez-Pose (2004), Bottazzi and Peri (1999, 2003), Moreno *et al*. (2005) and Rodriguez-Pose (1999, 2001).

provinces i and j are neighbors, and w_{ij} =0 otherwise. This simple contiguity matrix ensures that interactions between provinces with common borders are considered.[9] For ease of economic interpretation, a row-standardized form of the W matrix was used. Thus, the spatial lags terms represent weighted averages of neighboring values.

The value of I for the 2003 per capita GDP was 0.721, well above the expected value for this statistic under the null hypothesis of no spatial correlation, E[I]=-0.021. It appears that the per capita GDP is spatially correlated since the statistic is strongly significant with p=0.001. This result reveals the existence of a strong and statistically significant degree of positive spatial dependence in the distribution of regional per capita GDP in 2003. Figure 5 shows the spatial distribution of regional per capita GDP in 2003. Figure 6 provides a clearer view of the spatial autocorrelation in this year through the Moran scatterplot.[10] Figure 6 shows a strong geographic pattern and reveal the presence of positive spatial dependence.

<<insert figures 5, 6 here>>

On the other hand, being aware of the potential drawback coming from the simultaneity problem due to the fact that the market access variable is endogenous and simultaneously determined with per capita GDP, the instrumental variables estimation will be used. Again, we use the same set of instruments: the gravity indicator of efficiency, the location indicator and the mean time of access to a commercial airport. The goodness of the instruments is proved with the Sargan test, which contrasts the null hypothesis that a group of s instruments of q regressors are valid. This is a χ^2 test with (s–q) degress of freedom that rejects the null when at least one of the instruments is correlated with the error term (Sargan, 1964). In our case, the null hypothesis is not rejected at 5%, validating the us of the instruments. Table 3 shows the estimation results of equation (14) by IV for the Spanish provinces in the year 2003 (Model 1).

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[9] Other alternative definitions for the spatial weights matrix were considered. Specifically, defining their elements as the inverse of the distances, and considering the median of the great circle distance distribution, the lower quartile, the upper quartile and the maximum distance. These matrices generated results very similar to those presented in this paper.

[10] The Moran scatterplot displays the spatial lag *W log(GDPpc)* against *log(GDPpc)*, both standardized. The four quadrants of the scatterplot identify the four different types of local spatial association between a province and its neighbours (Anselin, 1996): quadrants I (*High* income-*High* spatial lag) and III (*Low* income -*Low* spatial lag) correspond to positive spatial autocorrelation while quadrants II (*Low* income -*High* spatial lag) and IV (*High* income -*Low* spatial lag) refer to negative spatial dependence.

<<insert table 3 here>>

From Model 1 in Table 3, all coefficients are highly significant, having the expected sign. Besides, no problems were revealed with respect to a lack of normality (residuals from this regression are normally distributed, since the Jarque-Bera test does not reject the null hypothesis of normality), and there is no evidence of the existence of heteroskedasticity (Breusch-Pagan and Koenker-Bassett test). Nevertheless, the value of the Morans' *I* for the residuals was 0.176, and the null hypothesis of no spatial correlation is rejected (P-value=0.012). As the Lagrangre Multiplier (error) test is not significant, and the Lagrange Multiplier (lag) test is significant, it becomes clear that there would be evidence for the adoption of a spatial lag model.

Thus the spatial lag model shown in equation (15) is considered and estimated (Model 2 of Table 3)

$$\ln w_{\mathrm{i}} = \theta + \rho W \ln\ w_{\mathrm{i}} + \sigma^{-1} \ln MA_i + \gamma_1 \ln\ lab_i + \gamma_2 \ln\ K_i + \gamma_3 \ln\ R\&D_i + \eta_i \tag{15}$$

Now, from results of the spatial lag model, the Likelihood Ratio test on the spatial autoregressive coefficient ρ rejects the null hypothesis, showing that the spatial dependence has been adequately dealt with by incorporating the spatial lag of log (GDP_{pc}). In addition, the value of the Morans' *I* for the residuals was 0.025, and the null hypothesis of no spatial correlation is not rejected (*P*-value=0.297). The Moran scatterplot (see figure 7) displays the spatial lag of the residuals from Model 2 against the residuals, both standardized. Now, as the Morans' *I* confirmed, there is no evidence of local spatial association between the residuals of a province and the residuals of its neighbours.

<<insert figure 7 here>>

Hence, the results derived from the final specification in Equation 15 (Model 2) capture the notion of a spatial wage structure, showing a direct relationship between nominal wages in a particular location and its market access, since the parameter is significant and positive. If market access increase by 1%, per capita GDP will increase by 0.08%, ceteris paribus. Thus, the estimates in table 3 support the theoretical predictions of the New Economic Geography literature. The control variables are individually significant, with the exception of the human capital variable.

A second remark should be made about the relevance of externalities across provinces, derived from the significance and magnitude of the estimated spatial autoregressive parameter $\left(\hat{\rho}=0.284\right)$. The spatial lag coefficient is, as expected, significant and positive. This coefficient measures the strength of the inter-provincial spillover effects (such as technological spillovers or factor mobility) and indicates that the per capita GDP of a province is related to those of its neighbor provinces after conditioning for the market access, the productive capital, the human capital and the R&D expenditures. From a spatial externalities perspective, it seems that provincial per capita GDP depends not only on the market access and its own conditioning variables but also on the per capita GDP of its neighboring provinces.

The political implications of such findings for the design of regional policy cannot be underestimated. Since our results point out that proximity to consumers, proxied by market access, is a key element to understand per capita GDP disparities among Spanish provinces and can be acting as a penalty for convergence, improving market access in the so called Spanish "periphery" would represent a major policy instrument to encourage economic and social cohesion.

4. Conclusions

In this paper, the relevance of market access on the Spanish wage disparities has been tested. GDP per capita is used as a proxy for differences in wages. Using a data set for the 47 peninsular Spanish provinces in the year 2003, strong evidence was found supporting the thesis that market access is a key variable for the explanation of per capita income disparities. Hence, the weight of the evidence suggests that proximity to consumers, proxied by market access, is indeed partly responsible for the wages experienced by the Spanish provinces. Moreover, three important channels through which market access might be affecting the wage disparities among Spanish provinces are also discovered. These are the levels of human capital, the stocks of productive capital and the expenditures on Research and Development.

There is also suggestive evidence that the increase in transport costs could cause low regional performance. Therefore, the analysis of the empirical results provides important elements to discuss the way in which market access affects per capita provincial income disparities.

With respect to the lessons for regional policy design that can be derived from the theoretical and empirical outcomes stated in this work, it is clear that the relevance and efficiency of the different policies can depend on the market access. The important aspect of this study is the important role played by second nature geography in explaining per capita GDP disparities.

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Table 1: Per Capita GDP by Province (in Euros)

Period	1995	2004	Period	1995	2004
La Coruña	10069	16569	Albacete	8634	14671
Lugo	9164	14801	Ciudad Real	9314	15961
Orense	8800		Cuenca	9082	15331
Pontevedra	9040	15885	Guadalajara	12199	16741
Asturias	10208	16994	Toledo	9593	15148
Cantabria	10786	19156	Badajoz	6817	12765
Álava	15553	27176	Cáceres	8312	13642
Guipúzcoa	14265	24973	Barcelona	13966	23276
Vizcaya	13210	23532	Gerona	14682	24094
Navarra	14614	24761	Lérida	13848	24317
La Rioja	13255	21370	Tarragona	15162	24486
Huesca	12157	20123	Alicante	10342	17472
Teruel	12325	19857	Castellón de la Plana	12820	20496
Zaragoza	12585	21366	Valencia	11093	18459
Madrid	15204	25818	Almería	9566	18565
Avila	9643	15750	Cadiz	8322	15232
Burgos	13445	22153	Córdoba	8264	13337
León	10153	16603	Granada	7964	13704
Palencia	11251	19213	Huelva	8916	15987
Salamanca	9715	16619	Jaén	8032	13180
Segovia	11832	19906	Málaga	8574	15726
Soria	12236	19792	Sevilla	8930	15539
Valladolid	12373	20427	Murcia	9506	16481
Zamora	8835	15376			
Av. pc GDP	10950	18449	% Area Core vs Total	25%	
Madrid/Av. Pc GDP	1,38	1,39	% Pop Core vs Total	48%	
Highest pc GDP	15553	27176	% GDP Core vs Total	58%	
Highest/Av. Pc GDP	1,42	1,47	Av. pc GDP core/Av. pc GDP periphery	1,39	1,39
Lowest pc GDP	6817	12765			

Source: National Institute of Statistics of Spain (INE) and authors' calculations based on INE

Table 2: Regression results for log (GDPpc) on log (Market Access) at the Spanish provincial level in the year 2003

Dependent Variable: log (GDPpc)		
Variable	(1)	(2)
	OLS	Instrumental Variables
Constant	-2.926	8.233***
	(4.665)	(0.400)
Log (Market Access)	1.265**	0.160***
	(0.478)	(0.042)
R^2	0.134	0.241
Prob (F-statistic)	0.011	0.0004
First stage R^2		0.784
Number of observations	47	47

Instruments: log (Gravitatory indicator of efficiency), log(Localization indicator) and log (Mean time of access to a commercial airport)
Standard errors in parenthesis.
** indicates coefficient significant at 0,05 level; *** significant at 0,01 level.
"First stage R^2" is the R^2 from regressing market access on the instrument set.

Table 3. Regression results for *log(GDPpc)* on *log(Market Access)* and control variables at the Spanish provincial level in year 2003

Dep.Variable: $\ln(GDPpc)$ Variable; parameter	Model 1 IV (p-value)	Model 2 ML-Spatial lag model (p-value)
Constant: (θ)	2.069 (0.002)	1.141 (0.118)
ln (MA); $\left(\sigma^{-1}\right)$	0.084 (0.001)	0.087 (0.000)
ln (*lab*): (γ_1)	0.152 (0.066)	0.114 (0.121)
ln (*K*) (γ_2)	0.556 (0.000)	0.403 (0.000)
ln (*R&D*): $((\gamma_3))$	0.061 (0.000)	0.046 (0.002)
Spatial lag of ln (GDP_{pc}): (ρ)		0.284 (0.020)
R^2	0.860	0.877
Akaike Inform. Crit. (AIC)	-96.114	-99.511
Schwarz Inform. Crit. (SC)	-86.863	-88.410
Jarque-Bera Normality test	0.865 (0.648)	
Heteroscedastic. Breusch-Pagan/ Koenker-Bassett test	1.468 (0.832) 2.193 (0.700)	1.489 (0.828)
Moran's I test (error)	0.176 (0.012)	0.025 (0.297)
Lagrange Multiplier (error)	3.202 (0.073)	
Spatial Dep. Robust LM (error) test	0.198 (0.655)	
Lagrange Multiplier (lag)	5.535 (0.018)	
Spatial Dep. Robust LM (lag) test	2.532 (0.111)	
Likelihood Ratio Test on spatial lag dependence		5.396 (0.020)

Note: The spatial weights matrix used in the calculations is a row-standardized form of the *W* matrix defined expressing for each province (row) those provinces (columns) that belong to its neighborhood. Formally, $w_{ij}=1$ if provinces *i* and *j* are neighbors, and $w_{ij}=0$ otherwise.

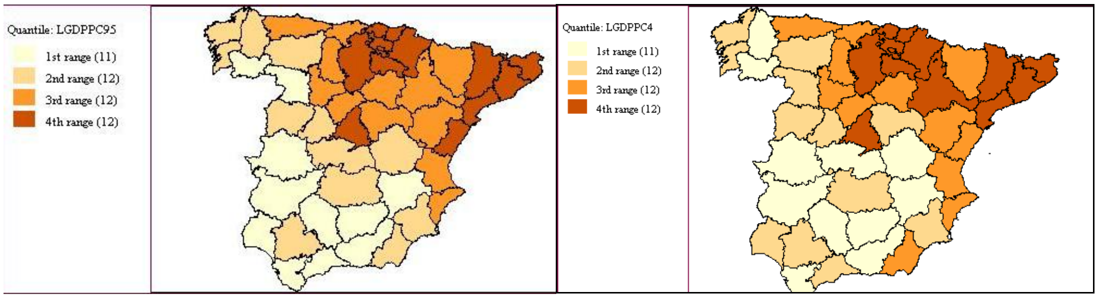


Figure 1: Provincial Per Capita GDP by Quantiles (years 1995 and 2004).

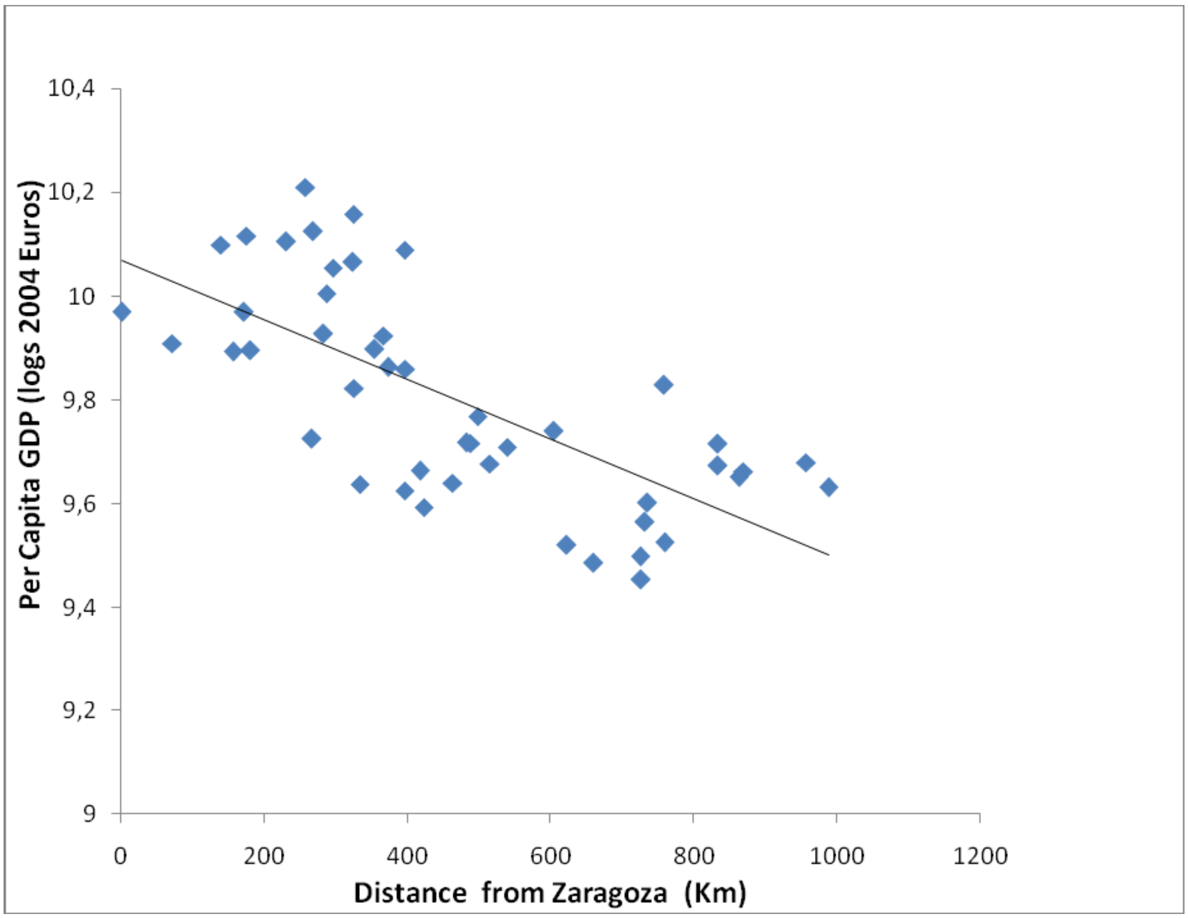
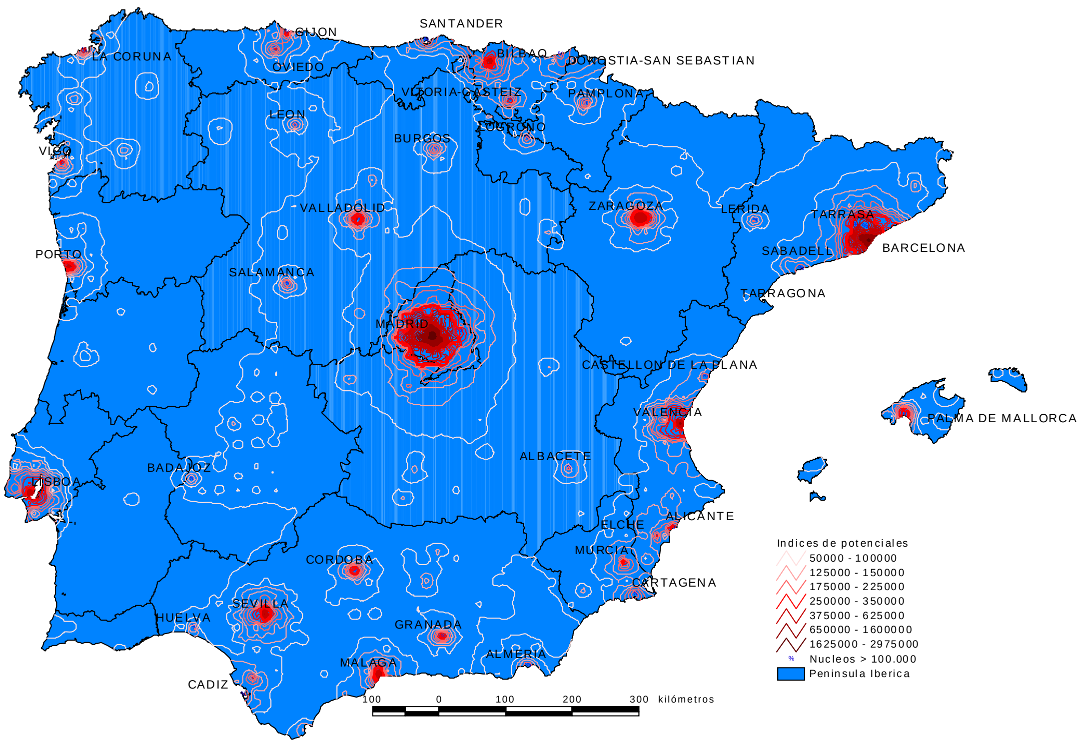


Figure 2: Per Capita GDP and Distance fron Zaragoza, 2004



Source: Faiña and Lopez-Rodriguez (2006)

Figure 3: Territorial Structure of Spain based on Market Potential Computations

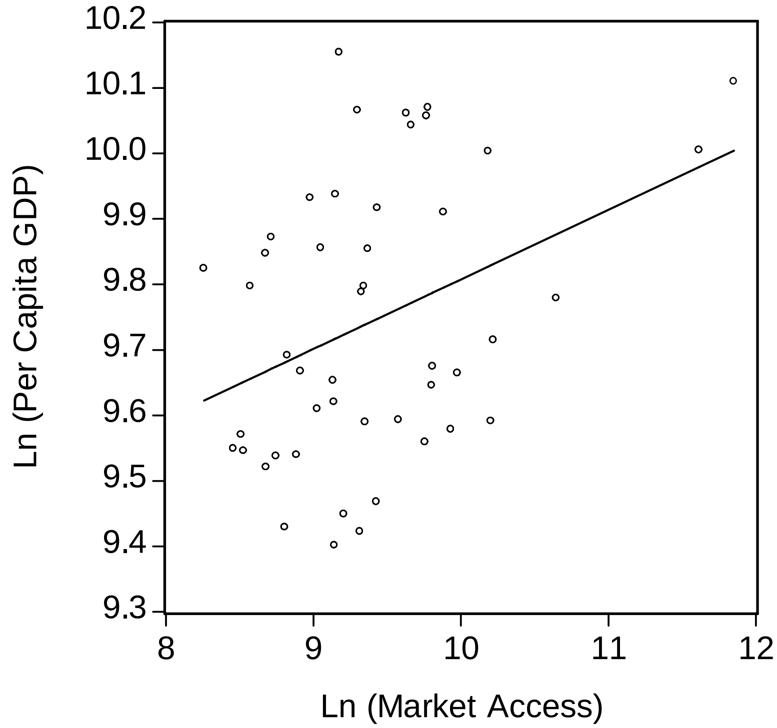


Figure 4: Provincial Per Capita GDP and Market Access , 2003

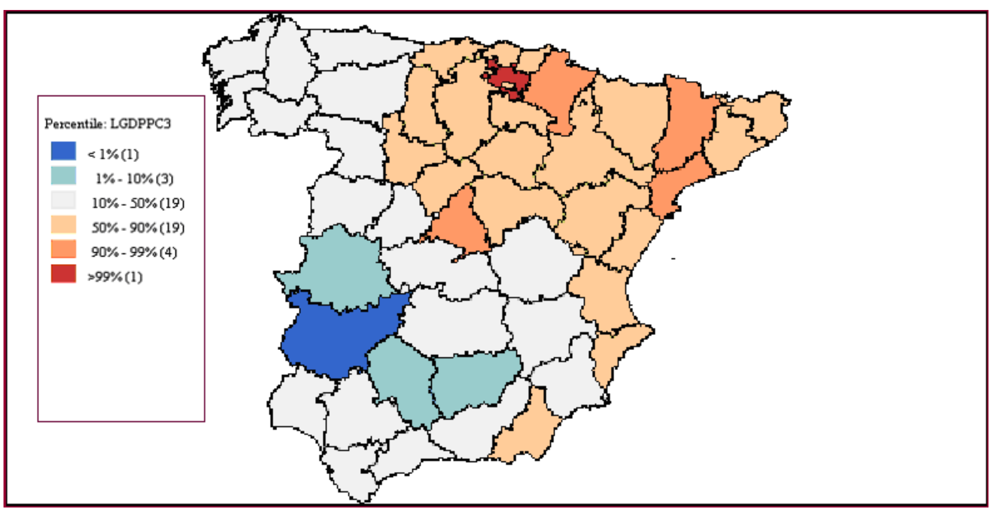


Figure 5: Spatial percentile distribution for the log of per capita GDP in 2003 (LGDPPC3).

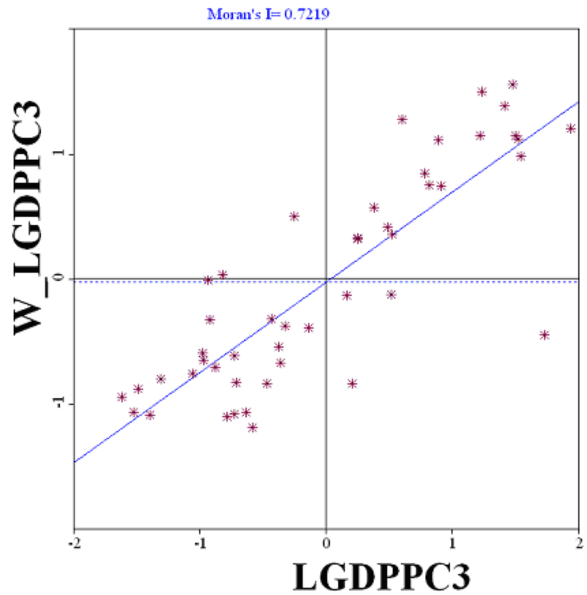


Figure 6: Moran scatterplot for the log of per capita GDP in 2003.

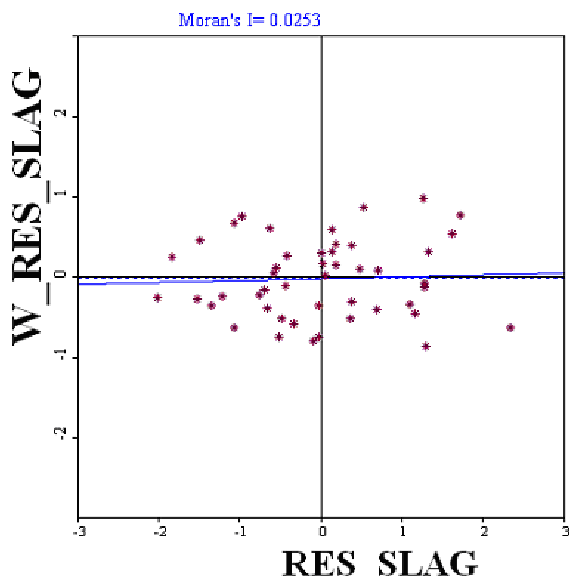


Figure 7: Moran scatterplot for the residuals from the spatial lag model showed in Model 2 of Table 3.