Urban growth and territorial dynamics in Spain (1985-2001): A spatial econometrics analysis*

José M. Mella-Márquez

jose.mella@uam.es

Coro Chasco-Yrigoyen

coro.chasco@uam.es

Universidad Autónoma de Madrid

ABSTRACT

The study of the territorial/regional development in Spain has nowadays a relatively long tradition, but from the point of view of cities development the number of studies and documents decreases drastically. This paper tries to improve the knowledge of the Spanish urban system. The aim of this paper is twofold: firstly, to determine the factors that explain the urban growth of Spanish cities; secondly, to observe the cities situation in terms of "winners" and "losers" after the long period of integration of Spain in the EU. A spatial conditional β-convergence equation is specified and the Durbin-Wu-Haussman exogeneity test is used to check on the existence of simultaneity between urban growth and the control variables. The classic problems of spread and backwash are studied by including a spatial autoregressive term and spatial regimes –convergence clubs- in the growth model.

Key words: urban growth, Spanish cities, conditional β-convergence, endogeneity, Durbin-Wu-Haussman, spatial autocorrelation, spatial heterogeneity, convergence clubs.

JEL classification: C21, C31, C51, D14, O18, R11

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* Research partially supported by the "Instituto de Estudios Fiscales" and "Universidad Autónoma de Madrid" (Spain). We would like to thank Julie Le Gallo for helpful discussions relating to certain issues considered in this paper. We also appreciate the collaboration of Esteban Sanromá and Raúl Ramos, who have provided us with some valuable data. Earlier versions of this paper were presented at the Regional Science Association International World Congress at the PE Technikon, Port Elisabeth (South Africa) in April, 2004, Regional Economic Applications Laboratory seminars at the University of Illinois at Urbana-Champaign in August 2004 and the First Seminar in Spatial Econometrics at the University of Zaragoza (Spain) in October 2004.

I. INTRODUCTION

The study of the territorial/regional development in Spain has nowadays a relatively long tradition, especially since the birth in the early eighties of the Autonomous Communities ("Comunidades Autónomas") or regions, considered as NUTS II in Eurostat nomenclature. There are plenty of articles and books written about the Spanish regional development, and in general they can be considered as a rich economic literature. But when one looks at the regional development topic from the point of view of cities, there are only a few documents and the studies are very scarce (Trullén, 2002; Trullén *et al.*, 2002; Viladecans, 2002; Mayor and López, 2003). This paper tries to improve the knowledge of the Spanish urban system and boost the urban studies in a country that has experienced a fast urbanization process within the last four decades. It is evident that at the international level this topic has received much more attention (in the EU, Cheshire, 2002; in the USA, Henderson, 1986, 1995; Glaeser *et al.*, 1992, Glaeser, 1998, among many other good references).

In fact, our intention is to test the same hypothesis formulated by Cheshire (2002): "The integration of Europe favours the core regions at the expense of the peripheral ones (…) removing protection as a result of economic integration works to the relative disadvantage of backward, peripheral regions and favours advanced core regions" (pp. 213). Several analysis have been presented in the literature (from Clark *et al.*, 1969 to Venables, 1996) that employing very different methodologies, converge on the 19th-century economic geographer's conclusion that "the best protection for a backward region is a bad road", interpreting "road" broadly as a shorthand for "costs of doing trade", including all transport costs and tariff and non-tariff barriers. On the other point, there are two exceptions in Steinle (1992) and Cheshire and Carbonaro (1996) that reach to the conclusion that the strongest gains from integration would continue, over time, to spread outwards from the core to the near periphery: for example, to Catalonia, Valencia, the Ebro valley and to the south and east of England.

In this paper, we analyse the performance of the 122 main Spanish cities (province capitals and those with more than 50,000 inhabitants) in terms of per capita GDP. The analysis period starts just before the Spanish adhesion to the EU in 1985 and ends in the year 2001, for which we have available data. As a result of this analysis we can also conclude that, in the group of main Spanish cities, there has been a significant spread from the core to the near periphery in terms of economic development during the EU integration period. In fact, if in 1985 the core was located in the northern and eastern cities, as well as in Madrid and some nearby metropolitan towns, by 2001 there has been a shift towards some near periphery cities, especially in the Castilian ones surrounding Madrid (Avila, Ciudad Real, Cuenca, Salamanca, Segovia, Valladolid, etc.).

Therefore the aim of this paper is twofold: firstly, to determine the factors that explain the urban growth of Spanish cities; secondly, to observe the situation of cities in terms of “winners” and “losers” after a long period of integration of Spain in the EU or at least how is the pattern of Spanish cities like after a period of Economic and Monetary Union -Single European Market and Euro- impact. For the former, we estimate a β-convergence model and spatial effects are tested with spatial econometric techniques. And for the later, we also use exploratory spatial data analysis (ESDA).

We use the formal tools of spatial econometrics to identify and include the relevant spatial effects in the estimation of the appropriate income growth model. In effect, in spatial cross-sectional contexts, it is almost inevitable to test for the presence of spatial spillovers. Spatial econometrics techniques, including the ESDA, are the appropriate tools to manage with spatial dependence in the error terms.

The paper proceeds as follows: section II introduces the theoretical framework to study the factors that explain the urban growth of Spanish cities and provides some insights into the β-convergence model and spatial effects upon which the empirical estimations described in the following sections rely. Section III presents the estimation of urban GDP and exploratory spatial data analysis (ESDA) is used to detect spatial autocorrelation and spatial heterogeneity among Spanish urban GDP data. In section IV, we present the model and set of control variables. The estimation, testing and re-specification process of the appropriate model is detailed in section V, in which these two spatial effects –spatial autocorrelation and spatial heterogeneity- are included. In section VI, some economic interpretations of the results are shown and the last section provides some concluding remarks.

II. <u>THEORETICAL APPROACH</u>

This section introduces the theoretical framework to study the factors that explain the urban growth of Spanish cities. The function that explains urban evolution in Spain could be a cross-sectional specification of the neo-classical growth model (Solow, 1956), which is considered as a natural starting point of the analysis of regional disparities, especially in

Europe (Fingleton, 2003). Although debatable, the neoclassical convergence specification have a sound empirical track record and that is why we considered it could also be useful to explain urban growth processes. The neo-classical model predicts that the growth rate of a city is positively related to the distance that separates it from its steady state. That is to say, if all urban economies are structurally identical and have access to the same technology, they are characterized by the same steady state, and differ only by their initial conditions:

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\alpha+\beta\ln\left(y_{i,0}\right)+\varepsilon_i \quad ; \quad \varepsilon_i \sim i.i.d.\left(0,\sigma_\varepsilon^2\right) \tag{1}$$

where $y_{i,t}$ represents per capita GDP in city i year t; α, β are parameters to be estimated and ε_i is a stochastic error term. This unconditional convergence model is very basic as it does not consider either the possible existence of spatial spillovers or other explicative variables for income growth or indeed the existence of spatial heterogeneity ("clubs") or even some endogeneity in the regressors.

It is interesting to analyse the economic meaning of the β parameter, which is computed as $\beta=\left(1-e^{-\theta t}\right)/t$, for θ the mean urban rate of convergence of the steady state. Urban units are said to exhibit absolute β-convergence if an estimation of (1) produces a significant negative estimate of β, i.e. poorer cities tend to grow faster than richer ones.

In fact, as shown in Rey and Montouri (1999), model (1) assumes that the error terms are independent, which is not very common in spatial cross-sectional contexts. That is why recent applications are increasingly testing for the presence of **spatial spillovers** in terms of spatial dependence in the errors[1]. Some of them have led to the conclusion of misspecification due to the presence of some kind of substantial spatial dependence, which is solved by the introduction of an endogenous variable spatial-lag (for EU regions, Vayá, 1997 and Ramajo *et al.*, 2003).

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\rho W\left[\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)\right]+\alpha+\beta\ln\left(y_{i,0}\right)+\varepsilon_i \quad ; \quad \varepsilon_i \sim i.i.d.\left(0,\sigma_\varepsilon^2\right) \tag{2}$$

where ρ is the scalar spatial autoregressive parameter. Spatial-lag model (2) focuses on the extent to which the growth rate of a city is related to the growth rate in its surrounding cities after conditioning on the starting year level of income[2].

Nevertheless most of the confirmatory analysis of this model has preferred the spatial error dependence specification: for US states (Rey and Montouri, 1999), Brazilian states (Magalhães *et al.*, 2000), EU regions (Moreno and Vayá, 2000; Le Gallo *et al.*, 2003), Spanish regions and provinces (Goicolea *et al.*, 1998; Toral, 2001).

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\alpha+\beta\ln\left(y_{i,0}\right)+\left[I-\lambda W\right]^{-1}u_i \quad ; \quad \mathrm{u}_i \sim i.i.d.\left(0,\sigma_u^2\right) \tag{3}$$

where λ is a scalar spatial error coefficient. This expression indicates that a random shock introduced into a specific city will not only affects its growth rate but, through the spatial multiplier [I - λW], will impact the growth rates of other cities, even if a given city has a limited number of neighbours. Given that OLS estimators, though unbiased, are not efficient, this model must be estimated by ML or General Moments Method.

The unconditional convergence model (1) could be tested for the hypothesis of **conditional convergence** to control for permanent cross-cities differences that could potentially explain urban growth income. The concept of conditional β-convergence is used when the assumption of similar steady states is relaxed and some explicative variables (different from $y_{i,0}$) that proxy the differences in steady state positions across different economies must be included in the model.

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\alpha+\beta\ln\left(y_{i,0}\right)+\gamma X_i+\varepsilon_i \quad ; \quad \varepsilon_i \sim i.i.d.\left(0,\sigma_\varepsilon^2\right) \tag{4}$$

where X_i is a vector of variables maintaining constant the stationary-state of city i. They can be referred to the beginning-of-sample period or can be the average over the sample period.

These "control variables" –as they are usually called (Barro and Sala-i-Martin, 1995; Vayá, 1996; Colino y Noguera, 2002; Johnson and Takeyama, 2003; Ramajo *et al.*, 2003)- can be **stock of physical capital** (e.g. ratio of public consumption/investment domestic to GDP, ratio of airports, banks, volumes in public libraries, telephone lines, hospital beds, institutions of higher education, electrical generation capacity, post offices, miles of railroad, registered motor vehicles), **stock of human capital** (e.g. fraction of population with a college degree, ratio of registered borrowers in public libraries), **factor productivity and technological progress variables** (fraction of total employees that are employed in farm/manufacturing/services, fraction of total employees that are employed in professional-technical occupations and in local-state-federal (civilians) government employees, occupation rate, ratio of farm/manufacturing/services GDP to total GDP, ratio of value of farm land, gross new firms as a fraction of existing firms, number of patent and designs issued) and **environment and state characteristics** (e.g. population density, median age of the population, fraction of population living in a family, birth/death/fertility/marriage rates, migration flows, rate of population living in the same house, rate of murders, degree of political instability, and fraction of population living in urban areas).

The conditional convergence model estimates two effects on economic growth. The first one is an expected negative effect of the initial per capita GDP through the estimated value of β in order to capture the convergence phenomenon. The second one corresponds to all other effects on growth of each explanatory variable introduced in X_i, i.e. which variables contribute or weaken economic growth and the way the convergence process is constrained by some explanatory variables.

Model (4) can also include some exogenous spatial lag variables –environmental ones- as explicative of income growth. This specification explicitly allows handling of spatial spillover effects (Le Gallo *et al.* 2003).

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\alpha+\beta\ln\left(y_{i,0}\right)+\gamma X_i+\xi WZ_i+\varepsilon_i \quad ; \quad \varepsilon_i \sim i.i.d.\left(0,\sigma_\varepsilon^2\right) \tag{5}$$

The influence of h spatially lagged exogenous variables contained in the ($N\times h$) matrix Z is reflected by the parameter vector ξ. A particular case of this last model occurs when Z includes only the $ln(y_{i,0})$ variable, which is exogenous and allows for an OLS estimation (Rey and Montouri, 1999).

The absolute convergence hypothesis, model (1), could also be rejected in favour of another kind of conditional specification or **club convergence hypothesis** (Durlauf and Johnson, 1995), which implies the existence of different urban economies (clubs) that are similar in structural characteristics and tend to converge within groups. The equilibrium each city will reach depends on the initial conditions of the group they belong to. In this case, income growth models include different parameters α, β, γ for each city club and X variables (called "split variables") should be referred to the beginning period of the sample (Ramajo *et al.*, 2003, López-Bazo *et al.* 2004).

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right)=\alpha^{g}+\beta^{g}\ln\left(y_{i,0}\right)+\gamma^{g}X_{i,0}+\varepsilon_{i} \quad ; \quad \varepsilon_{i} \sim i.i.d.\left(0,\sigma_{\varepsilon}^{2}\right) \tag{6}$$

where *g* represents the different clubs. A particular case of this model occurs when only the $ln(y_{i,0})$ variable is considered in the right-hand side of the expression (Baumont *et al.*, 2003).

The composition of the city clubs depends on the distribution of the split variable (e.g. sorting the sample of this variable into ascending order based on that variable and by dividing the sorted sample into thirds) but could be defined after an exploratory spatial data analysis (ESDA) as different spatial regimes, as in model (6). Nevertheless, there are not enough conceptual reasons to make an appropriate choice of the control/split variables as well to select either a conditional or club convergence model. And if it were, data on some of these variables may not be easily accessible or reliable.

In addition, it is also possible to find **endogeneity problems** due to system feedbacks in income growth models. Endogeneity is quite common in empirical regional and urban economic research but has not been received much attention in the particular context of income growth specifications (see some exceptions in Anselin *et al.*, 2000; Aronson *et al.* 2001, Henry *et al.* 2001, Fingleton *et al.* 2003, Stirboeck 2003). If this were the case, some endogenous explicative variables would also be correlated with the error terms invalidating the OLS estimates (and associated statistical inference), which would no longer be consistent similarly to what happens in systems of simultaneous equations. In this case, the estimation must be carried out by means of Instrumental Variables (IV) procedures, such as Two Stage Least Squares (TSLS), as shown in Anselin and Kelejian (1997).

We can assess this topic with the help of the **Durbin-Wu-Haussman** (DWH) test for the consistency of the OLS estimates when any endogeneity is present in a model; it has also been proposed as an “exogeneity test” (Anselin, 1999). In fact, it is an F test with ($k^{*}, n - k - k^{*}$) degrees of freedom on the null hypothesis of exogeneity of a k^{*} subset of the total k explicative variables, for n as the number of observations (for technical issues, see Davidson and McKinnon, 1993).

Anselin (1992) states that the principle of instrumental variables estimation is based on the existence of a set of instruments that are strongly correlated with 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 explicative endogenous variables that consists of their predicted values in a regression on both the instruments and the exogenous variables. In the standard simultaneous equations framework, the instruments are the "excluded" exogenous variables.

A conditional income growth model (4) can contain two types of control variables: exogenous and endogenous ones:

$$\frac{1}{t}\ln\left(\frac{y_{i,t}}{y_{i,0}}\right) = \alpha + \beta \ln\left(y_{i,0}\right) + \gamma X_i + \theta Y_i + \varepsilon_i \quad ; \quad \varepsilon_i \sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \tag{7}$$

where y_i is a vector of endogenous explicative variables and θ is its corresponding parameter vector.

In the next section, exploratory spatial data analysis (ESDA) is used to detect spatial autocorrelation and spatial heterogeneity among Spanish urban GDP data. These two spatial effects are then included in the estimation of the appropriate model.

III. EXPLORATORY SPATIAL DATA ANALYSIS

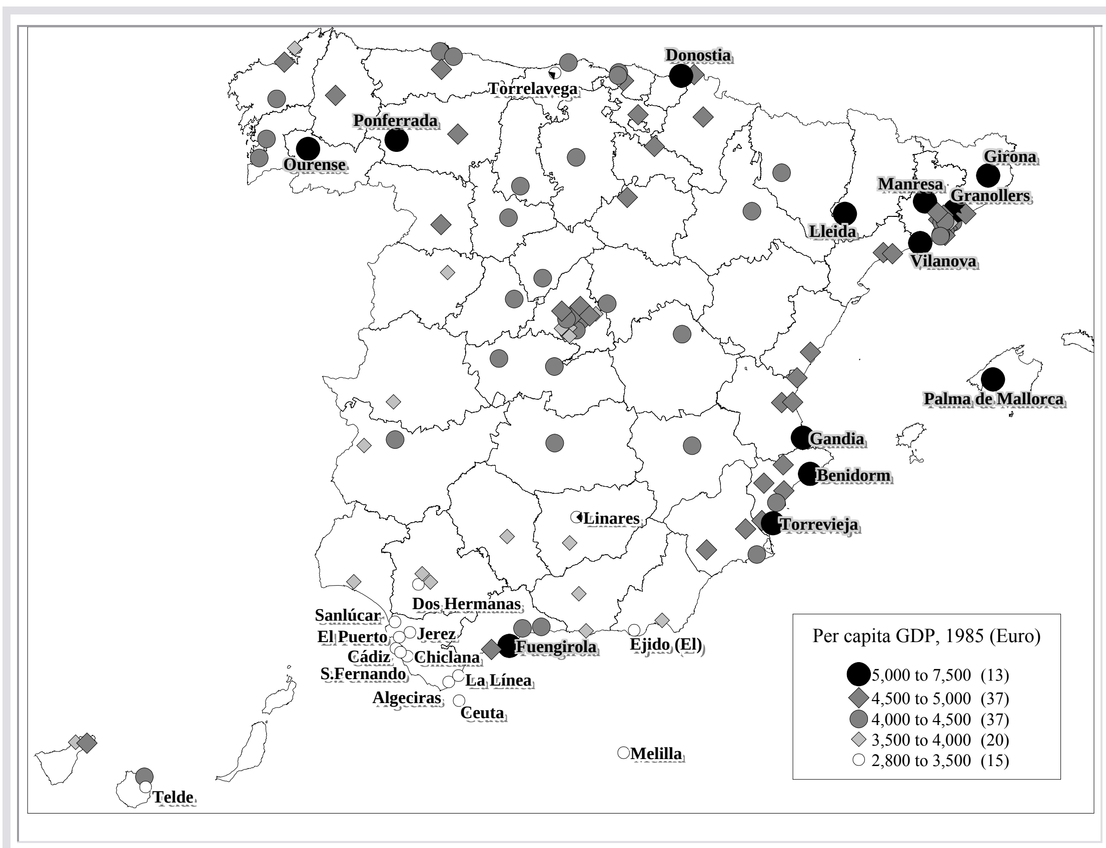
In this section we explore the geographic dimension of per capita GDP for the main Spanish cities: province capitals and those above 50,000 inhabitants in the 2001 Census (122 cities in total). As income data is not available for the local level, we first explain the estimation method we have followed to obtain these figures. Next, we test for global and local spatial autocorrelation for this variable in the initial and final period (1985 and 2001) using Moran's I statistic, Moran scatterplot and Local Moran Cluster Maps. Indeed, these tools are used to detect spatial trends and clusters, as well as some spatial discontinuities in the distribution of urban income. These spatial effects contribute to determine the factors that explain urban growth as well as to define spatial regimes, which are interpreted as spatial convergence clubs, in order to capture some kind of polarization pattern in the group of main Spanish cities (see an example of ESDA applied to regional convergence in Moreno and Vayá, 2000 and Le Gallo and Ertur, 2003).

II.1. Estimation of municipal GDP

As GDP data for all Spanish municipalities is neither available in the INE (Spanish Office for Statistics) nor in any other institution, we have had to estimate it using a spatial extrapolation model[3]. This is an ecological model of provincial per capita GDP, which is available in the INE only for current prices, over five per capita explanatory variables: telephone lines, car pool, fleet of lorries and vans, bank and savings bank offices and retail outlets (see Chasco, 2003 for more details on this kind of ecological inference process)[4]. Two regressions have been estimated for the 1985 and 2001 per capita GDP so that urban income growth can be calculated for this period[5].

We have estimated the per capita GDP for the cited sample of 122 main Spanish cities in the 17 Spanish regions from 1985 to 2001: Andalusia (23), Aragon (2), Asturias (3), the Balearics (1), the Canary Islands (4), Cantabria (2), Castile and Leon (10), Castile-La Mancha (6), Catalonia (21), Community of Valencia (12), Extremadura (3), Galicia (7), Madrid (14), Murcia (3), Navarre (1), the Bask Country (7) and La Rioja (1).

Figure 1 1985 urban per capita GDP



Source: Self-elaboration with MapInfo.

This estimation shows (Figure 1) that the highest per capita GDP in 1985 took place mainly in some of the Eastern Spanish cities, with an increasing size in terms of population, which it is the case of Fuengirola (Málaga), Gandia (Valencia), Benidorm (Alicante) and Vilanova i la Geltrú (Barcelona). On the contrary, other urban areas also had a highest income level but not a great population dynamism, which it is the case of some capital cities such as Donostia-San Sebastián, Girona and Lleida. It is well known that, as a consequence of congestion problems and increasing land/housing rents, some of the biggest Spanish cities are loosing strength in favour of neighbouring locations in terms of population, but not always in terms of per capita GDP. Nevertheless, there are some interesting exceptions to this rule constituted by three satellite towns, which have registered a lower –even negative- population growth but are also in the highest income group, as is the case of Manresa (Barcelona).

On the other end of the distribution, the lowest per capita income (Figure 1) is registered in southern up-growing population towns such as the tourist city of Chiclana de la Frontera, El Puerto de Santa María, Sanlúcar de Barrameda (Cádiz) and Telde (Las Palmas), as well as the metropolitan Dos Hermanas (Seville) and the town of El Ejido (Almería). There are also other poorer towns which have suffered from population losses, such as Cádiz and Torrelavega (Cantabria). Take note of the existence of some outliers such as Torrelavega (a low income point in the wealthier North) and Fuengirona (a high income point in the poorer South).

Similar results are found in the spatial distribution of per capita GDP in 2001, with a significant drift of higher per capita income from southeastern and northern towns towards the northeastern and central ones, respectively around the cities of Barcelona and Madrid. This is similar for Spanish provinces (as shown in FBBVA, 2002): although all the main cities have significantly improved in terms of per capita income and labour productivity, some of them have not taken advantage of this development to increase their population whereas others have almost duplicated their size.

II.2. Analysis of global spatial autocorrelation

Some kind of spatial trend can therefore be found in urban per capita income, as shown in Figure 1: from the Southwest (low income) to the East -and some northern cities- of the Peninsula (high income). This is a spatial effect called "spatial autocorrelation", which can be defined as the coincidence of value similarity with locational similarity (Anselin, 2000). There is positive spatial autocorrelation when high or low values of a random variable tend to cluster in space and there is negative spatial autocorrelation when geographical areas tend to be surrounded by neighbours with very dissimilar values. The measurement of global spatial autocorrelation is based on the Moran's *I* statistic, which is the most widely known measure of spatial clustering (Cliff and Ord, 1973, 1981). This statistic is written as follows:

$$I = \frac{n}{S_0} \frac{\sum_{i=1}^{n} \sum_{j=1}^{n} w_{ij} \left(y_i - \overline{y} \right) \left(y_j - \overline{y} \right)}{\sum_{i=1}^{n} \left(y_i - \overline{y} \right)^2} \tag{8}$$

where y_i: natural log of per capita GDP in city i.

$\overline{y}$: average value of variable "y".

w_{ij}: is an element of a spatial weights matrix, W, such that each element w_{ij} is set equal to 1 if city *j* is 130 km far from city *i*, according to the steepen slopes of the accessibility function (as shown in next chapter). Similar results have been obtained with other specifications for 150 and 220 km[6].

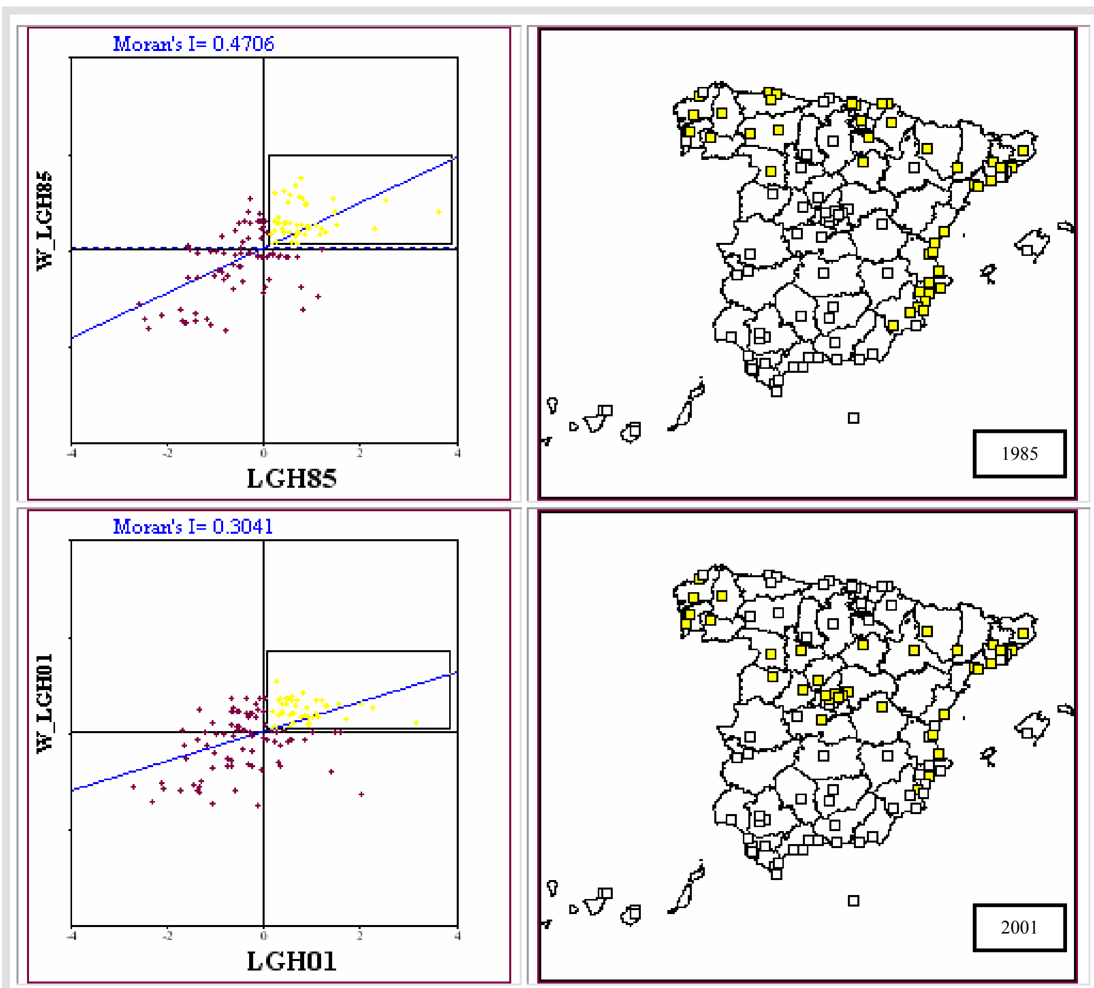
S_0: scaling factor equal to the sum of all elements of W.

In the given period, the urban income distribution displays a high degree of spatial autocorrelation: the magnitude of Moran's I test is high (I=0.4706 in 1985 and I=0.3041 in 2001) and strongly significant at p=0.0000, which is in both cases well above than its expected value under the null hypothesis of no spatial autocorrelation, E[I]= –0.008, in both cases. Inference is based on the permutation approach (999 permutations) though similar results have been obtained with the normalization and the randomization approaches (Anselin, 1995A, B). This result suggests that the evolution of urban income distribution appears to be somehow clustered in nature. That is, cities with very relatively high/low income levels tend to be located nearby other cities with high/low income levels more often than would be expected due to random chance. If this is the case, then each city should not be viewed as an independent observation. Similar results have also been found for the change of per capita GDP from 1985 to 2001; in this case, I=0.2968, with p=0.0000.

Figure 2 provides a more disaggregated view of the nature of spatial autocorrelation for per capita income through a Moran scatterplot, suggested by Anselin (1996), which plots the standardized income of a city (LGH85 for 1985 and LGH01 for 2001), against its spatial lag (also standardized), W_LGH85 and W_LGH01, respectively. A city's spatial lag is a weighted average of the incomes of its neighbouring states, with the weights being obtained from a row-standardized spatial weight matrix (W). The four different quadrants of the scatterplot identify four types of local spatial association between a city and its neighbours: HH ("High-High"), LL ("Low-Low"), LH ("Low-High") and HL ("High-Low").

In the first quadrant, HH, the Moran scatterplot represents the high-income cities that are surrounded by high-income neighbours: they have been highlighted in the map to the right of the Figure. It can be appreciated that they are all mainly located in the North and East side of Spain, though during the 1985-2001 period there was a shift from both sides towards the central cities. In quadrant 3, LL, we can find the group of low-income cities, which are surrounded by low-income neighbours. In quadrants II (LH) and IV (HL), we can find the group of low/high income cities surrounded by high/low income neighbours. Quadrants III and I belong to positive forms of spatial dependence while the remaining two represent negative spatial dependence.

Figure 2 Moran scatterplots (left) and maps (right) of urban per capita GDP



Source: Self-elaboration with GeoDa (Anselin, 2003).

II.3. Analysis of local spatial autocorrelation

Moran's I statistic, which is a global measure of spatial dependence, does not detect the presence of non-stationarity pockets ("hot-spots") which deviate from the overall pattern. On its side, Moran scatterplot not only shows a visual impression on the overall stability of the global pattern of dependence, but also a more disaggregated view of spatial dependence. Nevertheless, it does not provide any evidence on the statistical significance of the HH, HL, LH and LL links between one observations and its neighbour. This is the purpose of the Local Moran statistic I_i (Anselin, 1995B), which for each observation *i* gives an indication of significant spatial clustering of similar values around that observation. A positive value for I_i indicates clustering of similar values (high or low), whereas a negative value indicates clustering of dissimilar values. It takes the following form:

$$I_i = \frac{z_i}{m_2} \sum_{j=1}^{n} w_{ij} z_j \qquad (9)$$

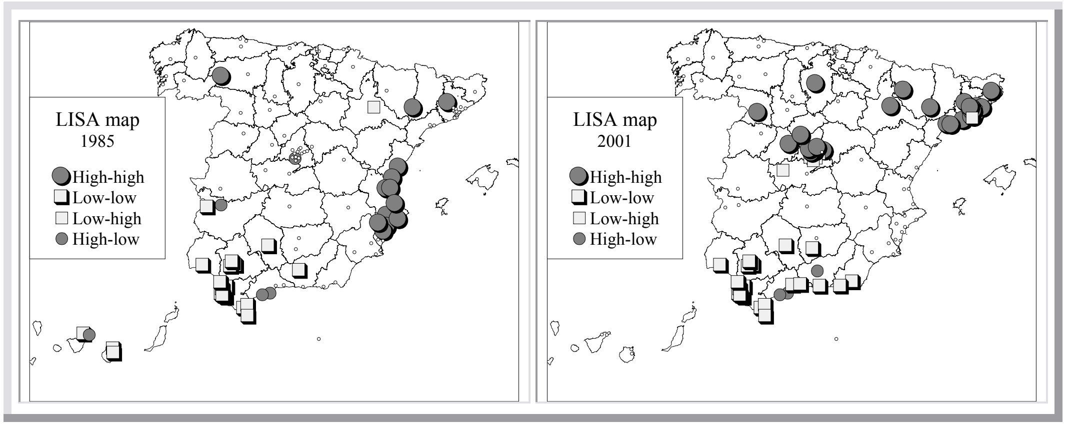
with $m_2 = \sum_{i=1}^{n} y_i^2$

z_j: natural log of per capita income growth in city i (measured as a deviation from the mean value)

w_{ij}: is an element of a spatial weights matrix, W, such that each element w_{ij} is set equal to 1 if city *j* is 130 km far from city *i*.

In Figure 3, a LISA Cluster Map is represented: it is a special choropleth map showing the locations associated to a significant Local Moran statistic classified by type of spatial correlation: big dark grey circle for the high-high association, big light grey square for low-low, small light grey square for low-high, and small dark grey circle for high-low. The high-high and low-low locations suggest clustering of similar values, whereas the high-low and low-high locations indicate spatial outliers.

Figure 3 LISA Cluster Map for p.c. GDP in 1985 (left) and 2001 (right)

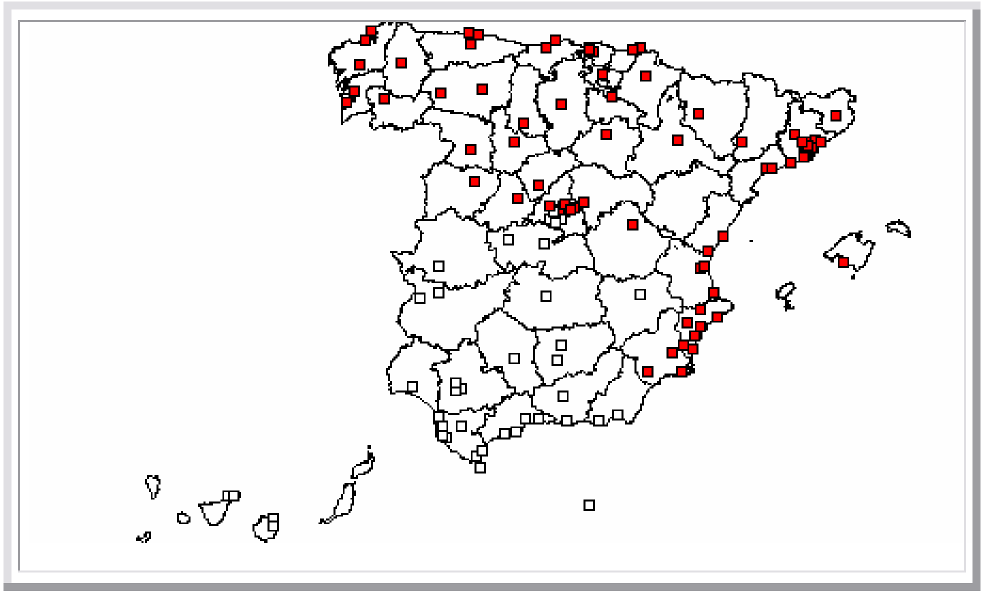


Source: Self-elaboration with MapInfo.

Both Moran scatterplot and LISA Cluster Map reveal the presence of two spatial clusters of urban income in Spain: a long arc which connects the North, Ebro Valley and eastern border cities, with the spatial discontinuity of Madrid and its metropolitan area (higher income) and the southern cities (lower income). Figure 3 shows that the high-high association, which was clearly located in the eastern coastal towns in 1985, has experienced a shift from the East to the central metropolitan towns around Madrid. Therefore, at present there are two poles of high-income concentration values in the trade area of Madrid and Barcelona.

Exploratory Spatial Data Analysis (ESDA) has shown the existence of some kind of spatial polarization of urban economies in Spain, which is expressed as two different spatial regimes in per capita GDP distribution of Spanish cities, based on the locations included –or nearby- quadrant I of positive spatial autocorrelation of Moran scatterplot of per capita GDP in 1985 (Figure 4). This spatial structure almost coincides with the spatial distribution of per capita income in the Spanish provinces.

Figure 4 Spatial regimes: red (higher per capita GDP)/white (lower per capita GDP)



Source: Self-elaboration with GeoDa (Anselin, 2003).

On the one hand, there is one regime of lower per capita GDP in the south: from the southern metropolitan towns of Madrid to the Canary Islands. On the other hand, higher per capita income towns are located in the northern cities of Galicia, Catabrian Coast, Castille and Leon, Ebro Valley, Catalonia, Balearics, Community of Valencia and Murcia, as well as Madrid and its northeastern metropolitan cities.

This shape almost coincides with other recent analysis for Spanish provinces (FBBA, 2000; Garrido, 2002), in which two spatial clubs are determined in terms of per capita income: the more developed north-eastern-central provinces versus the poorer south-western ones. This has been a persistent reality in Spain for a long time.

IV. MODEL AND DATA

Our aim is to determine the factors that explain urban growth of Spanish cities. For this purpose, we will take into account the neoclassical growth theory, the new growth theory and the spatial context of urban growth in Spain.

The neoclassical theory states that growth is a function of its **initial conditions** when all the economies are structurally similar and characterized by the same steady state (Barro and Sala-i-Martin, 1990), though the orthodox neoclassical theory focuses on two types of laws that govern the process of economic growth, **capital** and **labour** (Solow 1956). The classical economic development specifically points out that a **shift of labour and capital** from less productive to more productive sectors (e.g. from subsistence agriculture to industry and service) can accelerate growth. Moreover, the new growth theory emphasizes the positive external effects of **knowledge** in production (Lucas 1988). Finally, we are also interested in analysing the **spatial context** of urban growth in Spain, searching for any locational effect of the Spanish economy gravity centre.

In the case of Spanish larger cities, as their initial conditions are not similar, we initially propose a **conditional convergence model**, the same as (4), in which income per capita growth in period (*0,T*) is a function not only of per capita income in time *0* but also of some control variables (*X*) that proxy the differences in steady state positions across different economies. Following this, as demonstrated in the ESDA, we test for the presence of spatial effects (S_i) in the form of spatial autocorrelation –spatial spillovers- and/or spatial heterogeneity -spatial clubs convergence hypothesis.

$$\begin{aligned} g_{i,T} &= f\left(g_{i,0}, X_i, S_i, \varepsilon_{i,T}\right) \\ \varepsilon_i &\sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \end{aligned} \tag{10}$$

To test for this hypothesis, 5 basic explicative variables have been selected in order to explain per capita GDP growth during the period 1985-2001 for the 122 selected Spanish cities: per capita GDP in 1985, capital, human resources, productive structural change, technology, location and economic potential change.

1) **Initial conditions:** As shown in the table Annex 1, *G85* is GDP per capita in 1985 (in natural logarithms), which is the result of the already shown estimation process. Following the neoclassical paradigm, we expect that less developed cities are catching-up with the richer cities, so that urban levels of per capita income do tend to converge over the long run because of diminishing returns to capital. In a competitive environment, regional labour and capital mobility as well as regional trade will also work in favour of factor price convergence, reinforcing the negative relation between growth and regional inequality. However, other schools of thought tend to agree with Myrdal's basic claim (1957) that growth is a spatially cumulative process, which is likely to increase inequalities (divergence). Therefore, a negative value for the slope coefficient "β" indicates convergence of per capita GDP across territorial units of analysis, within a given time period, while a positive value indicates divergence. In the group of main Spanish cities, a negative sign for β is expected, as most of the cities with less per capita income growth in the 1985-2001 period were the ones with higher

per capita GDP in 1985. As is shown in Figure 1, these cities are located in the North, Catalonia and the Mediterranean coast.

2) **Capital:** it is a dummy variable (*K*) measuring per capita capital growth by provinces in the 1985-1998 period; it takes the value 1 for cities located in those Spanish regions that have registered an above average income growth and 0 for the rest. This variable, which has been published in FBBVA (2002), is the sum of both private (residential and productive) and public capital (transport infrastructure, urban equipments). The Spanish economy was strongly capitalized with growth rates of 4.25 (above the average) from 1964 to 2000. In general terms, the capital accumulation process has been more intense than the own population growth. This accumulation process has had two main sources: the residential sector (a 48% of the private investment has been concentrated in housing, hotels and apartments) and the public transport infrastructure (the 38% of the total public investment). We know that the improvement derived from this capital investment has been unequal by the Autonomous Communities and provinces. A positive sign is expected in this variable, as cities located in highly capitalized provinces should benefit from this general tendency, and vice versa. In effect, the highest growth rates have been located in Madrid, Castile-La Mancha, Andalusia, some inland Northern regions (Navarre, La Rioja), the Archipelagos and, in general, the Mediterranean coastal territories[7].

   We have also selected another capital indicator, *B*, which is the growth rate in the number of banks and savings banks per capita registered in a city during the 1991-2001 period that should positively be related to the per capita income growth. This variable has been obtained from Banesto (1992) and "la Caixa" (2004). In this case, the higher rates are concentrated in Madrid and its surrounding area, whereas the lower ones are in some northern and eastern cities (Portugalete, Irún, Granollers, Torrent…)

3) **Human resources:** it is a quite complex conception made up of two elements: labour market and entrepreneurship. Glaeser and Maré (2001) found out that there is a direct relationship between human capital and productivity of city workers, as salaries of largest cities are higher than those of smaller ones. This is why human capital growth is expected to have a great explaining power on income growth. As for entrepreneurship, this can be defined as the availability of an entrepreneurial spirit which makes the introduction of the technological and productive innovations possible, i.e. a strategic factor for the transformation, adaptation and improvement of a city's competitivity

level. Otherwise not only the old sectors and activities can loose the possibility of being substituted, but the own city could miss the opportunities of raising new activities with more development potential (Vázquez, 1993).

As an indicator of human resources, we have introduced the percentage of population with a university degree in 2001, *D*, which is positively related to income growth as it expresses the final result of an accumulation process of human resources in a city. The 2001 Census (INE) provides this variable. Improvements within education not only imply human capital accumulation, but also augment labour technical change. In this way, this variable registers the highest scores in most central and northern cities of Spain, as well as in some south-easterly cities. Likewise, we have also considered the change in unemployment rate of cities in the 1991-2001 period, *U*, which is available from the Spanish Ministry of Labour and should be negatively related with income growth.

4) **Productive structure change:** it has played, at the provincial level, a main role in order to explain income growth and productivity convergence (Garrido, 2002). Therefore, we will use the non-agrarian occupation rate change in 1985-1999, *E*, as an expression not only of the agrarian weight loss during the last few decades in Spain, but also of the industry and service dynamism. This indicator has been obtained from Funcas (2001). We expect this variable to be positively related to per capita GDP growth, as cities located in non-agrarian-trend provinces should have been benefited from this tendency, and vice versa. The highest rates of non-agrarian employment growth are located in Madrid and some central provinces, Extremadura, the south-eastern Mediterranean provinces and the Islands.

   Finally, the tourism index[8] in 2001, *T*, has been considered as a way of taking into account the outstanding and positive role of tourism in the economic development of Spain. Not only the so-called "sun and beach tourism", which some cities of the Mediterranean coastline and the Islands specialize in (Benidorm, Marbella, Palma de Mallorca), but also the cultural tourism (Granada, Santiago de Compostela…) as well as business tourism, which cities- such as Madrid, Barcelona, Seville, Valencia, Bilbao and some others- are key players.

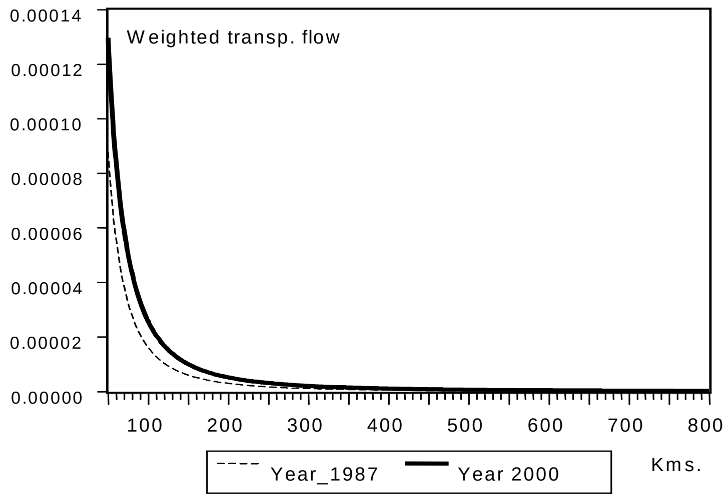
5) **Technology:** it is measured by the number of per capita patents by cities, *P*, and by the per capita R&D expenses by provinces, *R*, both referred to in 2000. The number of patents has been taken from the Spanish Patents Office files (Sáiz, 2003). The R&D

expenses data is only published at NUTS2 level but the INE has estimated it for the provincial level[9]. Audretsch (1998) points out that knowledge-based economic activity is generated and transmitted more efficiently via local proximity (actually via face-to-face interaction and through frequent and repeated contact) and has a high propensity to cluster within a geographic area: “intellectual breakthroughs must cross hallways and streets more easily than oceans and continents” (quoting Glaeser *et al.*, 1992), emphasizing that concerning the innovative output measured by the number of patents, “states aren’t really the right geographical units, the relevant geographic unit of observation is at the city level” (quoting Krugman, 1991).

The cities with higher concentration of patents per capita are located in the main productive centres, such as capitals and metropolitan towns around Madrid, Barcelona, Valencia and Seville. As for higher per capita R&D expenses, these are mainly located in the provinces of Álava, Barcelona, Burgos, Cáceres, Castellón/Castelló, Guipúzcoa, Lugo, Madrid, Pontevedra and Saragossa.

6) **Location and economic potential change:** The importance of the cities location in income growth can be approached using an accessibility index, which is an economic potential indicator (*EP*) of absolute location to the gravity centres of a country. The accessibility index is inspired in Clark *et al.* (1969) and Keeble *et al.* (1988), but estimated in a different way. Instead of an arbitrary choice of an exponent to the distance, we have estimated empirically the distance effects on the road transport of commodities for the particular Spanish case. We have calculated this index with data from the “Encuesta Permanente de Transporte de Mercancías por Carretera 2000” (Ministerio de Fomento) on regional exchanges of goods between the Spanish Autonomous Communities. After dividing the transport flow by the production of the origin and destination province, we have represented in a graph the weighted transport flow as a function of the distance and then we have adjusted the function taking into account several distances inside the same province (its radio, its surface and its third of the radio: the best one has been the first)[10]. The best exponent (measured by the R^2 measure of fit) was 2.5 for the year 1987 and 2.2 for the year 2000 (Figure 4).

Figure 5 Accessibility Index



Firstly, this shows that distance is becoming less reluctant to the economic transactions in Spain. Secondly, we can also conclude that for the same distance more and more trade is made per distance unit. And thirdly, these results show that trade is becoming more concentrated in shorter distances, because of the steepen slopes of this function for the year 2000, up to distances of approximately 130 km In other words, road transport improvements and diminution of transaction costs among EU regions in the last few decades have been prone to deepen the spatial disparities between the Spanish provinces.

"*EP*" is the accessibility index percentage change between 1987 and 2000 and measures the development of provinces towards the gravity centres of the Spanish economy. The provinces that have improved their economic accessibility are those which had a worse situation in the begining (1985) that is, those located in Andalusia (Almería, Cordoba, Granada, Jaén, Seville), Castile-La Mancha, Extremadura and Murcia.

"*X*" is a dummy variable capturing the most expansive Spanish provinces in terms of real per capita GDP in this period. It takes into account the fact that the gravity centre of the Spanish economy has been moving since the sixties to the Northeast (Funcas, 2002). This dummy takes the value 1 for all the provinces located on the east side of a diagonal line drawn from Navarre to Cadiz, crossing through Madrid (including the Canary Islands). They are regarded as being virtuous provinces as they have enjoyed better growth rates than the remaining ones. There are no theoretical reasons for including this variable, but empirical evidence shows that the economic activity is increasingly concentrating in the eastern side of the country.

V. EMPIRICAL RESULTS

The analysis is done over urban geographic scale which is defined as the group of Spanish cities that are province capitals and those with more than 50,000 inhabitants in the 2001 Census (in total, 122 cities or observations), as in ESDA.

Table 1 Pearson correlation coefficients of the model variables

	G	G85	B	D	P	K	U	E	T	R	EP	X
G	1	-0.59	0.42	0.30	0.08	-0.02	0.46	0.11	0.11	0.06	0.11	0.09
G85	-0.59	1	-0.24	0.10	0.22	-0.25	-0.41	-0.04	0.18	0.19	-0.06	0.14
B	0.42	-0.24	1	-0.01	0.03	0.17	0.27	0.39	0.07	0.08	0.32	0.22
D	0.30	0.10	-0.01	1	0.13	-0.15	0.32	-0.06	0.14	0.08	0.07	-0.04
P	0.08	0.22	0.03	0.13	1	-0.18	-0.11	0.00	0.23	0.08	0.04	0.10
K	-0.02	-0.25	0.17	-0.15	-0.18	1	-0.11	0.56	0.10	-0.30	0.37	0.19
U	0.46	-0.41	0.27	0.32	-0.11	-0.11	1	-0.02	-0.05	0.02	0.09	-0.06
E	0.11	-0.04	0.39	-0.06	0.00	0.56	-0.02	1	0.08	-0.25	0.63	0.43
T	0.11	0.18	0.07	0.14	0.23	0.10	-0.05	0.08	1	0.03	-0.02	0.08
R	0.06	0.19	0.08	0.08	0.08	-0.30	0.02	-0.25	0.03	1	-0.17	0.38
EP	0.11	-0.06	0.32	0.07	0.04	0.37	0.09	0.63	-0.02	-0.17	1	0.27
X	0.09	0.14	0.22	-0.04	0.10	0.19	-0.06	0.43	0.08	0.38	0.27	1

Urban growth is explained by the 11 already shown explicative variables. Nevertheless, in Pearsonian terms (Table 1), the relationship between some explicative variables is strong leading to multicollinearity problems in the model which produce large estimated variances in the coefficients (a first full model has been estimated producing a multicollinearity number of 46, well above the acceptable limit of 20/30 proposed by Anselin, 1995A). To avoid this situation we have previously tested the behaviour of the explicative variables and finally selected those with a higher explicative power, and a non-misleading correlation with urban growth, as well as less multicollinearity problems.

Therefore, as previously stated, our initial model is a conditional β-convergence model which relates urban per capita GDP growth with per capita GDP in the initial moment and the following control variables: banks per capita growth rate, percentage of population with university degrees and per capita number of patents:

$$\begin{aligned} &G_{i,T} = \alpha + \beta G85_{i,0} + \phi_1 B_i + \phi_2 D_i + \phi_3 P_i + \varepsilon_{i,T} \\ &\varepsilon_i \sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \end{aligned} \qquad (11)$$

where $G_{i,T} = G01_{i,T} / G85_{i,0}$ is the average growth rates of log per capita GDP in city i between date 0 and T (1985 and 2001, respectively); $G85_{i,0}$ represents the vector of log per capita GDP levels in city i in 1985; B is per capita banks and savings banks growth rate; D is the percentage of population with university degrees in 2001; P is per capita patents in 2000; α, β, ϕ_1, ϕ_2, ϕ_3 are parameters to be estimated and ε_i is a stochastic error terms with the usual properties.

In the case of conditional β-convergence models, there is conditional β-convergence if the estimate of β is significantly negative once X is held constant. These models estimates two effects on the transitional economic growth to the steady state. The first one is an expected negative effect of the initial per capita GDP through the estimated value of β in order to capture the convergence phenomenon. The second one corresponds to all other effects on the transitional growth of each explanatory variable introduced in X_i, that is, which variables contribute in weakening economic growth and how the convergence process is constrained by some explanatory variables.

We have estimated this model (11) by **OLS**. We have also carried out various tests to detect the presence of spatial dependence using the spatial weight matrices previously defined in (8). As it is shown in Annex 2, all the coefficients are quite significant, especially β= -0.0398, the negative sign of which indicates the existence of some kind of convergence trend in the set of main Spanish cities. Using the estimated β coefficient, the convergence process can be characterized by two additional concepts: on the one hand, the convergence speed, which can be defined as $b = -\ln(1+T\beta)/T$, and on the other hand, the half-life or the time necessary for the economies to fill half of the variation that separates them from their steady state: $\tau = -\ln(2)/\ln(1+\beta)$. In this estimation, the associated speed of convergence is 6.3%, significantly higher than the 2% usually found in the convergence literature, which indicates a quicker process (the half-life is 17 years). These results indicate that the convergence process is indeed stronger in urban contexts than in regional ones, probably due to the fact that there are more similarities between the economic conditions of big cities each other than the development which exists between regions[11].

As for the rest of the explicative variables (capital growth, human capital and patents), they are also strongly significant and positively related to urban per capita GDP growth. By standardizing the estimated coefficients, it is possible to know the relative influence of each explicative variable on the endogenous one, which is headed by per capita

GDP in 1985 (-0.59) and followed by human capital growth (0.34), capital growth (0.28) and knowledge of external effects (0.15).

In this case, the multicollinearity number is 26, which could be considered as acceptable. The Jarque-Bera non-normality statistic on the residuals takes on a relatively high value, which can be considered with certain precaution as being non-significant (p=0.08). Consequently, we will also accept with caution the results of the misspecification tests which depend on the normality assumption, such as the various Lagrange Multiplier (LM) tests. This is the case of the LM tests for **spatial autocorrelation** in the residuals, LM test for spatial error dependence and LM test for spatial lag dependence, which are highly significant in this model. Nevertheless, only the "robust LM test for spatial error", robust to the presence of spatial lag dependence in the model, is significant while its counterpart is not (p=0.0008)[12]. This result indicates that there is an evidence of the presence of spatial dependence in the OLS error terms (also stated by the Kelejian-Robinson test, which is not affected by non-normal errors), which would suggest a possible respecification of model (11) into a spatial error model.

However, it is also necessary to check for the presence of **heteroskedasticity** in the OLS error terms to evaluate the extent of which the significance in spatial autocorrelation tests may be influenced by potential heteroskedasticity, or vice versa. In our model, both the Breush-Pagan and White tests are quite significant (p=0.0378 and p=0.08) showing firstly the possibility of some heteroskedastic explicative variables in the model (it is performed against all them) and secondly, the existence of an unspecified form of heteroskedasticity possibly due, as it was shown in the ESDA, to the existence of two spatial regimes in the distribution of the endogenous variable, per capita GDP growth (Figure 4)[13].

Therefore OLS errors are not random and the sources of this non-randomness could on one hand, be the presence of some omitted variables which can be spatially correlated (spatial autocorrelation) and on the other hand, the existence of spatial heterogeneity due to one or more heteroskedastic explicative variables or to the existence of two spatial regimes, as suggested by both ESDA and the White test. It must be said that the links between spatial autocorrelation and spatial heterogeneity are strong so that they can be somehow equivalent (Anselin, 1995A).

That is why at this point, we could follow two strategies. First, we could estimate both a spatial lag (2) and spatial error model (3) to prove if the LM tests perform well in the absence of clear normality in the OLS errors; that is to check if the spatial error model is really a better specification than the spatial lag. Once decided about the best model, we should correct it from remaining spatial heterogeneity. Second, we could try to control first spatial heterogeneity, by the means of a spatial regimes model and afterwards to correct it of possible remaining spatial autocorrelation. Both lines lead to the same conclusion about the detected spatial effects in OLS residuals: They are a combination of two sources, not only spatial autocorrelation but also spatial heterogeneity. The main results of spatial error and spatial lag specifications (first strategy) are in the Annex 2. As it can be seen, the Wald test on common factor hypothesis clearly rejects the null, indicating an inherent inconsistency in this specification –the spatial error model is no longer appropriate- and the need of explicitly considering some kind of substantive spatial dependence (e.g. spatial lag model[14]). From now on, we follow the second strategy and show its main results in detail.

Therefore, will first test for the existence of structural instability in the form of two spatial clusters in order to capture the polarization pattern previously observed in the distribution of the per capita GDP growth among the 122 main Spanish cities: higher/lower income growth cities. Consequently, spatial heterogeneity can be controlled by allowing cross-region parameter variation in a **spatial regimes model** with two spatial regimes corresponding to higher-lower per capita GDP cities, as it was defined in Figure 4, and based on quadrants III and I of positive spatial autocorrelation of Moran scatterplot of per capita GDP in 1985:

$$\begin{aligned} G_{i,T} = {} & \alpha^H + \beta^H G85_{i,0} + \phi_1^H B_i + \phi_2^H D_i + \phi_3^H P_i + \\ & + \alpha^L + \beta^L G85_{i,0} + \phi_1^L B_i + \phi_2^L D_i + \phi_3^L P_i + \varepsilon_{i,T} \\ & \varepsilon_i \sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \end{aligned} \qquad (12)$$

where *H*, *L* represent high/low income regimes, respectively. We can interpret the parameter instability between the regimes (spatial heterogeneity) as indicative of cities belonging to two spatial convergence clubs: higher income growth cities, located on the Galician coast, Ebro Valley, East coast and surroundings of Madrid, versus lower income growth cities, which are mostly concentrated in the Cantabric North and South.

Model (12) has been estimated with OLS and we have computed a **spatial Chow test** proposed by Anselin (1990). This is a test on the null hypothesis where the coefficients are the same in all regimes and it is based on an asymptotic Wald statistic, distributed as a χ^2 distribution with *[(m-1).k]* degrees of freedom (*m* as the number of regimes). As it can be seen in Annex 2, the null hypothesis on the joint equality of coefficients is clearly rejected by the Chow-Wald test: its value (2.44) is sufficiently extreme for a distribution with 5 degrees of freedom. The same indication is provided by the test on the individual coefficients for *G85* and *B*. Regarding to *G85*, this model also estimates different negative coefficient values for each regime, highlighting the existence of two convergence speed for Spanish cities depending on the spatial regime in which they are located: cities located in the higher income area grow faster (12.1%) towards their steady state than the ones in the lower income growth area (5.5%). This result shows an interesting but worrying event: though both groups are experimenting a convergence process, the group of richer cities is growing faster than the group of poorer ones, so the former will reach its corresponding steady state first than the later ones.

In terms of fit, this model is much better than the other (AIC=-870.855)[15] but the evidence against of the normality of errors is stronger (Jarque-Bera=7.83) so the LM tests for heteroskedasticity and spatial dependence should be interpreted with caution, since they are based on the normal assumption. The Breush-Pagan test clearly accepts the homoskedasticity hypothesis (p=0.5337) possibly due to the explicit consideration of the two spatial regimes in the model. We should also be cautious with the LM tests results for spatial autocorrelation in the residuals: Once again, the robust tests must decide the best specification for the spatial autocorrelation pointed out by the ordinary LM tests. In this occasion, as in the other, only the robust LM test for spatial error dependence is significant, which could be pointing out the need of a respecification for model (12) as a spatial error spatial regimes model. The Kelejian-Robinson test, which is not affected by non-normal errors, also highlights the strong evidence of the presence of spatial dependence in the OLS error terms (though it is not capable of suggesting the proper alternative -spatial-lag or spatial error- model).

Therefore, the spatial regimes equation (model 12) although controls for the presence of spatial heterogeneity cannot totally absorb the existence of spatial dependence in the error terms, which has been clearly pointed out by both the LM and Kelejian-Robinson tests. Besides, we should also test for the potential **endogeneity** of some right hand side variables, which could be simultaneously determined with per capita GDP growth in a system feedback. In fact, as stated in Anselin and Kelejian (1997), there is some evidence that ignored endogeneity significantly affects the power of OLS-based tests against spatial error autocorrelation, Moran's I and LM tests.

If it were the case, these endogenous explicative variables would also be correlated with the error terms invalidating the OLS estimates (and associated statistical inference), which would no longer be consistent similarly to what happens in systems of simultaneous equations. We can assess this topic with the help of the **Durbin-Wu-Haussman** (DWH) test for the consistency of the OLS estimates when any endogeneity is present in a model; it has also been proposed as an "exogeneity test" (Anselin, 1999). In fact, it is an F test with (k^*, $n - k - k^*$) degrees of freedom on the null hypothesis of exogeneity of a k^* subset of the total k explicative variables, for n as the number of observations (for technical issues, see Davidson and McKinnon, 1993).

To check for the presence of potential endogenous variables on the right hand side of model (12), as OLS estimates would no longer be achieves consistency, we must estimate it with an instrumental variables approach (IV) such as **two stage least squares** (2SLS). The principle of the instrumental variables estimation is based on the existence of a set of instruments that are strongly correlated to the original endogenous variables but asymptotically uncorrelated to the error term. Once these instruments are identified, they are used to construct a proxy for the explicative endogenous variables which consists of their predicted values in a regression on both the instruments and the exogenous variables. In the standard simultaneous equations framework, the instruments are the "excluded" exogenous variables. In our case, we have analysed the potential system feedbacks between the dependent variable –per capita GDP growth in 1985-2001- and three explicative variables (initial per capita GDP has been excluded as has not an economic sense): per capita banks growth rate (*B*), percentage of population with university degrees (*D*) and per capita patents (*P*). So the instruments have been the excluded explicative variables (*K*, *U*, *E*, *T*, *R*, *EP*, *X*), as recommended in Anselin 1995A.

We have run this estimation for model (12) considering the existence of potential endogenity in *B*, *D* and *P*, but the DWH test only allows for the rejection of the null hypothesis on exogeneity in the case of *P* or per capita patents (p=0.0379)[16]. Consequently, in the current empirical exercise, we must treat *G85*, *B,* and *D* as exogenous, leading to the following model:

$$
\begin{aligned}
G_{i,T} &= \alpha^H + \beta^H G85_{i,0} + \phi_1^H B_i + \phi_2^H D_i + \phi_3^H P_i + \\
&\quad + \alpha^L + \beta^L G85_{i,0} + \phi_1^L B_i + \phi_2^L D_i + \phi_3^L P_i + \varepsilon_{i,T} \\
\varepsilon_i &\sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \\
IV(P) &= (K, U, E, T, R, EP, X)
\end{aligned}
\tag{13}
$$

It can be appreciated in Annex 2 that in this case the 2SLS estimation of model (13) produces more consistent but not necessarily more efficient estimates (with smaller variances), as it is the case, and a worse fit, in terms of the SSR (sum of squared residuals) coefficient value. The β parameters have experienced some changes, especially in the *H* sub-zone, where convergence speed rises to 19.8%, but it remains more or less the same in *L* (5.3%). The Chow-Wald test detects more instability over the regimes in the *G85* regression coefficient and patent coefficient is not significant in *L* space. We also have information about remaining spatial dependence in the errors thanks to an LM error dependence test (p=0.0173).

Therefore 2SLS errors are not random in model (13) confirming the results obtained by the LM and Kelejian-Robinson's autocorrelation tests in the previous models. The sources of this non-randomness could be, as suggested by the robust LMLAG test in model (12), an omitted spatial-lagged endogenous variable, leading to a spatial-lag spatial regimes model, or some omitted unobservable variables which could be spatially correlated, leading to an autoregressive structure in the error terms. At this point, in order to choose the best option, we should estimate both a spatial error spatial regimes and a spatial-lag spatial regimes model.

Firstly, the correction of the spatial autocorrelation present in the latter model with some endogenous explicative variables could imply a re-specification of model (13) into a **spatial regimes spatial error model**:

$$
\begin{aligned}
G_{i,T} &= \alpha^H + \beta^H G85_{i,0} + \phi_1^H B_i + \phi_2^H D_i + \phi_3^H P_i + \\
&\quad + \alpha^L + \beta^L G85_{i,0} + \phi_1^L B_i + \phi_2^L D_i + \phi_3^L P_i + \left[I - \lambda W\right]^{-1} u_{i,T} \\
u_i &\sim i.i.d.\left(0, \sigma_u^2\right) \\
IV(P) &= (K, U, E, T, R, EP, X)
\end{aligned}
\tag{14}
$$

where λ is a scalar spatial error coefficient. This expression indicates that a random shock introduced into a specific city will not only affect the growth rate in that city but, through the spatial multiplier $[I - \lambda W]^{-1}$, will impact the growth rates of other cities even if a given city had a limited number of neighbours. Model (14) has been estimated by 2SLS, as there is an endogenous explicative variable (*P*), using the same instruments for them as before. The spatial autoregressive error coefficient, λ, is treated as a "nuisance parameter" and therefore no inference or standard errors (or t-tests, etc.) are computed. In general, the results do not improve the later in terms of fit (RSS=0.0056) and patents variable is not significant at all. Neither the Chow-Wald test is significant (p=0.28). In this case, no remaining spatial autocorrelation error tests are reported.

Secondly, as model (14) does not provide an improvement at all, we should also re-specified model (13) as a **spatial regimes spatial-lag model**, in which spatial heterogeneity is explicitly considered with spatial autocorrelation through the dependent spatial lag variable (*WG*) as one explanatory variable which is as follows:

$$\begin{aligned} G_{i,T} &= \rho WG_{i,T} + \alpha^H + \beta^H G85_{i,0} + \phi_1^H B_i + \phi_2^H D_i + \phi_3^H P_i + \\ &\quad + \alpha^L + \beta^L G85_{i,0} + \phi_1^L B_i + \phi_2^L D_i + \phi_3^L P_i + \varepsilon_{i,T} \\ \varepsilon_i &\sim i.i.d.\left(0, \sigma_\varepsilon^2\right) \\ IV(P) &= \left(WG85, WB, WD, K, U, E, T, R, EP, X\right) \end{aligned} \qquad (15)$$

where ρ is a scalar spatial lag coefficient, *WG85* represents the log per capita GDP spatial lag variable; *WB* is per capita banks growth rate spatial lag variable; *WD* is the percentage of population with university degrees spatial lag variable.

The presence of the spatial lag on the right hand side part of the expression is similar to the inclusion of endogenous variables on the RHS in systems of simultaneous equations. So two endogeneity sources are present in model 7, spatial simultaneity (*WG*) and feedback simultaneity (*P*), which implies the use of proper instruments for the 2SLS estimation method. For the case of spatial-lag models, Kelejian and Robinson (1992) recommend the use of the set of spatially lagged exogenous variables as good instruments for the endogenous spatial lag variable. For this reason, these spatially lagged exogenous variables have been incorporated into the group of instrumental variables already used in the previous models.

This model is better in terms of fit (SSR=0.050) and there is no longer any problem with residual spatial dependence, as stated by the LMERR test (p=0.99). The spatial Chow test rejects the null hypothesis of joint equality of the coefficients, so in this model the two spatial regimes are quite significant. In the case of β coefficient, it exhibits some differences according to the regimes and the speed of convergence in cities located in the higher income space (11.9%) is more than the double above the convergence speed estimated for the lower income zone over the analysed period (4.5%). The implied half-life for the cities of the lower income space, i.e. the time necessary for this group of economies to fill half of the variation, which separates them from their steady state, is about 21 years while for the rest of cities the half-life decreases to 13 years.

This reformulation also has an interesting interpretation from an economic perspective: the growth rate of a city (*G*) is positively influenced by the average growth rate of its neighbouring cities, through the endogenous spatial lag variable (*WG*), after conditioning on the starting levels of income (*G85*), capital growth (*B*), human capital (*D*) and technological investment (*P*). This is a spillover effect which, together with the spatial regimes, indicates that the spatial association patterns are not neutral for the economic performances of Spanish cities. The more a city is surrounded by dynamic cities with high growth rates, the higher its growth rate will be. In other words, the geographical environment has an influence on growth processes.

In general terms, the standardized coefficients shows that per capita GDP at the initial moment and technology level (patents) are the most decisive variables on urban growth. But it is also interesting to highlight the different performance of the explicative (control) variables in both spatial regimes. In the high per capita GDP zone (North-East-Centre cities), the most important variables are in the following order: per capita GDP in the initial moment (-0.79), technology level (0.35), capital growth (0.28), human capital growth (0.25) and per capita GDP in neighbouring cities (0.15). And in the lower per capita GDP area (Southwest), the most influential variables follow a slightly different sequence: per capita GDP at the initial moment (-0.48), technology level (0.32), human capital growth (0.29), capital growth (0.25) and per capita GDP in neighbouring cities (0.15). As we can see, initial GDP levels have a strong weight in income growth in both sub-regions, but it is not so determinant in the backward zone.

VI. ECONOMIC INTERPRETATIONS OF THE RESULTS

The analysis of per capita income growth of the Spanish cities clearly indicates four important features to be emphasized. Firstly, per capita income level at the starting situation strongly determines the evolution of per capita GDP of the cities throughout the period, but there are also other factors, such as capital growth and technology level, that are also important. Secondly, there is some kind of spatial polarization of urban economies within Spain, which is expressed as two different spatial regimes in per capita GDP distribution of Spanish cities: lower per capita GDP in the Southwest of Spain, and higher per capita income towns, located in the North-East, as well as in Madrid and its north-central surrounding cities. Thirdly, a spillover effect has been detected in income growth or a spread movement from each city towards the surroundings areas. Fourthly, it has been demonstrated that there is a convergence trend in the set of the main Spanish cities.

In general terms, the growth of per capita GDP in the group of the main Spanish cities depends heavily on their **per capita income level in a previous moment of time**. In effect, income growth has been particularly intense in those cities located in provinces with lower income levels in 1985, where the non-agrarian sectors (particularly industry, public and private services) have registered a higher growth, have enjoyed a more intense capitalization process, and have also demonstrated an outstanding entrepreneurship capacity in an appropriate environment of competition (e.g. Almería, Balearic Islands, Madrid, Málaga…). As seen in Figure 2, it is also evident that the more dynamic cities –the "winners"- in terms of per capita GDP, are being increasingly concentrated in some Mediterranean provinces (Catalonia and Castellón) and in central Spain, leaded by Madrid. Nevertheless, in the Mediterranean Arc area, there are differences in terms of per capita GDP growth that should be highlighted. In fact, most south-eastern Mediterranean cities in the Community of Valencia and Murcia, which have received a vast amount of people within the last fifteen years, have grown at a lower speed compared to some Catalan and especially all the Andalusian cities that had lower income levels in 1985. It goes without saying that the Iberian North (Asturias, Lugo, Bask Country), and part of the West (León, Zamora) –the "losers"-, is declining due to the lack of thriving cities capable of regenerating their influence territories.

Capital growth, which has been approach in this paper by per capita banks variable, has also been a influent explicative factor in income growth during this period. In effect, the current real estate boom (rapid increase of prices) is harshly reflected in places with the highest per capita GDP -so the principal Spanish cities-, strong capitalization processes, utmost firms creation, economic potential or accessibility, located in the Mediterranean area (and in general in the eastern part of the country) are highly motivated by a strong demand for houses on the coastline, second residences and a high pressure of the land rents/costs on the final housing prices mainly in the chief cities. In fact, this last point is one of the main factors responsible for expelling people- particularly the youngest- from the biggest cities and from the old main economic activity centres, and for enhancing new market areas demanding more and more distance from the traditional city centres. Even so, two important aspects need to be pointed out: firstly, housing demand increase over years since the average size of dwellings is decreasing (but it is still much higher than the European size), the arrival of immigrants within the country is intense, and the forecast for the demand of the second residences (included the ones coming from foreign countries) is on the rise. Only the first factor would have needed two millions houses in the 1991-2001 period- for the same population volume- to offset the reduction of dwellings sizes. This strong demand is backed up or supported by an income consisted of two salaries, stemmed from increasing female activity rates particularly intense in the most dynamic urban areas. Secondly, sometimes the population growth of these new areas is faster than GDP expansion in per capita terms. So the urbanization process of the core Spanish areas and their surroundings cities will remain the same for years to come.

From the point of view of the **technology and human capital factors**, it can be said that urban income growth is also explained by the R&D activities and higher-education level; even though it should be underlined that the correlation of these last variables and per capita GDP is stronger in the lower income regime, which were more technologically backward in 1985. It might mean that in the higher income cities regime, as technological progress and human capital levels are stronger, they are not as equally followed by income spread as in the lower income ones. In other words, although there is a sort of decentralization process –to the South, East and metropolitan cities around Madrid- which are growing very fast, the traditional cores continue to retain the main keys of the economic growth, in terms of the employment qualification and technology. The same thing can be said about accessibility, TIC embedding, and its spread effects, which are highly correlated and positively affecting the market share growth, but concentrated around the main urban areas (as shown in Fig. 2, Barcelona and Madrid metropolitan towns).

In model (15), it is also shown that **per capita GDP in neighbouring cities** is also a decisive variable on urban income growth. That is, a spillover effect has been detected in income growth or a spread movement from each city towards the surroundings areas; in other words, the closer the cities are to higher/lower income cities the faster/slower they will grow. That has been especially the case of some “winner” cities that have benefited from a privileged location, closer to the economic cores, to reach dominant positions (Ávila, Guadalajara, Segovia, Toledo, Valladolid, Zaragoza, as seen in Figure 1). And the contrary situation affects some "loser" cities, which are in decline (e.g. Bilbao, Gijón, Ponferrada, Portugalete).

Finally, it has been demonstrated that there is a general **convergence trend** in the set of the main Spanish cities, which are growing much faster than Spanish regions (2%); this fact could be explained, in part, by the existence of more homogeneity existent in a group of main cities compared to a set of regions. Nevertheless, there are two convergence speed for Spanish cities depending on the spatial regime in which they are located. In effect, cities located in the higher income area grow almost three times faster (11.9%) towards their steady state than the ones in the lower income growth area (4.5%), what points out that cities located in the prosperous sub-space are growing much faster than those that are farther the main economic cores. Therefore, Spanish northern-eastern-central cities will reach its corresponding steady state 1.5 times first than southern-western ones, only due to their relative location.

VII. FINAL REMARKS

Bigger cities are not dying. These cities keep concentrating and absorbing in the latest backwash process the newest frontiers in matters such as technology, information society, communications and financial systems, headquarters and companies strategies, public and private decision hubs and accessibility to the international networks. This means that the bigger cities keep taking over the whole urban system through their higher economic potential and their stronger political power.

It remains to be tested to what extent the convergence trend detected in the group of Spanish main cities might be worse if the public administration was not so decentralized. The Spanish regional governments have been implementing many sector and transversal policies, which allow in different degrees of success, a development from the bottom, based on strategy planning, public/private cooperation and a more local commitment of officers, policy makers, law makers and politicians. In addition, the UE regional policy has contributed to a more rationale way of making and organizing policies by means of Regional Development Plans as a requirement to be beneficiary of the Structural Funds.

In conclusion, the empirical evidence of this paper is quite clear to the effect that there are two possible but contradictory forces in the evolution of the group of main Spanish cities. On the one hand, there is an empirical evidence of a general convergence process in the group of main Spanish cities that ensures more equality in terms of per capita income in a future time. But on the other hand, convergence speed is much slower in those cities located in the Spanish south-western backward regions. Therefore, although the integration in the EU economic environmet have derived strongest gains to the Spanish cities, as stated in Cheshire (2002), the integration of Europe could be favouring the core regions at the expense of the peripheral ones. Thus, the market mechanism and a more open competition are stronger than the regional policies and dominate the economies working, driving them to an accrued concentration of the production factors and the economic activity.

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Annex 1: Definition of the variables used in the models

Variable		Source
Dependent variable:		
G	Average per capita GDP growth rate, period 1985-2001 (in natural logarithms)	Estimation.
Explicative variables:		
Initial conditions:		
G85	GDP per capita in 1985 (in natural logarithms)	Estimation.
Capital:		
K	Dummy variable measuring the private and public capital growth by provinces (1985-1998).	FBBVA (2002)
B	Number of banks and savings banks per capita growth rate, period 1984-2001	Banesto (1985) "la Caixa" (2004)
Human resources:		
D	Percentage of population with university degrees	2001 Census (INE)
U	Unemployment rate growth rate, period 1991-2001	Ministry of Labour
Productive Structure:		
E	Percentage of no-agrarian employment change, by provinces (1985-2000)	Funcas (2003)
T	Tourism indicator: Taxes on tourism activity, taking into account number of rooms, category and occupation rate.	"la Caixa" (2004)
Technology:		
P	Per capita number of patents, 2000's	Spanish Patents Office, OEPM
R	Per capita R&D expenses, by provinces in 1994	INE (2000)
Location and economic potential:		
EP	Economic Potential Change: Index of growth of the economic potential, period 1987-2000	Keeble *et al.* (1988) Ministerio de Fomento (2000)
X	Dummy variable for the most expansive Spanish provinces in terms of real per capita GDP rate, period 1985-2000	Funcas (2003)

Annex 2: Estimation results for different models of urban per capita GDP growth

Model		Standard model (11)	Spatial Error	Spatial Lag	Spat. Regimes (12)	Spat. Regimes (13)	S.Reg. Sp.Error (14)	S.Reg. Spat-lag (15)
Estimation		OLS	ML	ML	OLS	2SLS	2SLS	2SLS
$\hat{\alpha}$	H	0.392648 (0.0000)	0.392088 (0.0000)	0.29935 (0.0000)	0.41497 (0.0000)	0.422717 (0.0000)	0.527722 (0.0002)	0.290157 (0.0000)
	L				0.387745 (0.0000)	0.386879 (0.0000)	0.321053* (0.0798)	0.255412 (0.0000)
$\hat{\beta}$ $\underline{\hat{\beta}}$	H	-0.039838 -0.592881 (0.0000)	-0.03939 (0.0000)	-0.034339 (0.0000)	-0.053478 (0.0000)	-0.059848 (0.0000)	-0.049695 (0.0000)	-0.053143 -0.790891 (0.0000)
	L				-0.036499 (0.0000)	-0.035920 (0.0000)	-0.030898 (0.0004)	-0.032271 -0.480257 (0.0001)
$\hat{\phi}_1$ $\underline{\hat{\phi}_1}$	H	0.000066 0.282272 (0.0000)	0.000067 (0.0000)	0.000064 (0.0000)	0.000103 (0.0019)	0.000080 (0.0264)	0.000092 (0.0170)	0.000066* 0.284756 (0.0583)
	L				0.000056 (0.0005)	0.000055 (0.0012)	0.000060 (0.0000)	0.000058 0.247798 (0.0004)
$\hat{\phi}_2$ $\underline{\hat{\phi}_2}$	H	0.000511 0.341149 (0.0000)	0.000534 (0.0000)	0.000482 (0.0000)	0.000459 (0.0001)	0.000450 (0.0002)	0.000528 (0.0005)	0.000377 0.251849 (0.0016)
	L				0.000486 (0.0015)	0.000439 (0.0374)	0.000435 (0.0204)	0.000430 0.287023 (0.0286)
$\hat{\phi}_3$ $\underline{\hat{\phi}_3}$	H	0.049788 0.154401 (0.0137)	0.035511 (0.0444)	0.047263 (0.0122)	0.051141 (0.0196)	0.126156 (0.0045)	0.046868** (0.2617)	0.112874 0.350044 (0.0073)
	L				0.025931** (0.6104)	0.070731** (0.6252)	0.164648** (0.8448)	0.104662** 0.324577 (0.4175)
$\hat{\lambda}$ / $\underline{\hat{\lambda}}$		-	0.488211 (0.0000)	-	-	-	0.994156	-
$\hat{\rho}$ $\underline{\hat{\rho}}$		-	-	0.249275 (0.0054)	-	-	-	0.362395 0.147839 (0.0017)
AIC		-868.265	-882.553	-873.663	-870.855	-	-	-
SSR		0.0053	0.0044	0.0047	0.0039	0.0053	0.0056	0.0050
JB		5.0292* (0.0809)	-	-	7.8252** (0.0200)	-	-	-
Breush-Pagan		10.1584 (0.0378)	10.0704 (0.0393)	9.1392* (0.0577)	0.3874** (0.5337)	-	-	-
White test		29.8710 (0.0080)	-	-	-	-	-	-
Wald common factor hipot.		-	1.0567** (0.9011)	-	-	-	-	-
LMERR		19.9944 (0.0000)	-	6.9681 (0.0082)	16.4453 (0.0000)	5.6612 (0.0173)	-	0.000001** (0.9993)
Robust LMERR		11.2249 (0.0008)	-	-	5.7566 (0.0164)	-	-	-
LMLAG		8.9033 (0.0028)	0.0552 (0.8142)	-	12.2068 (0.0005)	-	-	-
Robust LMLAG		0.1338** (0.7145)	-	-	1.5181** (0.2179)	-	-	-

Model		Standard model (11)	Spatial Error	Spatial Lag	Spat. Regimes (12)	Spat. Regimes (13)	S.Reg. Sp.Error (14)	S.Reg. Spat-lag (15)
Estimation		**OLS**	**ML**	**ML**	**OLS**	**2SLS**	**2SLS**	**2SLS**
Kelejian-Rob.		59.3803 (0.0000)	-	-	52.4706 (0.0000)	-	-	-
DWH		-	-	-	-	4.4155 (0.0379)	-	-
Spatial Chow		-	-	-	2.4351 (0.0390)	10.7051* (0.0576)	6.2675** (0.2810)	12.3316 (0.0305)
conv. speed	**H**	6.3%	6.2%	5.0%	12.1%	19.8%	9.9%	11.9%
	L				5.5%	5.3%	4.3%	4.5%
half-life	**H**	17	17	20	13	11	14	13
	L				19	19	22	21

Notes: *Standardized* coefficients are underlined and *p-values* are in parentheses. *Standard* is a non-spatial model. *Spat. Error* is a spatial error model. *Sp.Regimes* is a spatial regimes model. *S.Reg. S.Error* is a spatial regimes spatial error model. *S.Reg. S.Lag* is a spatial regimes spatial-lag model. *OLS* indicates ordinary least squares estimation. *2SLS* indicates two stage least squares estimation. *AIC* is the Akaike Information Criterion. *SSR* is sum of squared residuals. *JB* is the Jarque-Bera non-normality test on the residuals. *Breush-Pagan* is the Breush-Pagan test for heteroskedasticity. *White* is the White test for unspecified heteroskedasticity. *Wald common factor hipot.* is the Wald test on common factor hypothesis. *LMERR* is the Lagrange multiplier test for spatial autocorrelation in the error term. *Robust LMERR* is the Lagrange multiplier test for spatial autocorrelation in the error term robust to the presence of spatial lag dependence. *LMLAG* is the Lagrange multiplier test for an additional spatially lagged endogenous variable in the model. *Robust LMLAG* is the Lagrange multiplier test for an additional spatially lagged endogenous variable robust to the presence of spatial error dependence. *DWH* is the Durbin-Wu-Haussman test for exogeneity of *U*, *R* variables ($\hat{\phi}_2$ $\hat{\phi}_3$ coefficients). *Spatial Chow* is the spatial Chow-Wald test on spatial instability of the coefficients in two regimes: high per capita GDP (*H*) and low per capita GDP (*L*). *Conv.speed* is the convergence speed. *Half-life* is the time necessary for the group of cities to fill half of the variation, which separates them from their steady state.

[1] There are other non-spatial analysis, as in Chua (1993) and Carrington (2003). They propose a formulation that makes convergence conditional upon location. More specifically, they place output per effective unit of labour in a Cobb-Douglas relationship to the broad capital per effective unit of labour in one region and in its neighbourhood, with incorporation of a constant externality parameter. Here, we will follow the spatial econometric approach.

[2] In this case, the use of Ordinary Least Squares (OLS) in the presence of non-spherical errors would yield inconsistent estimators due to the presence of a stochastic regressor *"Wy"*. Therefore, this model must be estimated by ML or Instrumental Variables method (for a more extent review, Anselin, 1988).

[3] There are other similar experiences of urban growth models that include an estimated urban GDP data, as Thomas (2001) for French cities.

[4] GDP provincial data has been obtained from the Regional Accounts (INE), which provides two series of data, one from 1985 to 1996 and another from 1995-2001 (these two series have been standardized so they can be comparable). Both provincial and municipal data corresponding to the explicative variables in 1985 and 2001 have been taken from Banesto (1985) and "la Caixa" (2004), respectively.

[5] After a first non-spatial OLS estimation and the application of the corresponding hypothesis tests, especially on spatial autocorrelation over the OLS residuals, a spatial-lag model has been used for estimating per capita GDP in both 1985 and 2001 (for the classical spatial modelling strategy, see Anselin, 1988).

[6] The role of the spatial weight matrix is to introduce the notion of a neighbourhood set for each city. It has been also used other spatial weight matrices. These include an inverse distance matrix (such that each element w_{ij} is set equal to the inverse of the squared distance between cities i and j) and a matrix obtained from a 200 km. distance threshold to define a city's neighbourhood set (Rey and Montouri, 1999).

[7] According to FBBVA (2003), in Aragon, Castile-La Mancha, Castile and León, Ceuta and Melilla, the public capital protagonism has been more important than in the other Autonomous Communities. On the contrary, the Balearics Islands have registered a higher speed in the private capital accumulation. The Northeast of Spain - the closest Spanish area to the main markets of the EU- are those that attract more easily the private capital, whereas the peripheral territories situated in the Centre- South and West of the Iberian Peninsula- find more attraction difficulties. Besides, the smallest public capital endowment belongs to the most populated provinces (Madrid, Barcelona, Valencia, Seville, Alicante and Malaga).

[8] This index, which is published in "la Caixa" (2004) has been calculated from taxes imposed on this economic activity, which take into account not only the category of the establishments and the number of rooms, but also its occupation time.

[9] This estimation has been obtained by the request of our colleagues from the University of Barcelona, Esteban Sanromá and Raúl Ramos (to whom we are grateful for having provided us with this data).

[10] Concerning marine distances, we have taken the formula used by Keeble et al. (1988): the distance is equal to the distance by road to the closest port+ 150+ marine distance divided by 1,5. The formula does work quite well, except for Ceuta and Melilla.

[11] For the moment, it is not possible to connect our results with another experiences in Europe or the US. As for Mayor and López (2003) results, they cannot be directly compared with ours because they do not analyse big cities but the set of municipalities –urban and rural ones- existent in a Spanish region, much more heterogeneous.

[12] Several different spatial weights matrices have been used –all of them row-standardized- that reflect different a prori notions of the spatial structure of dependence: some k-nearest neighbours matrices for 4, 5, 6, 7 neighbours (Anselin, 1988), the square inverse distance weights matrix and the so-called distance based matrices for 100, 130, 150 and 220 km between the cities. But in this paper we will only refer to the results obtained with the distance based matrix for 130 km between cities.

[13] We have re-estimated model (11) but specifying the dummy of the previously defined 2 spatial regimes as a heteroskedastic variable. The resulting Koenker-Basset heteroskedasticy test clearly accepts the null ($p=0.67$), which confirms that the existence of these 2 spatial regimes is the main cause of spatial heterogeneity in the model.

[14] An unconstrained Durbin model was estimated (Anselin 1995A), as suggested by the Wald common factor test, but the corresponding explanatory variables spatial lags were neither significant.

[15] The Akaike Information Criterion (AIC) is a ML-based statitic that, as well as the log likelihood (LIK) measure, is appropiate to compare models estimated by different methods (e.g. OLS and ML). But AIC corrects the LIK for overfitting, which is very important when also comparing models with different number of regressors (Anselin, 1992). As it is well-known, the best model is the one with the lowest value for an information criterion.

[16] When running model (12) considering the existence of potential endogeneity in *B* and *D* explicative variables, it is not possible to reject the null hypothesis of exogeneity with the DWH test. In the case of per capita banks (*B*), the associated F-test *p-value*=0.4514 and in the case of percentage of population with university degrees (D), the corresponding *p-value*=0.3548. Note that per capita banks variable (*B*)

has also been used –with any other indicators- in the spatial extrapolation of provincial per capita GDP to municipal one, so some kind of simultaneous relationship (endogeneity) between *B* and per capita GDP growth could be previously expected (but cannot be confirmed by DWH test at all).