Urban Studies
46(9) 1969–1993, August 2009

Provincial Wages in Spain: Convergence and Flexibility

Adolfo Maza and Jose Villaverde

[Paper first received, July 2006; in final form, June 2008]

Abstract

This paper analyses the performance of provincial Spanish wages during the period 1985–2003. Using both traditional and recent analytical approaches, related to spatial econometrics, non-parametric and semi-parametric techniques, the paper reveals that a process of wage convergence has taken place. However, this does not preclude that mobility within the provincial wage distribution is small. The paper also shows that the degree of provincial wage flexibility is very low, which is in accordance with the institutional framework of wage formation in Spain.

Introduction

Spatial disparities are one of the most pervasive keynotes of the world economic scene. Although they do show up in a great number of economic variables, these spatial disparities are particularly worrying in variables such as per capita income and the unemployment rate. Thus, it is no wonder that the analysis of territorial convergence for these variables is a critical economic issue which has become very popular since the late 1980s. Although there are several reasons for this popularity, there is no doubt that the processes of economic integration being undertaken all over the world (globalisation), but mainly within the context of the European Union (EU), are among the most important (Ezcurra *et al.*, 2005). In particular, the latest enlargements of the EU represent a shock which, according to specific circumstances (regional differences in areas such as industry mix, technological development and social capital), might be termed idiosyncratic and have diverse effects in different regions, thus either fostering or hindering regional convergence.

Several factors have been put forward in the literature to explain territorial disparities and their changes over time. To a given extent, these disparities are the result of differences in the performance of regional labour markets, related to issues such as labour mobility, labour market flexibility and the

Adolfo Maza and **Jose Villaverde** are in the Department of Economics, University of Cantabria, Avda Los Castros, s/n, Santander, Cantabria, 39005, Spain.
E-mail: mazaaj@unican.es and villavej@unican.es.

0042-0980 Print/1360-063X Online

DOI: 10.1177/0042098009106018

sectoral composition of employment (for example, Armstrong and Taylor, 2000, p. 211; Martin, 2003).[1] In particular, labour market flexibility describes the functioning of labour market, in the sense of how quickly and to what extent it adapts to economic shocks. By this, we are referring not only to adjustments in quantities (in workers and working time) but also to adjustment in prices (wages), both factors being heavily influenced by the institutional setting of the labour market.[2]

Regarding wage adjustment, it is well known that wage flexibility is a key adjustment mechanism in determining labour market flexibility (Blanchflower, 2001) and, hence, the overall performance of each economy. This is even more evident in the case of the less developed countries, where most of their competitive advantage is related to sectors with low relative wages. Referring to the European Union of 15 members (EU-15), it is apparent that, due to factors such as its latest enlargements, the need for wage flexibility, especially for its less developed members, is even more patent in order to cushion the potential negative effects of these shocks on the performance of their economies and, then, on their rate of convergence with the EU average. As should be obvious, the need for this wage flexibility applies not only at national but also—and possibly to an even greater degree[3]—at subnational level.

Since Spain is one of the less advanced countries in the EU-15, we are able to use it as a sort of laboratory to analyse the evolution of real provincial wages and its determinants during the period 1985–2003. Thus, focusing on provincial real wages, this paper tries to provide an answer to two relevant and somewhat closely related questions: first, whether a wage convergence process has taken place in Spain; and, secondly, what the degree of wage flexibility is. The interest in analysing wage convergence lies in the fact that if this had happened, then a provincial convergence in per capita income should also have taken place (Emerson *et al.*, 1992). On the other hand, it seems that, given the supply shocks related to the most recent EU enlargements—along with the persistence of high unemployment rates in most Spanish provinces—the attainment of an appropriate degree of wage flexibility should be considered as a desirable goal. The links between wage convergence and wage flexibility are diverse. For instance, it might be the case that convergence is the result of high wage flexibility, in the sense that wage evolution reflects the particular conditions of the labour market of each province, such as changes in its unemployment rate and productivity level; however, if the wage-setting mechanism fixes the wage growth rate for each sector and the differences in the provincial industry-mix are declining, wage convergence across provinces might also be linked to the presence of wage rigidity (Buti and Sapir, 1998).

This paper departs from some previous studies (Jimeno and Bentolila, 1998; Bajo *et al.*, 1998) in at least two ways. First, we provide evidence of provincial wage convergence in Spain, an issue that has received scant attention in the literature; additionally, we do that not only by means of applying a classical approach to convergence (Sala-i-Martín, 1996) but, considering its drawbacks, also by paying attention to the whole distribution and—given that provincial wages are likely to be interdependent-its spatial characteristics. Secondly, we assess how flexible provincial wages in Spain are; although this is a much more researched issue (see, for instance, the aforementioned Jimeno and Bentolila, 1998; and Bajo *et al.*, 1998), we go beyond standard models of wage flexibility in two respects: on the one hand, by considering, once again, the spatial properties of the distribution and, on the other, by applying more refined analytical techniques of estimation.

The rest of the paper is organised as follows. To begin with, the main features of the institutional framework of wage bargaining in Spain are briefly discussed in the next section. After that, we present the data and conduct a classical analysis of convergence. Then, we investigate whether the phenomenon of spatial dependence is present in the provincial wage distribution, for which we carry out exploratory and confirmatory analyses. Next, and given the limitations of the previous convergence analysis, we consider the external form of the provincial distribution of real wages, its evolution over time and the degree of mobility for the Spanish provinces in that distribution. Afterwards, we analyse the degree of wage flexibility of the Spanish provinces, by means of using parametric and semi-parametric estimation techniques. Finally, we outline the main conclusions and some policy recommendations.

Institutional Setting of the Spanish Labour Market

Institutional differences between countries in the labour market can lead to divergent wage and income developments. Considering the existing interest in the performance of provincial real wages in Spain, a short but accurate description of the Spanish labour market institutions seems fitting.

To begin with, we should remember that, generally speaking, countries with either strong centralised or decentralised wage bargaining systems are better equipped to respond to supply shocks than countries with intermediate degrees of centralisation (Calmfors and Drifill, 1988). However, it is also true that the relationship between the degree of wage centralisation and wage flexibility is not a linear process. There are other factors—like the degree of openness of the economy, the degree of competition in product markets and other institutional characteristics of the labour market[4]—that can affect this relationship.

The institutional framework of the Spanish labour market can be considered as having an intermediate degree of centralisation. This being the case, the system is relatively different for private- and public-sector employees. The pay-setting conditions in the private sector are established by formal collective bargaining, where negotiations take place at different levels. In particular, the main characteristic of the wage-setting mechanism which applies to the private sector is the existence of a three-tier bargaining system, at the national, sectoral and company level. In fact, most of the collective bargaining is currently performed at sectoral level, although following rules established at national level. Thus, insofar as the industry-mix differs across provinces, this wage-setting scheme also allows for some provincial wage dispersion (Bande *et al.*, 2001). Other interesting characteristics are related to the minimum wage, unemployment benefits and firing costs. First, there is a minimum wage (wage floor) set up at national level that, although very low as a percentage of the average wage, only affects a small fraction of wage-earners. Secondly, the unemployment benefits, both their level and duration, are higher in Spain than in the EU. Finally, firing costs are similar in Spain to those in the EU when they refer to fair dismissals, but much higher when they are applied to unfair dismissals (Jackman *et al.*, 1999).

As for the public sector, it is essential to distinguish between two types of employees: civil servants (*funcionarios*) and state employees (*personal laboral*). The wage arrangements of the *personal laboral* are determined in a similar way to those of the private sector, which is to say, by formal collective bargaining. On the other hand, the wage conditions of civil servants are fixed by the employer (central administration, regional governments and local corporations), which apparently would suggest that wages should be uniform across Spain. This is clearly not the case, though, because "there is a wide regional variation in public wage gaps" (García and Jimeno, 2005, p. 15). This is for two reasons: first, there are, up to a certain point, informal negotiations between public employers, trade unions and employees representatives; and, secondly, regional governments and local corporations are allowed to be somewhat flexible when setting up the remuneration packages offered to their employees. For this reason, our analysis of provincial wages is focused on total wages and not only on private-sector wages as would be the case be if the public sector's wage growth were not set up by collective bargaining.

Taking into consideration these mechanisms of wage-setting, we can conclude—as was previously mentioned—that the Spanish labour market has an intermediate degree of centralisation. Even so, the wage-bargaining system in Spain is closer to being centralised than decentralised, because, although most of the negotiations take place at sectoral level, they follow guidelines which are being set up at national level. Therefore, sectoral wages tend to evolve in a relatively similar way in all Spanish provinces, this having seemingly adverse effects on wage flexibility.

Data

As is shown in Figure 1, Spain is organised into 50 provinces (NUTS 3) which make up 17 autonomous regions (NUTS 2). Considering that aggregation problems may arise when dealing with regions due to the fact that they are of widely differing sizes, we have chosen to carry out the analysis at provincial level, as mentioned in the introduction.

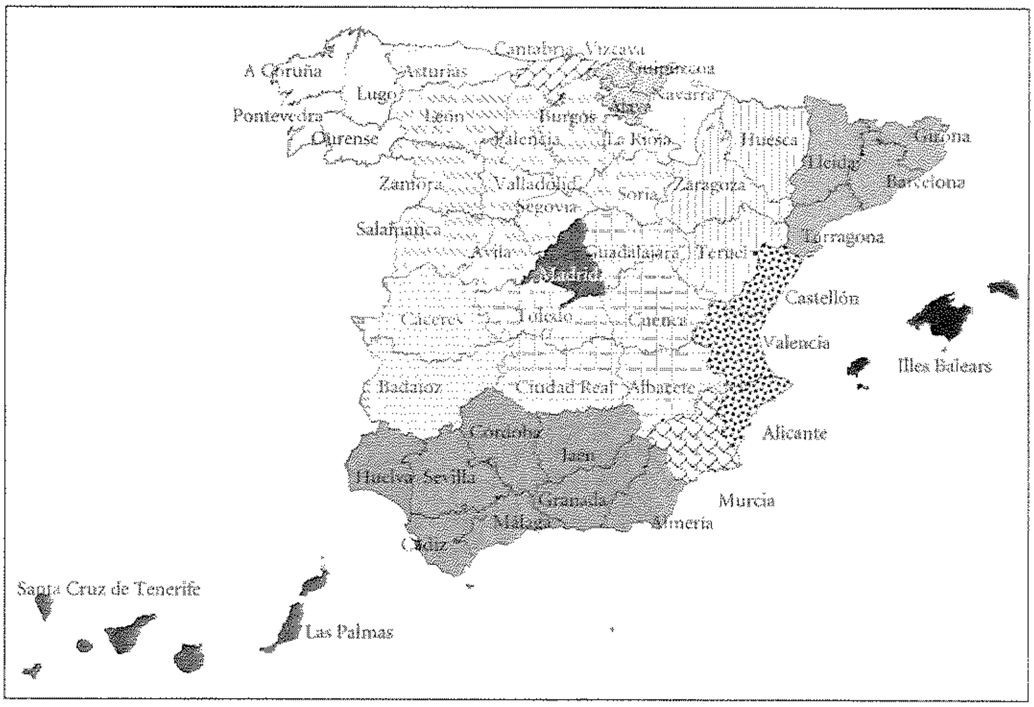

Figure 1. Provincial and regional map of Spain

Note: The provinces in same grey tone and trace belong to the same autonomous region.

In fact, the geographical size of Spanish provinces is relatively similar, which helps to minimise aggregation problems.

In this paper, we use two data sources: FUNCAS (Spanish Savings Bank Foundation) for data on real wages, unemployment rates and productivity and IVIE (Valencian Institute for Economic Research) for human capital. Total real wages have been calculated by deflating provincial nominal wages (ratio between total wages and the number of wage-earners) with the corresponding provincial price indices.[5] On the other hand, in order to compute unemployment rates, we have divided the figure for total unemployment by the active population, while productivity has been computed as the quotient between real gross added value (GAV) and total employment. For the human capital indicator (h) we have used the expression

$$h = \sum_{i=1}^{5} \phi_i A_i \qquad (1)$$

where, ϕ_i stands for the weight attached to human capital of level i; and A takes the values 0, 4, 8, 12 and 16 for i = 1, 2, 3, 4 and 5.[6]

Table 1 provides descriptive statistics of the four variables used in the empirical analysis of the paper plus the per capita GDP. Three main characteristics are worth mentioning: first, that a certain spatial clustering seems to exist in relation to all five variables, as can be conjectured by a simple glance at the table; secondly, that although provincial disparities are present in the five variables, these are the greatest in the unemployment rate; and, thirdly, that there is a clear correlation between the level of wages and those of the other variables, except for unemployment rates. This last result seems to anticipate that unemployment rates have no significant influence on wages, an outcome will be confirmed later on.

Classical Analysis of Wage Convergence

Convergence is an interesting but rather imprecise concept, with many interpretations. However, the most generally accepted measures of convergence are σ and β convergence. The first of these occurs when the dispersion between the economies diminishes;

Table 1. Spanish provinces: Descriptive statistics (national average 1985–2003 = 100)

	Wage	*GDPpc*	*Unemployment rate*	*Productivity*	*Human capital*
Almería	82.8	79.9	92.9	86.6	90.6
Cádiz	92.2	***69.6***	***187.7***	90.0	93.5
Córdoba	86.1	***69.1***	137.2	85.3	92.4
Granada	84.9	***62.7***	***143.5***	82.3	98.2
Huelva	85.1	76.4	137.8	91.2	89.4
Jaén	***81.9***	***67.3***	142.5	84.7	91.2
Málaga	92.9	77.4	***149.3***	94.5	94.7
Sevilla	91.3	70.8	***148.4***	91.5	99.6
Huesca	92.0	102.0	**42.3**	94.1	99.0
Teruel	98.8	101.5	**43.6**	97.9	93.2
Zaragoza	101.4	112.2	67.7	99.7	102.0
Asturias	106.5	89.4	110.0	95.6	97.7
Baleares	96.9	**132.3**	71.4	**117.9**	97.1
Las Palmas	96.0	98.6	113.3	101.5	96.2

(Continued)

(**Table 1.** *Continued*)

	Wage	*GDPpc*	*Unemployment rate*	*Productivity*	*Human capital*
Tenerife	94.9	96.5	111.8	103.0	96.8
Cantabria	101.1	94.5	93.0	94.1	100.9
Albacete	***81.5***	73.1	100.9	***78.1***	93.7
Ciudad Real	85.4	78.5	103.6	91.1	99.4
Cuenca	***79.7***	75.7	77.7	82.2	97.1
Guadalajara	93.3	107.9	76.4	103.6	94.7
Toledo	85.3	83.6	79.9	86.7	97.5
Ávila	87.4	79.2	78.3	***76.7***	97.8
Burgos	93.9	107.0	78.1	95.4	98.5
León	103.6	83.7	87.3	85.8	101.9
Palencia	94.0	96.1	94.2	93.3	***89.1***
Salamanca	92.7	87.1	121.0	91.2	94.4
Segovia	85.7	91.6	75.4	83.6	91.3
Soria	***80.1***	95.6	**50.2**	86.1	***89.3***
Valladolid	104.2	102.3	111.1	99.1	98.0
Zamora	88.5	75.0	96.1	82.8	***88.7***
Barcelona	**107.4**	**123.0**	92.4	**111.9**	**104.7**
Girona	96.0	**137.7**	**52.7**	105.1	98.8
Lleida	92.9	110.5	**30.4**	96.0	97.1
Tarragona	100.7	123.0	68.9	**119.3**	97.0
Alicante	90.5	93.0	106.2	96.1	94.0
Castellón	88.1	112.4	54.5	96.8	95.3
Valencia	94.6	105.4	103.5	100.6	101.0
Badajoz	***82.0***	***62.0***	***148.5***	***75.6***	92.6
Cáceres	82.4	77.6	118.5	90.2	90.2
A Coruña	93.7	88.3	94.3	87.3	90.9
Lugo	84.4	72.3	60.6	***57.4***	***84.1***
Orense	83.2	72.4	57.2	***61.6***	***85.1***
Pontevedra	90.7	84.2	99.8	79.5	91.2
Madrid	**118.9**	**128.5**	81.0	**118.5**	**114.2**
Murcia	84.9	83.1	105.5	89.0	95.4
Navarra	103.6	120.7	63.8	106.8	**104.8**
Álava	**113.2**	**135.7**	78.1	110.9	103.0
Guipúzcoa	**109.5**	109.9	99.9	108.6	**103.7**
Vizcaya	**114.9**	111.6	125.0	**119.6**	**110.7**
Rioja	93.1	118.8	60.8	99.1	98.4
CV	0.10	0.22	0.35	0.14	0.06
CC	1.00	0.70	-0.11	0.76	0.77

Notes: CV = coefficient of variation; CC = coefficient of correlation between wages and all variables. Provinces have been grouped by regions. The five best provinces for each variable are shown in bold while the five worst provinces are reported in bold italics.
Sources: FUNCAS, IVIE and own elaboration.

the second, when the poorer economies grow more quickly than the richer ones (Villaverde, 2006).

As a starting-point in this study, we calculate both these measures for wages in the Spanish provinces. First, σ convergence is computed by means of the coefficient of variation (Figure 2). As can be seen, the disparities have declined significantly—the coefficient fell by 31.7 per cent in the sample period (1985–2003), which represents an annual rate of convergence of 2.1 per cent.

With respect to the second type of convergence—less restrictive than the first—we begin by estimating a traditional absolute β convergence equation as follows:

$$\frac{1}{T}\log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right)=\alpha+\beta_1\log(\omega_{i,85})+\varepsilon_i \quad (2)$$

where, $\omega_{i,t}$ is the wage of province i in year t; T the number of sample years; and ε is the error term.

As is well known (Barro and Sala-i-Martín, 2004), in order for the hypothesis of convergence to be satisfied, there must be an inverse relation between the growth rate of wages and their initial level—i.e. β_i must be negative and significant at the standard levels. The results obtained (Table 2) demonstrate that there has been a process of convergence between provincial wages in Spain during the period 1985–2003. Moreover, the value of the β_i coefficient enables us to say that the convergence took place at an annual rate of 2.14 per cent.[7] Consequently, the time that would be required to cover half the gap separating the Spanish provinces from their stationary state (half-life) is 25.7 years.[8]

Considering that human capital is among the most prominent of the various factors that seem to hold a greater sway on wages, we have also estimated a conditional β convergence equation as follows

$$\frac{1}{T}\log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right)=\alpha+\beta_1\log(\omega_{i,85}) + \beta_2\log(h_{i,85})+\varepsilon_i \quad (3)$$

where, as wsas previously indicated, h stands for human capital.

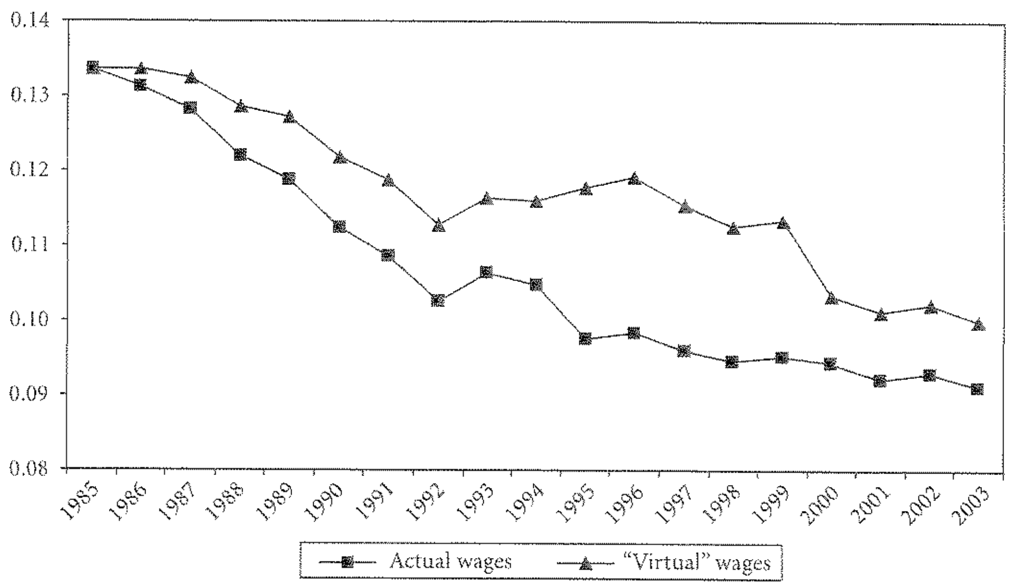

Figure 2. Plot of σ convergence

Not surprisingly, the estimated coefficient (β_2) on this variable is positive, showing that there is a direct relationship between the level of education and wages. Apart from this, the estimated speed of convergence increases from 2.14 to 2.79 per cent, thereby reducing the half-life by more than 8 years (Table 2).

So far we have used actual wages computed as a weighted average of the wages paid in different sectors. Thus, it should be clear that a portion of the aforementioned convergence might have been due to changes in the industry-mix. In order to remove this effect, we also consider 'virtual' wages computed under the assumption that each province has, for every year, the same employment structure as in the first year (1985) of the sample period;[9] therefore, provincial convergence between these 'virtual' wages simply reflects the evolution of provincial differences due to factors other than changes in the industry-mix of each province. The results obtained indicate that, as far as σ convergence is concerned, provincial differences have also decreased, but to a lesser degree (Figure 2); as for the β convergence, it is also shown that the speed of absolute convergence decreases from 2.14 to 1.75 per cent while that of conditional convergence goes from 2.79 to 2.36 per cent (Table 2). These results imply that changes in the industry-mix have enhanced wage convergence up to a point between the Spanish provinces, but not to any great extent. Thus, and for the sake of clarity, the rest of the paper will only deal with actual wages.

Spatial Analysis of Wage Convergence

So far, the convergence analysis has not considered the spatial distribution of wages. In other words, it has not taken into account the role potentially played by the geographical situation of each province. However, and for various reasons, consistent with assumptions and predictions based on endogenous growth models and new economic geographical theory, including similarities in the industry-mix, spillover effects, centrifugal movements due to saturation effects, stronger migration processes, etc., it seems logical to expect a certain spatial dependence to exist—i.e. we can expect provinces with higher (lower) real wages to tend to be geographically close to each other.

This idea—already mentioned in the descriptive analysis based on Table 1—seems to be corroborated by a simple glance at Figure 3, which shows provincial wage differentials in Spain taking the national average as equal to 100. Here it can be seen that provinces tend to be concentrated around similar wage levels; therefore, it does seem that a spatial analysis may be useful if we want to gain more insight on the situation of Spanish provincial wages.[10]

For this reason, we carry out the so-called exploratory spatial data analysis (ESDA). The most representative statistic of this kind of analysis is known as Moran's I.[11] The results obtained for this statistic, using as a distance matrix the inverse of the standardised distance, reveal the existence of a positive and statistically significant spatial dependence (see Figure 4); furthermore, spatial dependence, although undergoing many swings, has slightly increased over the sample period. This could suggest that there exists a globally upward tendency, although not very marked, towards geographical clustering of similar provinces in terms of relative wages.

The most widely used graphical representations of the presence of spatial dependence are the so-called Moran scatterplots and scattermaps. We have opted to use the scattermap, which divides the provinces into four groups: those with high real wages surrounded by provinces with high wages (high-high); those with high wages surrounded by provinces with low wages (high-low); those

Table 2. β convergence: Dependent variable: $\frac{1}{T} Log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right)$

	Actual wages				*'Virtual' wages*			
	Absolute convergence		*Conditional convergence*		*Absolute convergence*		*Conditional convergence*	
	Coefficients	*t-statistics*	*Coefficients*	*t-statistics*	*Coefficients*	*t-statistics*	*Coefficients*	*t-statistics*
Constant	0.277	7.94	0.346	9.69	0.217	6.56	0.275	7.84
Log($\omega_{i,85}$)	−0.027	−7.62	−0.040	−8.78	−0.021	−6.17	−0.031	−7.06
Log($h_{i,85}$)	—	—	0.029	3.82	—	—	0.024	3.25
R^2 adjusted	0.54		0.64		0.44		0.53	
LIK	215.09		221.85		217.70		222.77	
AIC	−426.18		−437.70		−431.40		−439.54	
SC	−422.36		−431.96		−427.57		−433.81	
Speed of convergence	2.14		2.79		1.75		2.36	
Half-life	25.7		17.4		33.0		22.1	

Sources: FUNCAS, IVIE and own elaboration.

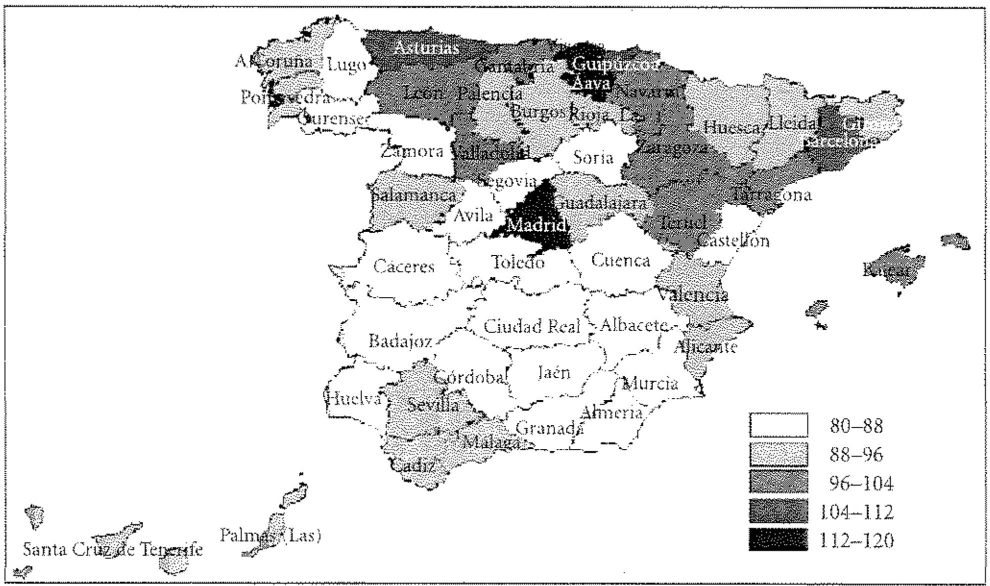

Figure 3. Provincial wage differentials (average 1985–2003)

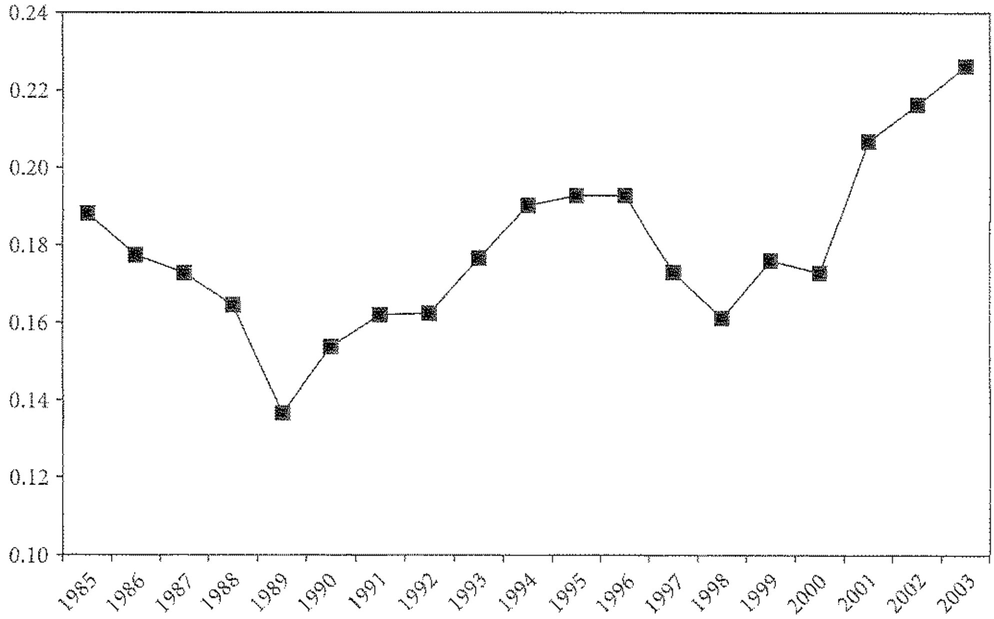

Figure 4. Spatial dependence (Moran's I statistic) on provincial real wages
Note: Moran's I statistics are significant at the 99 per cent level for all years.

with low wages surrounded by provinces with high wages (low-high); and, finally, those with low wages surrounded by provinces with low wages (low-low). The results displayed in Figure 5 show that, in general, the wage rate in a province is related to the wage rate in neighbouring provinces. From this, it can be seen that there is a phenomenon of positive

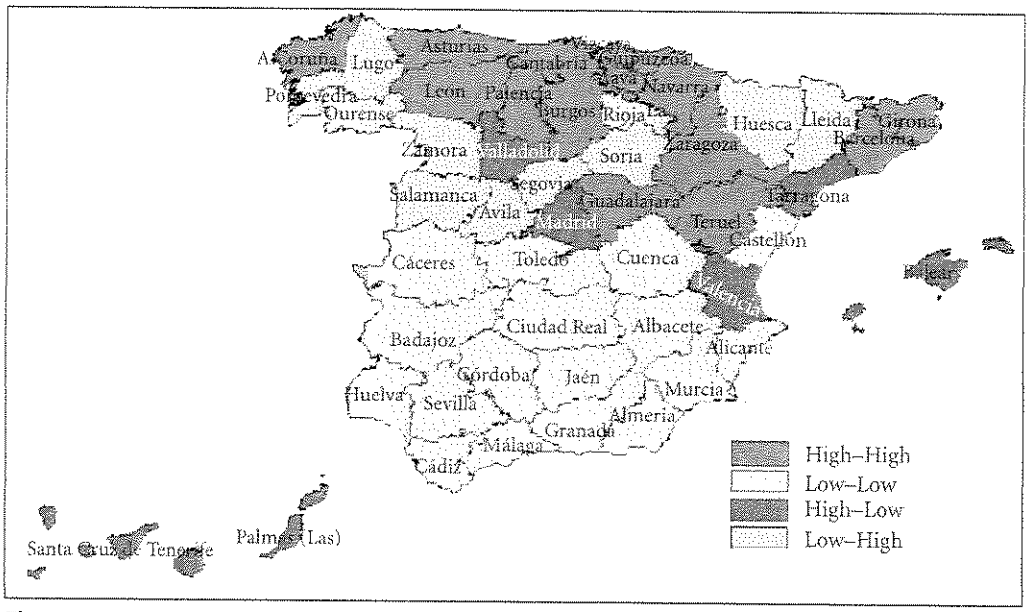

Figure 5. Moran scattermap (average 1985–2003)

spatial autocorrelation between the Spanish provinces, revealing a relatively clear north-east–south-west polarisation.

After having established the presence of spatial dependence in the distribution of wages, it is mandatory to revise the previous β convergence analysis, for if there are any problems of spatial autocorrelation in the estimated equations the results become inconsistent (Anselin, 1988). To conduct this confirmatory analysis, we perform a series of tests, with the Lagrange multipliers standing out, based on the principle of maximum likelihood.[12] The LM-ERR test, in particular, along with the associated robust LM-EL, tests for the absence of residual spatial autocorrelation, which would be caused by not including a structure of spatial dependence in the error term. The LM-LAG test is also used; this test, along with the associated robust LM-LE, tests for the absence of substantive spatial autocorrelation, which would be caused by the presence of spatial dependence in the endogenous variable. On observing the robust contrasts, it can be seen (Table 3) that the null hypothesis of absence of residual spatial dependence can be rejected at the 95 per cent level in both absolute and conditional convergence; on the other hand, there is no evidence of substantive spatial dependence. In short, these results indicate that it is necessary to correct the spatial autocorrelation of the residuals in the β convergence equations estimated earlier.

The model must be modified to introduce an autoregressive structure in the error term (ε_i) of equations (2) and (3), where $\varepsilon = \pi W \varepsilon + \tau$ with $\tau \approx N(0, \sigma^2 I)$. In this expression, π is the autoregressive parameter measuring the intensity of spatial autocorrelation in the error term, while W is the distance matrix: its elements $w_{i,j}$ reflect the intensity of the interdependence between provinces i and j. In this model, the effects of spatial dependence (diffusion) are exhibited in two ways, since the wage growth rate for province i is influenced, on the one hand, by the growth rates of the remaining provinces and, on the other, by its own initial wage level, weighted in both cases by W.[13]

Table 3. β convergence: Spatial contrasts

	Absolute convergence		Conditional convergence	
	Values	*P-values*	*Values*	*P-values*
I-Moran	2.203	0.0276	2.113	0.0356
LM-ERR	2.306	0.1288	1.613	0.2040
LM-EL	3.848	0.0480	4.081	0.0433
LM-LAG	0.031	0.8506	0.333	0.5639
LM-LE	1.572	0.2099	2.802	0.0942

Sources: FUNCAS, IVIE and own elaboration.

Table 4 reports the results of the maximum likelihood estimation of the new convergence equations.[14] The following points stand out

1. For both convergence equations, all the goodness-of-fit measures that are comparable between the aspatial and spatial models, such as the logarithm of maximum likelihood (LIK), Akaike's Information Criterion (AIC) and Schwartz's Criterion (SC),[15] reveal that equations including the autoregressive structure in the error term achieve a better fit than equations (2) and (3).
2. The value of the β_i coefficient, once the spatial autocorrelation in the residuals has been corrected, rises slightly in the two convergence equations. Thus the speed of convergence also increases and, consequently, the number of years required to cover half the gap from the stationary state then falls (see Tables 2 and 4).
3. The coefficients π are positive and statistically significant, confirming the results of the earlier spatial autocorrelation tests. This means that a shock in a specific province will not only affect the wage growth rate of that province but spills over to all or part of the national territory (Rey and Montouri, 1999).

Provincial Distribution of Wages

Both classical analysis and spatial analysis of convergence are very informative, but neither is without its problems. Perhaps the best-known criticism refers to the fact that they offer no information about the internal dynamics of the distribution examined (Quah, 1996, Durlauf and Quah, 1999, Magrini, 1999; Le Gallo, 2004), since they only capture some moments of it. In order to overcome this drawback, the paper tries to look more deeply into the provincial distribution of wages both in terms of its external form and of the changes occurring within it. This should provide not only (additional) evidence on the process of convergence, but also some hints as to the formation of poles of provinces with clearly differentiated wage rates.

Before doing that, and taking into account the spatial dependence previously observed in relative wages, we use the filtering methodology proposed by Getis (1995), based on a spatial statistic developed by Getis and Ord (1992).[16] This filtering procedure is designed to convert spatially dependent variables (X) into spatially independent variables (X^F); in this case, it implies that the difference between ω and ω^F is a new variable, L, which represents the spatial effects embedded in ω (Getis, 1995, p. 175). The filtering formula is defined as follows

$$\omega_i^F = \omega_i \frac{\sum_j w_{ij}(\delta)}{(N-1)G_i(\delta)} \qquad (4)$$

with

$$G_i(\delta) = \frac{\sum_j w_{ij}(\delta)\omega_j}{\sum_j \omega_j},\ i \neq j \qquad (5)$$

Table 4. β convergence: Spatial estimation: Dependent variable: $\frac{1}{T} Log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right)$

	Absolute convergence		*Conditional convergence*	
	Coefficients	*t-statistics*	*Coefficients*	*t-statistics*
Constant	0.291	8.17	0.363	9.96
Log($\omega_{i,85}$)	−0.029	−7.86	−0.041	−9.03
Log($h_{i,85}$)	—	—	0.028	3.86
π	0.453	1.99	0.462	2.01
LIK	216.26		222.78	
AIC	−428.51		−439.56	
SC	−424.69		−433.83	
Speed of convergence	2.26		2.87	
Half-life	23.9		16.8	

Sources: FUNCAS, IVIE and own elaboration.

and where δ is a distance parameter indicating the extent to which further distant observations are down-weighted.

To apply this filter, the inverse of the standardised distance is, once again, used as a distance matrix. Therefore, we assume that

$$w_{ij}(\delta) = (d_{ij})^{-\delta}$$

where, $\delta = 1$; and d_{ij} is the distance between province capitals i and j.

With respect to the external form of the provincial wage distribution, we focus our attention on the initial and final years of the sample, for which we estimate the density functions by using a Gaussian kernel with optimal bandwidth, computed as in Silverman (1986, p. 40).[17] The outcomes obtained with the filtered data are plotted in Figure 6, which offers relevant information about the changes that have occurred during the sample period. A close look at the figure shows that the shape of the distribution has undergone two main changes: first, there is a clear trend towards the concentration of mass probability around the average, this confirming the convergence process referred to earlier; secondly, a small peak (or pole) occurs at wage rates well below the national average in 2003. In order to examine the extent to which the entire cross-sectional distribution is modified when the spatial dimension is included in the analysis, we also present the density functions obtained by using actual data (Figure 7);[18] concordant with the rising value of Moran's I statistic, it can be seen that, although there are no major differences in 1985, the external shape of the distribution is somewhat different in 2003: in particular, the concentration of mass probability around the average is much greater with filtered rather than actual data.

Analysing the external form of the distribution based on density functions offers additional information to that provided by classical convergence analysis. It is, though, also true that it neglects one key factor: it does not give any information about the changes occurring within the distribution. Thus, in order to analyse the intradistribution dynamics, we use a continuous approximation based on the calculation of stochastic kernels.

The results obtained are displayed in Figures 8 and 9. The x-axis represents wages for the year 1985 and the y-axis wages for the year 2003, while the z-axis measures the density (or conditioned probability) of each point on the x–y plane. The lines parallel to the y-axis show the probability of moving from a point considered on the x-axis to any other

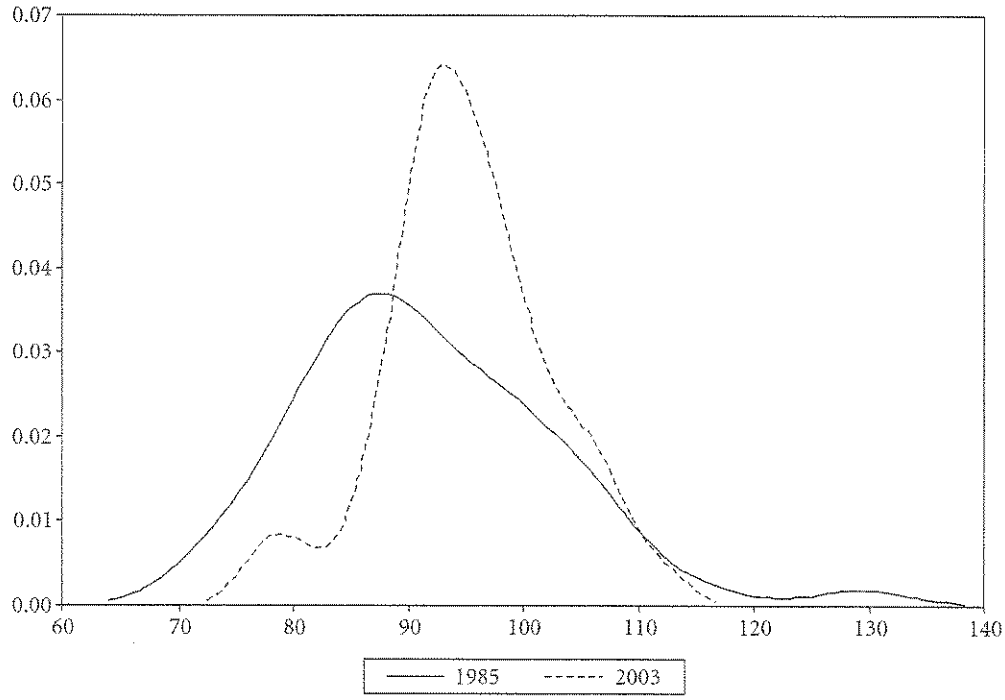

Figure 6. Density functions (filtered data)

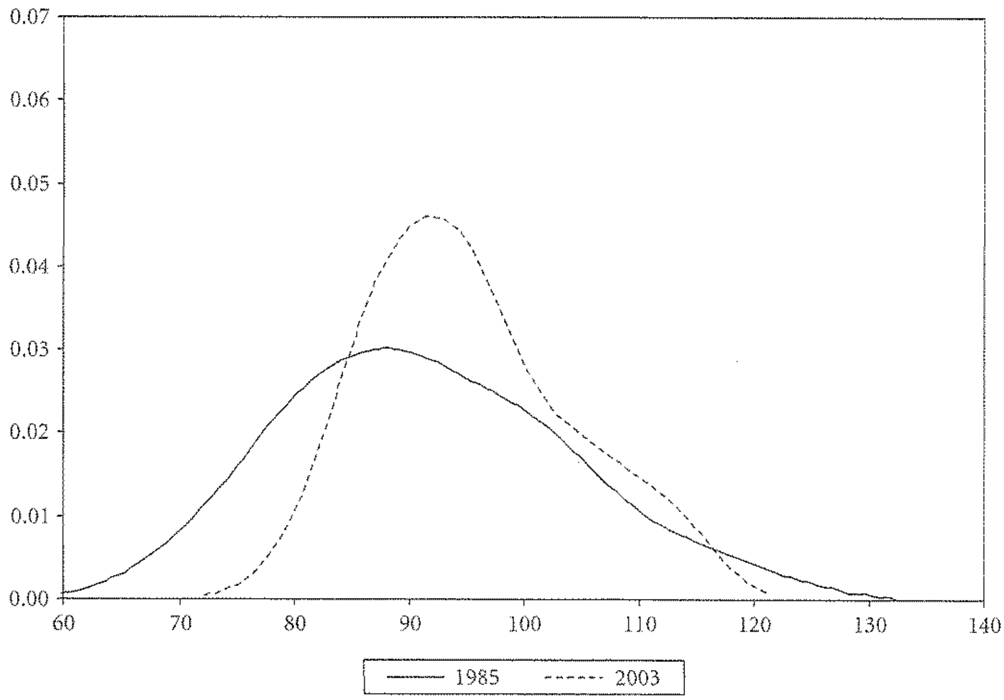

Figure 7. Density functions (actual data)

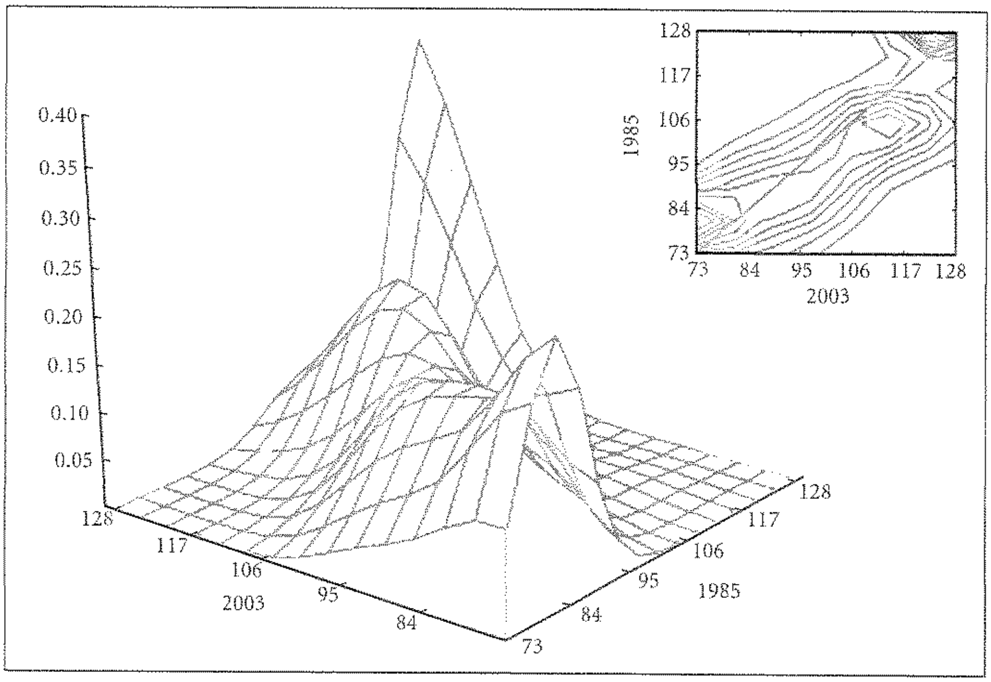

Figure 8. Intradistributional dynamics (filtered data)

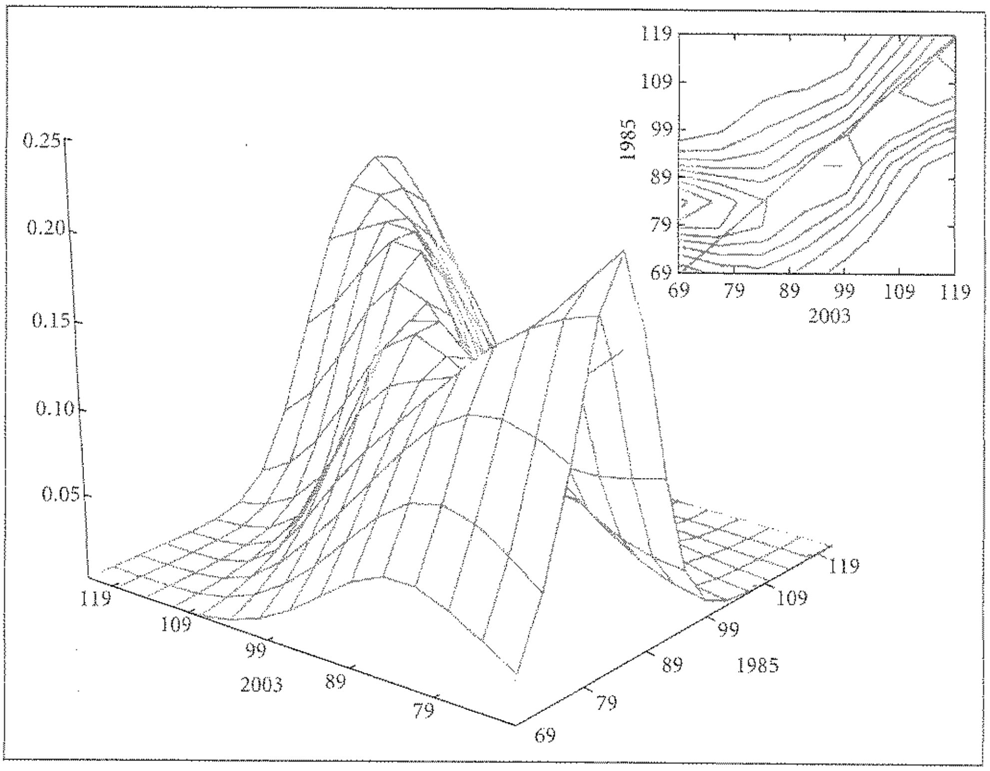

Figure 9. Intradistributional dynamics (actual data)

point on the y-axis. The two-dimensional figure on the right is the contour plot, obtained by taking a cut parallel to the x-y plane for a particular density value; the lines of this contour plot connect points reflecting the same density on the three-dimensional graph. Interpreting the kernel is easier if we look at this latter graph. If the contour lines are situated along the positive diagonal, the level of mobility is low; and the tighter the contour lines are to the diagonal, the lower this level is: this means persistence. If this is not the case, there has been mobility, with more mobility the further the lines are from the diagonal. According to this, Figure 8 shows that, generally speaking, the distribution of wages in the Spanish provinces has been characterised by persistence; however, the graph also shows that some provinces have greatly changed their position (note the width of the contour lines), thereby indicating a certain degree of mobility. Regarding this, a close look at the data shows that, for example, Ávila, Albacete and Cuenca have significantly improved their position, while Baleares, Asturias and Tenerife have worsened theirs. In this respect, the differences between using filtered and actual data (Figure 9) are not very significant.

Wage Flexibility of Spanish Provinces

Apart from having shown some changes in the external shape of provincial wage distribution together with a general tendency towards persistence in positions, the previous sections of the paper have also revealed that there has been a process of provincial convergence in wages in Spain between 1985 and 2003. As was pointed out in the introduction, several factors can be put forward to explain this result. In particular, there is no doubt that this convergence process is related to issues such as provincial differences in the sectoral distribution of employment, sectoral productivities and the degree of both labour mobility and wage flexibility across provinces.

In this section, we assess the degree of wage flexibility characterising the Spanish provinces, which is related to the institutions of the labour market and, in particular, to the wage-setting mechanism. Wage flexibility mainly refers to the speed and extent to which wages react to macroeconomic conditions, mainly unemployment and productivity changes. As regards unemployment, it is interesting to note that—as shown in Table 5—Spain is characterised by very high unemployment rates at national level and a high and increasing degree of provincial dispersion. Another major issue that can also be gleaned from the table is the persistence of the distribution of unemployment in the Spanish provinces, this meaning a substantial stability in the ranking; the need for wage flexibility seems to be—at least for this reason—quite clear. Regarding productivity, Table 5 shows there are striking differences in provincial disparities when they are compared with the unemployment rate—they are lower and decreasing—but the persistence in the ranking is also present, mainly in relation to provinces with the lowest productivity levels.

In order to test whether wages have been flexible at the provincial level in Spain, we estimate the following equation, obtained from a model of wage negotiation along the lines of Abraham (1996)

$$\dot{\omega}_{it} = \chi_i + \gamma_1 \dot{\omega}_t + \gamma_2 \dot{u}_{it} + \gamma_3 \dot{\lambda}_{it} + \eta_{it} \qquad (6)$$

where, the wage growth rate of province i ($\dot{\omega}_{it}$) depends on factors specific to that province (χ_i) and on the growth rates of the national wage ($\dot{\omega}_t$), provincial unemployment rate ($\dot{u}_{it}$) and provincial productivity ($\dot{\lambda}_{it}$).

If the influence of the national wage growth rate were very strong, wages would be characterised by rigidity, with very similar increases in each province regardless of its specific conditions. In contrast, if it was the effect of changes in unemployment and

Table 5. Evolution of unemployment rate and productivity in Spain

	National average	*Coefficient of variation*	*Maximun*	*Minimum*
Unemployment rate				
1985	17.4	0.31	24.4 (Badajoz)	5.3 (Lleida)
1990	16.4	0.35	31.5 (Cádiz)	4.6 (Lleida)
1995	19.4	0.35	34.9 (Cádiz)	6.1 (Lleida)
2000	11.9	0.42	24.4 (Cádiz)	3.5 (Lleida)
2003	10.7	0.45	24.1 (Cádiz)	2.9 (Lleida)
Productivity				
1985	26948	0.20	133.3 (Baleares)	46.2 (Orense)
1990	32258	0.16	123.8 (Madrid)	50.5 (Lugo)
1995	33020	0.14	120.6 (Vizcaya)	59.6 (Lugo)
2000	33554	0.11	119.6 (Tarragona)	66.1 (Lugo)
2003	33637	0.10	122.3 (Vizcaya)	71.9 (Lugo)

Note: Productivity levels are measured in constant 1995 euros. Provincial productivity levels are computed considering the national average equal to 100.
Sources: FUNCAS and own elaboration.

productivity in each province on the growth rate of its wage that was strong, then wages would be considered as flexible.

Before estimating equation (6), and considering once again the recent attention paid to invalid inference in regression models with spatial dependence in the residuals, it seems necessary to investigate the spatial properties of the data included in this equation. The results (Table 6) show that there is little sign of spatial autocorrelation in the data as Moran's I statistic is only significant at the 95 per cent level in 10 out of 54 cases. This result could seem surprising, but it must be taken into consideration that we are using variables in growth rates and the presence of spatial dependence tends to be more likely in levels than in growth rates. For this reason, we have opted not to filter these variables and estimate equation (6) using actual variables.[19]

We have computed equation (6) by OLS with fixed effects and the results obtained are reported in Table 7. It can be seen that the factor having the greatest effect on the evolution of wages in each province appears to be the growth rate of the national wage (with a coefficient of 0.876); as we have just mentioned, this is a clear sign of rigidity.[20] With respect to the provincial variables, it seems that, on the one hand, the influence of unemployment growth is negligible and, on the other, that increases in productivity are partially embedded in the evolution of provincial wages, since the coefficient associated with this variable has a value of 0.140 and differs statistically from zero. Finally, and with regard to the fixed effects, the table shows that they are statistically significant in only 10 provinces; this result means that, for these provinces, the rate of growth for wages is affected by idiosyncratic factors such as differences in industry-mix,[21] savings and/or investment rates, population growth and public capital. When these fixed effects are negative (positive) they reflect that, *ceteris paribus*, the corresponding rates of growth for wages are lower (greater) in these provinces than those obtained in the rest. In particular, the provinces with positive fixed effects

Table 6. Spatial properties of provincial variables included in equation (6)

	$\dot{\omega}_{it}$		$\dot{u}_{it}$		$\dot{\lambda}_{it}$	
Years	*Moran's I*	*P-values*	*Moran's I*	*P-values*	*Moran's I*	*P-values*
1986	0.023	0.454	0.110	0.023	0.070	0.117
1987	0.094	0.047	0.074	0.100	0.107	0.026
1988	0.062	0.153	−0.014	0.907	−0.009	0.838
1989	0.207	0.000	0.044	0.263	0.058	0.171
1990	0.053	0.204	0.079	0.085	−0.025	0.930
1991	−0.045	0.674	0.006	0.644	−0.049	0.614
1992	0.062	0.151	−0.038	0.755	−0.071	0.378
1993	0.117	0.017	−0.044	0.680	0.111	0.022
1994	−0.036	0.781	0.019	0.490	0.031	0.374
1995	0.053	0.204	0.048	0.235	−0.029	0.881
1996	−0.003	0.765	0.014	0.547	0.053	0.200
1997	0.045	0.254	0.054	0.196	−0.020	0.993
1998	0.106	0.028	0.115	0.019	0.090	0.056
1999	0.052	0.210	−0.087	0.245	0.010	0.597
2000	0.052	0.210	0.046	0.249	0.033	0.352
2001	0.019	0.488	−0.119	0.087	−0.123	0.075
2002	−0.065	0.434	−0.042	0.710	−0.014	0.908
2003	−0.027	0.907	0.217	0.000	0.353	0.000

Sources: FUNCAS and own elaboration.

(Albacete, Ciudad Real, Ávila, Cuenca, Cáceres and Jaén) are located in the centre-south of Spain while those with negative effects are either islands (Baleares and Tenerife) or are located in the north-west of the country (Lugo and Orense). To a certain extent, this is related to the notable rise in the ranking of Ávila, Albacete and Cuenca and the previously mentioned fall of Tenerife and Baleares.

According to the previous results, it appears that wages do not respond very much to the specific conditions of each province; they only respond to some extent to variations in productivity. Moreover, it could be added that improvements in productivity only have important effects on wages when the former are very substantial (Maza and Moral-Arce, 2006), a circumstance that a parametric estimation does not reveal. Thus, in order to know whether this is the case, there are more flexible ways of estimating the regression function of provincial productivity growth on provincial wage growth. In fact, this can be done without having to make any prior assumptions about its form, except that it is a smooth function (Hardle *et al.*, 1999). To deal with this issue, we have used a semi-parametric estimation technique, which merges a traditional parametric procedure with a non-parametric approach. Therefore, we estimate a new version of equation (6) as follows

$$\dot{\omega}_{it} = \chi_i + \gamma_1 \dot{\omega}_t + \gamma_2 \dot{u}_{it} + m(\dot{\lambda}_{it}) + \eta_{it} \quad (7)$$

where the only change is the inclusion of the variable representing increase in regional productivity in a non-parametric form—i.e. we allow its effects on real wages to be non-linear.[22] To compute this equation, we have followed the semi-parametric estimation process detailed in Li and Stengos (1996).

In this case, it can be seen (Table 8) that the coefficient associated with the national

Table 7. Wage flexibility (parametric estimation): Dependent variable: $\dot{\omega}_{it}$

	Coefficients	*t-statistics*[a]		*Coefficients*	*t-statistics*[a]
$\dot{\omega}_t$	0.876	30.52	León	0.000	–0.04
$\dot{u}_{it}$	–0.005	–0.81	Palencia	0.003	0.93
$\dot{\lambda}_{it}$	0.140	7.80	Salamanca	0.006	1.84
Fixed effects			Segovia	0.001	0.25
Almería	0.003	1.01	Soria	0.000	0.15
Cádiz	0.002	0.63	Valladolid	0.001	0.43
Córdoba	0.004	1.34	Zamora	0.004	1.15
Granada	–0.001	–0.33	Barcelona	0.000	–0.10
Huelva	0.004	1.25	Girona	–0.002	–0.53
Jaén	0.008	2.71	Lleida	–0.003	–1.01
Málaga	–0.001	–0.28	Tarragona	0.000	–0.08
Sevilla	0.003	1.01	Alicante	0.004	1.15
Huesca	0.001	0.38	Castellón	0.003	1.13
Teruel	0.000	–0.08	Valencia	0.001	0.16
Zaragoza	0.000	0.12	Badajoz	0.003	0.99
Asturias	–0.006	–1.88	Cáceres	0.009	2.83
Baleares	–0.009	–2.89	A Coruña	–0.002	–0.55
Las Palmas	–0.004	–1.35	Lugo	–0.007	–2.31
Tenerife	–0.006	–2.02	Orense	–0.008	–2.58
Cantabria	0.000	–0.14	Pontevedra	–0.006	–1.81
Albacete	0.011	3.69	Madrid	–0.004	–1.28
Ciudad Real	0.007	2.34	Murcia	0.000	–0.08
Cuenca	0.009	2.87	Navarra	0.002	0.77
Guadalajara	0.002	0.56	Álava	–0.001	–0.26
Toledo	0.006	1.85	Guipúzcoa	0.002	0.73
Ávila	0.010	3.29	Vizcaya	–0.002	–0.52
Burgos	0.002	0.68	Rioja	0.005	1.61
			R^2 adjusted	0.68	

[a] The results using the White heteroscedasticity consistent covariance matrix estimator for statistical inference in the OLS estimation are equivalent to those shown in the table.

Sources: FUNCAS and own elaboration.

wage growth rate increases slightly in value, which reinforces our previous diagnosis of wage rigidity. Likewise, the provincial unemployment growth rate still has no influence on wages, while the fixed effects largely maintain similar values. With regards to the non-parametric variable, Figure 10 not only shows that the influence of productivity on wages is indeed non-linear but also that it turns out to be particularly intense when productivity growth approaches or exceeds 10 per cent.

Conclusions

This paper has provided new insights into the nature of provincial wage performance in Spain for the period 1985–2003. To be precise, the paper analyses two different, yet related questions: the convergence and degree of flexibility of wages. To this end, our empirical strategy has been based on using both traditional and recent approaches related to spatial econometric, non-parametric and semi-parametric techniques. The most

Table 8. Wage flexibility (semiparametric estimation): Dependent variable: $\dot{\omega}_{it}$

	Coefficients	*t-statistics*[a]		*Coefficients*	*t-statistics*[a]
$\dot{\omega}_i$	0.895	31.66	León	−0.001	−0.22
$\dot{u}_{it}$	−0.004	−0.75	Palencia	0.002	0.61
$\dot{\lambda}_{it}$	n.p.v.[b]	n.p.v.[b]	Salamanca	0.005	1.57
Fixed effects			Segovia	0.000	0.06
Almería	0.002	0.75	Soria	0.000	−0.06
Cádiz	0.001	0.26	Valladolid	0.000	0.04
Córdoba	0.003	0.96	Zamora	0.003	0.97
Granada	−0.002	−0.60	Barcelona	−0.002	−0.56
Huelva	0.003	0.82	Girona	−0.003	−0.92
Jaén	0.007	2.38	Lleida	−0.004	−1.35
Málaga	−0.003	−0.88	Tarragona	−0.002	−0.52
Sevilla	0.002	0.68	Alicante	0.002	0.64
Huesca	0.000	−0.04	Castellón	0.003	0.84
Teruel	−0.001	−0.28	Valencia	−0.001	−0.23
Zaragoza	−0.001	−0.23	Badajoz	0.002	0.72
Asturias	−0.007	−2.30	Cáceres	0.008	2.64
Baleares	−0.011	−3.48	A Coruña	−0.002	−0.80
Las Palmas	−0.006	−1.89	Lugo	−0.007	−2.42
Tenerife	−0.008	−2.48	Orense	−0.008	−2.62
Cantabria	−0.001	−0.46	Pontevedra	−0.007	−2.19
Albacete	0.011	3.43	Madrid	−0.006	−1.80
Ciudad Real	0.006	2.07	Murcia	−0.001	−0.47
Cuenca	0.009	2.88	Navarra	0.001	0.46
Guadalajara	0.000	−0.04	Álava	−0.002	−0.69
Toledo	0.005	1.59	Guipúzcoa	0.001	0.33
Ávila	0.009	3.08	Vizcaya	−0.003	−0.92
Burgos	0.001	0.40	Rioja	0.004	1.31

[a] The results using the White heteroscedasticity consistent covariance matrix estimator for statistical inference in the OLS estimation are equivalent to those shown in the table.
[b] The symbol 'n.p.v.' denotes the non-parametric variable.
Sources: FUNCAS and own elaboration.

relevant conclusions deriving from our analysis are as follows.

First, and contrary to the expected increase in provincial wage differentials related to the ever more prominent role of globalisation and the shift in production towards services, the results obtained by carrying out traditional analysis of σ and β (absolute and conditional) convergence reveal the existence of a wage convergence process between the Spanish provinces. As has been shown by the estimation carried out with 'virtual' wages, this process is partly explained by changes in the provincial industry-mix. This is obviously related to the wage-setting mechanism operating in Spain which is mainly defined at sectoral level.

However, the presence of spatial autocorrelation problems in the residuals of the classical β convergence equations has shown that these are mis-specified, which led us to re-estimate them. The results obtained

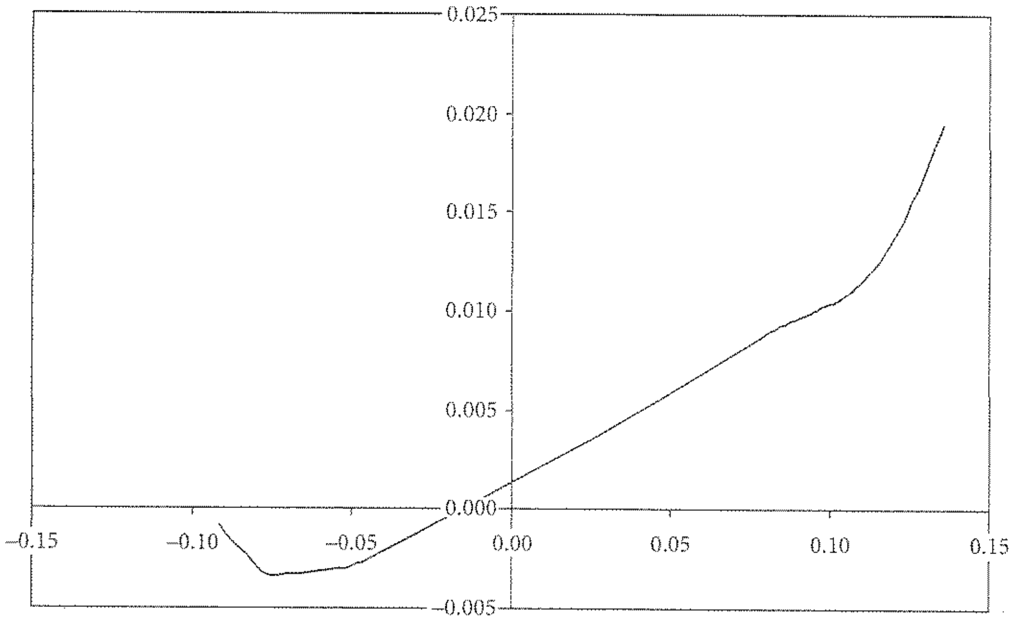

Figure 10. Semi-parametric estimation, $m(\dot{\lambda}_{it})$

confirm the existence of a convergence process, slightly higher than that obtained in the classical models.

Secondly, the external form of the provincial distribution of wages has varied significantly over time, with more provinces positioned around the mean in 2003 than in 1985. However, the analysis of the intra-distributional mobility—based on the estimation of stochastic kernels—demonstrates that, in general terms, provinces with high (low) wages in 1985 also tend to have high (low) wages in 2003; yet, this persistence does not mean that some provinces have not undergone great positional changes.

Thirdly, the evolution of provincial wages is found to be indexed, above all, to the evolution of the national wage and, to a lesser extent, to the productivity growth of each province, with the growth of provincial unemployment rates playing no role whatsoever in wage negotiations. While these results corroborate previous findings (for example, Bajo *et al.*, 1998, for the period 1989–92 in Spain; and Decressin and Fatás, 1995, for the EU), the paper has also been able to provide additional insights through the using of semi-parametric techniques. In this case, the analysis shows that the response of wages to variations in productivity is stronger when productivity growth is around or higher than 10 per cent. As a result of all of these findings, it can be said that the degree of wage flexibility in Spain is very low.

As mentioned in the introduction, the links between wage convergence and wage flexibility are many and diverse. However, this paper has revealed that, as far as the Spanish provinces are concerned, there is an inverse relationship between these two variables. This tends to confirm the suggestion by Buti and Sapir (1998) that the presence of wage convergence can be considered as proof of wage rigidity.

Behind this lack of flexibility, there is the institutional framework of the Spanish labour market, not only related to the wage-bargaining mechanism—sectoral and strongly influenced by national guidelines—but also by factors such as the tough hiring-and-firing rules, the generosity of unemployment benefits and the high level of minimum wage floors. This is in accordance with what theoretical models predict: that labour market institutions, through strict regulations, tend to compress wage differentials—especially if these institutions are more effective for unskilled workers—and then to reduce wage flexibility.

Finally, it is convenient to note that the lack of wage flexibility might be rather harmful for the performance of provincial labour markets in Spain. This is so because, as indicated in the introductory remarks, provincial wage flexibility is necessary to cope with issues such as the persistence of large differences in provincial unemployment rates and potential supply shocks related to the most recent EU enlargements. However, according to our results, it appears unlikely that the Spanish provinces will be able to respond to these problems. This being the case, it seems that—in order to act as an effective adjustment instrument—the Spanish wage-setting mechanism should be more flexible at the provincial level; considering what was mentioned in the previous paragraph, we think that Spain should move towards a much more decentralised system in which wage claims have an absolute direct effect on competitiveness. Thus, this paper calls for an institutional reform of the Spanish labour market, as it happens that the previous ones have only served to entrench a three-tier bargaining system. Indeed, if this reform were carried out, it would be possible to attain, at least in the long run, a higher degree of wage flexibility together with a provincial wage convergence process.

Notes

1. Other factors that can contribute to explaining territorial disparities are related to differences in the sources of economic growth such as factor endowments, technology and human capital (see, for instance, Nelson, 2005).
2. Blanchard and Wolfers (2000) and Bertola *et al.* (2002) have highlighted the role of institutions as potential explanatory factors of the difference in response of labour markets to shocks.
3. This is due to the fact that, as a general rule, the likelihood of suffering shocks with asymmetric results is greater the smaller the economy is.
4. As is mentioned later on, we refer to such factors as, for instance, unemployment benefits, minimum wages and firing costs.
5. From now on we use the term 'wages' to refer to 'real wages'.
6. Note: $i = 1$ = illiterate; $i = 2$ = no formal education or primary education; $i = 3$ = compulsory secondary education; $i = 4$ = pre-university education; $i = 5$ = higher education.
7. This is calculated according to the following expression

$$\varphi = (1- e^{-\beta T})/T$$

where T is the number of years of the period under analysis.
8. Defining this number of years by h, this can easily be calculated using the expression

$$e^{-\beta h} = 1/2$$

9. We are considering six sectors: primary sector, energy, industry, construction, private services and public services.
10. Other recent papers using spatial econometric techniques are, for example, Rey and Montouri (1999), López-Bazo *et al.* (1999), Le Gallo and Ertur (2003), Fingleton *et al.* (2004), Maza and Villaverde (2004), Villaverde (2005) and Ertur *et al.* (2006).
11. This is expressed as follows

$$I = \frac{n}{\sum_i \sum_j w_{i,j}} \frac{\sum_i \sum_j w_{i,j}(\omega_i - \omega)(\omega_j - \omega)}{\sum_i (\omega_i - \omega)^2}$$

where, ω_i (ω_j) is the real wage of province i (j), ω the mean real wage of the country; and $w_{i,j}$ an element of the distance matrix (W) between each pair of provinces.

In order to compute the significance level of Moran's I statistic, we have followed Anselin (1992) and assumed that the standardised statistic follows a normal distribution; for the sake of robustness, we have also used two other approaches (the randomisation and permutation approaches) and the results are roughly the same.

12. Tests that require the normality assumption in the residuals to be satisfied. In this respect, the results obtained from the Bera–Jarque test are satisfactory.
13. For instance, rearranging equation (2) we obtain the following equation, in which the third and fourth terms on the right-hand side refer to the aforementioned spatial effects (Toral, 2002):

$$\frac{1}{T} Log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right) = \alpha + \beta\, Log(\omega_{i,85}) + \rho W \frac{1}{T} Log\left(\frac{\omega_{i,03}}{\omega_{i,85}}\right) + \psi W \frac{1}{T} Log(\omega_{i,85}) + \tau$$

where, $\rho = \pi$, $\psi = -\pi\beta$

14. Spatial dependence invalidates the traditional ordinary least squares estimation method because of the two-dimensional and multi-directional nature of dependence in space. Specifically, the estimation of the spatial error model by ordinary least squares leads to inefficient parameters due to the non-diagonal structure of the disturbance variance matrix (Anselin, 1988). Likewise, it is convenient to point out that, according to the tests carried out, there are no problems of heteroscedasticity in this model.
15. The R^2 is not the appropriate measure to compare them, since it does not have the same meaning in the two cases.
16. Besides the filtering approach used in this paper, the literature considers alternative spatial conditioning schemes. Among these, another possibility to be used here is the 'neighbouring regions' approach applied, among others, by Quah (1996), Le Gallo (2004) and Tortosa-Ausina *et al.* (2005). The use of this approach would imply the construction of a new wage series in which each province wage is normalised by the average wage of the neighbouring provinces. Although both approaches are suitable for our purposes, we have opted for the filtering method as it seems to be more general: in fact, in the filtering method each value of the new wage series is constructed taking into consideration all other values in the sample while in the 'neighbouring' approach it only considers those of neighbouring provinces.
17. This bandwidth (h_0) has the following form

$$h_0 = 1.06\hat{\sigma} n^{-1/5}$$

where, n is the sample size, and $\hat{\sigma}$ is the standard deviation of provincial wages in Figures 6–9 and provincial productivity growth in Figure 10.
18. The scale of Figures 6 and 7 is the same to ease comparison.
19. Even so, we have checked for the presence of spatial autocorrelation in the residuals of equation (6) and found there is no sign of it. In addition, the computed value of the Durbin–Watson statistic (2.03) shows that there is no autocorrelation on the disturbances of the model. Finally, we have also tested for endogeneity problems between the provincial productivity growth and wage growth variables and found them not present in equation (6).
20. Although it is not shown in the paper, we have also computed the value of γ_1 for each province and, even though there are some differences between them, we can assert that national wage growth has a strong impact on provincial wage growth in almost all Spanish provinces. The results are available upon request.
21. When the same analysis of wage flexibility is carried out at sectoral level, it is shown that both the industrial sector and—not surprisingly, according to what was said in relation to the wage-setting mechanism—the public sector are the most flexible. This means that provinces in which these sectors are more strongly represented are also more flexible. Most of these provinces (Álava, Navarra, Guipúzcoa, Zaragoza ...) are located in the north of Spain, where the unemployment rates tend to be lower than average.

22. This non-linear relationship is demonstrated by the fact that a simple neglected non-linearity test (conditioned on regional productivity growth) rejects the null hypothesis of no neglected non-linearity at any standard level of significance. A Fan–Ullah (1999) test has been utilised in which the conditional expectation of the residuals took the form

$$E(\hat{\xi}_{it} | \dot{\lambda}_{it}) = m(\dot{\lambda}_{it})$$

where, m(.) was estimated using a Nadaraya–Watson estimator. The Fan–Ullah (t-test) statistic is 8.28, which clearly surpasses the 5 per cent critical value of 1.96.

Acknowledgements

The authors would like to thank four anonymous referees for their useful comments and suggestions.

References

Abraham, F. (1996) Regional adjustment and wage flexibility in the European Union, *Regional Science and Urban Economics*, 26, pp. 51–75.

Anselin, L. (1988) *Spatial Econometrics: Methods and Models*. Dordrecht: Kluwer.

Anselin, L. (1992) *SpaceStat tutorial: a workbook for using SpaceStat in the analysis of spatial data*. University of Illinois.

Armstrong, H. and Taylor, J. (2000) *Regional Economics and Policy*. Oxford: Blackwell.

Bande, R., Fernández, M. and Montuenga, V. M. (2001) Regional disparities in the unemployment rate: the role of the wage setting mechanism, *Regional Studies*, 41, pp. 235–251.

Bajo, O., Rabadan, I. and Salas, R. (1998) Regional wage flexibility in Spain, 1989–1992, *Regional Science Review*, 18, pp. 63–73.

Barro, R. and Sala-i-Martín, X. (2004) *Economic Growth*. Cambridge, MA: The MIT Press.

Bertola, J., Blau, F. and Kahn, L. M. (2002) Comparative analysis of labor market outcomes: lessons for the US from international long-run evidence, in: A. Krieger and R. Solow (Eds) *The Roaring Nineties: Can Full Employment Be Sustained?*, pp. 159–218. New York: Russell Sage and Century Foundations.

Blanchard, O. and Wolfers, J. (2000) The role of shocks and institutions in the rise of European unemployment: the aggregate evidence, *Economic Journal*, 110(462), pp. 1–33.

Blanchflower, D. G. (2001) Unemployment, well-being, and wage curves in eastern and central Europe, *Journal of Japanese and International Economics*, 15, pp. 364–402.

Buti, M. and Sapir, A. (1998) *Economic Policy in EMU: A Study by the European Commission Services*. Oxford: Oxford University Press.

Calmfors, L. and Drifill, J. (1988) Bargaining structure, corporatism and macroeconomic performance, *Economic Policy*, 6, pp. 13–61.

Decressin, J. and Fatás, A. (1995) Regional labor market dynamics in Europe, *European Economic Review*, 39, pp. 1627–1655.

Durlauf, S. N. and Quah, D. (1999) The new empirics of economic growth, in: J. B. Taylor and M. Woolford (Eds) *Handbook of Macroeconomics, Vol. 1*, pp. 231–304. Amsterdam: North-Holland.

Emerson, M., Gros, D., Italianer, A. *et al.* (1992) *One Market, One Money*. Oxford: Oxford University Press.

Ertur, C., Le Gallo, J. and Baumont, C. (2006) The European regional convergence process 1980–1995: do spatial regimes and spatial dependence matter?, *International Regional Science Review*, 29(1), pp. 3–34.

Ezcurra, R., Gil, C., Pascual, P. and Rapún, M. (2005) Inequality, polarisation and regional mobility in the European Union, *Urban Studies*, 42(7), pp. 1057–1076.

Fan, Y. and Ullah, A. (1999) Asymptotic normality of a combined regression estimator, *Journal of Multivariate Analysis*, 71, pp. 191–240.

Fingleton, B., Igliori, D. C. and Moore, B. (2004) Employment growth of small high-technology firms and the role of horizontal clustering: evidence from computing services and R&D in Great Britain, 1991–2000, *Urban Studies*, 41(4), pp. 773–799.

García, J. I. and Jimeno, J. F. (2005) *Public sector wage gaps in Spanish regions*. Working Paper No. 0526, Bank of Spain.

Getis, A. (1995) Spatial filtering in a regression framework: examples using data on urban crime, regional inequality, and government expenditures, in: A. Anselin and R. J. Florax

(Eds) *New Directions in Spatial Econometrics*, pp. 172–185. Berlin: Springer-Verlag.

Getis, A. and Ord, J. K. (1992) The analysis of spatial association by use of distance statistics, *Geographical Analysis*, 24, pp. 189–206.

Hardle, W., Müller, M., Sperlich, S. and Werwatz, A. (1999) *A course on non- and semiparametric modelling*. Humboldt University of Berlin.

Jackman, R., Layard, R. and Nickell, S. (1999) Combating unemployment: is flexibility enough?, in: R. Layard (Ed.) *Tackling Unemployment*, pp. 257–288. London: Macmillan.

Jimeno, J. F. and Bentolila, S. (1998) Regional unemployment persistence (Spain, 1976–1994), *Labour Economics*, 5, pp. 25–51.

Le Gallo, J. (2004) Space–time analysis of GDP disparities among European regions: a Markov chain approach, *International Regional Science Review*, 27, pp. 138–163.

Le Gallo, J. and Ertur, C. (2003) Exploratory spatial data analysis of the distribution of regional per capita GDP in Europe, 1980–1995, *Papers in Regional Science*, 82, pp. 175–201.

Li, Q. and Stengos, T. (1996) Semiparametric estimation of partially linear panel data models, *Journal of Econometrics*, 71, pp. 389–397.

López-Bazo, E., Vayá, E., Mora, A. and Suriñach, J. (1999) Regional economic dynamics and convergence in the European Union, *Annals of Regional Science*, 33, pp. 343–370.

Magrini, S. (1999) The evolution of income disparities among the regions of the European Union, *Regional Science and Urban Economics*, 29, pp. 257–281.

Martin, R. (2003) *A study on the factors of regional competitiveness*. Report for the European Commission Directorate General Regional Policy, Cambridge Econometrics and ECORYS-NEI.

Maza, A. and Moral-Arce, M. (2006) An analysis of wage flexibility: evidence from the Spanish regions, *Annals of Regional Science*, 40(3), pp. 621–637.

Maza, A. and Villaverde, J. (2004) Regional disparities in the EU: mobility and polarisation, *Applied Economics Letters*, 11(8), pp. 517–522.

Nelson, R. R. (2005) *The Sources of Economic Growth*. Cambridge, MA: Harvard University Press.

Quah, D. (1996) Regional convergence clusters in Europe, *European Economic Review*, 40(3–5), pp. 951–958.

Rey, S. J. and Montouri, B. D. (1999) US regional income convergence: a spatial econometric perspective, *Regional Studies*, 33(2), pp. 143–156.

Sala-i-Martín, X. (1996) The classical approach to convergence analysis, *Economic Journal*, 106(437), pp. 1019–1036.

Silverman, B. W. (1986) *Density Estimation for Statistics and Data Analysis*. London: Chapman and Hall.

Toral, A. (2002) Measuring distance in the Spanish regional convergence process, *Revue d'Économie Régionale et Urbaine*, 5, pp. 789–804.

Tortosa-Ausina, E., Pérez, F., Mas, M. and Goerlich, F. J. (2005) Growth and convergence profiles in the Spanish provinces (1965–1997), *Journal of Regional Science*, 45 (1), pp. 147–182.

Villaverde, J. (2005) Provincial convergence in Spain: a spatial econometric approach, *Applied Economic Letters*, 12, pp. 697–700.

Villaverde, J. (2006) Indicators of real economic convergence, in: P. de Lombaerde (Ed.) *Assessment and Measurement of Regional Integration*, pp. 146–161. London: Routledge.