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MISMATCH AND INTERNAL MIGRATION IN SPAIN, 1962-1986

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(*) Prepared for the volume *Mismatch and Labor Mobility*, edited by Fiorella Padoa Schioppa, to be published by the Centre for Economic Policy Research.

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1. Introduction

High European equilibrium unemployment is the most important economic development in the last fifteen years. Macroeconomists have put a lot of effort into trying to explain it. The studies made for the first Chelwood Gate conference, collected in the book *The Rise in Unemployment*, analysed the roles of real wages and demand contraction in the increase of unemployment after 1973, but then the persistence of high unemployment over time became the main issue. This task was picked up by the second Chelwood Gate conference, which studied the role of capital constraints and insider wage setting in generating persistence. Among the common findings across countries in the latter conference, as summarised by Drèze (1990), was that employment was consistently and significantly below the minimum of labour supply, classical and keynesian employment.

Deficient matching between labour supply and demand became a natural suspect for this finding, and so the first steps are being taken, for example in several chapters of this book, in developing a theory of mismatch. At the same time, it is useful to provide empirical evidence on the various dimensions of mismatch, at the very least to find out which are more relevant for each country. This paper is devoted to the latter task for the case of Spain.

Spain had very low unemployment, around 1%, in the 1960s but then experienced a sustained increase in the 1970s and the first half of the 1980s, reaching a 21.5% rate in 1985, by far the highest in the OECD. Since then the unemployment rate has drifted downwards, reaching 17.2% in 1989. Thus, unfortunately, Spain is an interesting country to study the persistence of unemployment and of mismatch as a potential cause for it.

Bentolila and Blanchard (1990) analyse the causes of the rise in unemployment in Spain and argue, in line with *insider-outsider* theories, that one of these causes is that high unemployment induces changes in the labour market through which increases in actual unemployment can lead to rises in equilibrium unemployment, i.e., the so-called *hysteresis* effect. In particular, they argue that the prolonged period of high unemployment in Spain has contributed to a reduction of the search intensity of the unemployed, and in particular of the willingness to look for work in regions other than their own, i.e. a source of mismatch from the labour supply side.

In order to learn about the likelihood of mismatch being an important cause for the rise in Spanish unemployment, we follow an eclectic approach, approaching the issue from several angles. One view, Jackman, Layard and Savouri's (1990) in one of the chapters in this volume, equates mismatch to relative unemployment rate dispersion. In the first part of Section 2 we document that, in fact, unemployment rate imbalances in absolute terms have greatly widened in Spain as the national rate was rising. Nevertheless, when we compute a relative unemployment rate dispersion measure we find that it has fallen over time according to most characteristics of the labour force, seemingly implying that mismatch has been falling, not rising, over time.

In the second part of Section 2 we pursue a different strategy. Disequilibrium models interpret mismatch as arising from heterogeneity of regimes (classical, keynesian, repressed inflation) across different sectors in the economy. Thus, in line with the Chelwood Gate II conference approach, we take the estimated measure of mismatch derived from the disequilibrium model fitted to the Spanish economy, and regress it on variables related to mismatch, such as the proportion of long-term unemployed in the labour force, regional unemployment rate dispersion, interregional migration flows, relative energy prices, or employment turbulence, finding a very good fit. This suggests that this measure might be a good proxy for overall mismatch in Spain. Unfortunately, it behaves exactly in the opposite way that the dispersion indices commented on above do: it rises steadily until the mid-1980s, and then falls a little. It clearly implies that mismatch is today at historically high levels in Spain. At the end of Section 2 we list a number of reasons leading us to think that the latter index might be giving a more accurate view than the former ones.

Finally, the fact that regional variables play an key role in explaining the latter index of mismatch reinforces our previous belief that the geographical aspect of the labour market is important in understanding the rise in unemployment in Spain. There have always been genuine differences in language, uses, etc. across Spanish regions, and this could easily lead to the segmentation of markets. Also, migration flows, which in the early 1960s were high both towards Europe and within Spain, have fallen dramatically coinciding with the rise in unemployment. Bentolila and Blanchard (1990) stress the rise in overall unemployment as the main factor inhibiting labour mobility in Spain. In Section 3 we analyse this proposition by setting up an econometric model of internal migration in Spain. We find that net interregional migration flows respond to economic incentives, in particular to unemployment and wage differentials, but with low elasticities and long lags. Therefore the observed fall in these flows can be partially explained by the reductions in such differentials that has taken place in Spain. We also find that migration flows are deterred by housing price differentials[1] and by the overall unemployment rate. In consequence, while labour mobility will increase as the overall unemployment rate continues to fall in the next few years, this process is bound to be slow, and there is a role for policy in speeding up the process. We dwell on some such measures at the end of Section 3.

Finally, Section 4 draws some preliminary conclusions and opens issues for future research.

2. Stylized facts of Spanish unemployment and mismatch indices

2.1. Evolution over time

In this subsection we sketch an account of the evolution of Spanish unemployment over time, depicted in Figure 1. From 1960 to 1974 Spain made the transition from an agricultural to an industrial economy, and experienced high growth, low unemployment, moderate inflation and high productivity growth. In 1975 Franco died, and the two years it took to get political institutions back in order saw a wage explosion and the transmission of wage costs and the oil price shock into booming inflation (25% in 1977). The inflationary momentum was broken by contractionary monetary policy and a series of nationwide agreements on

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1 Echoing the theme of Muellbauer and Murphy (1989).

FIGURE 1: National Unemployment Rate

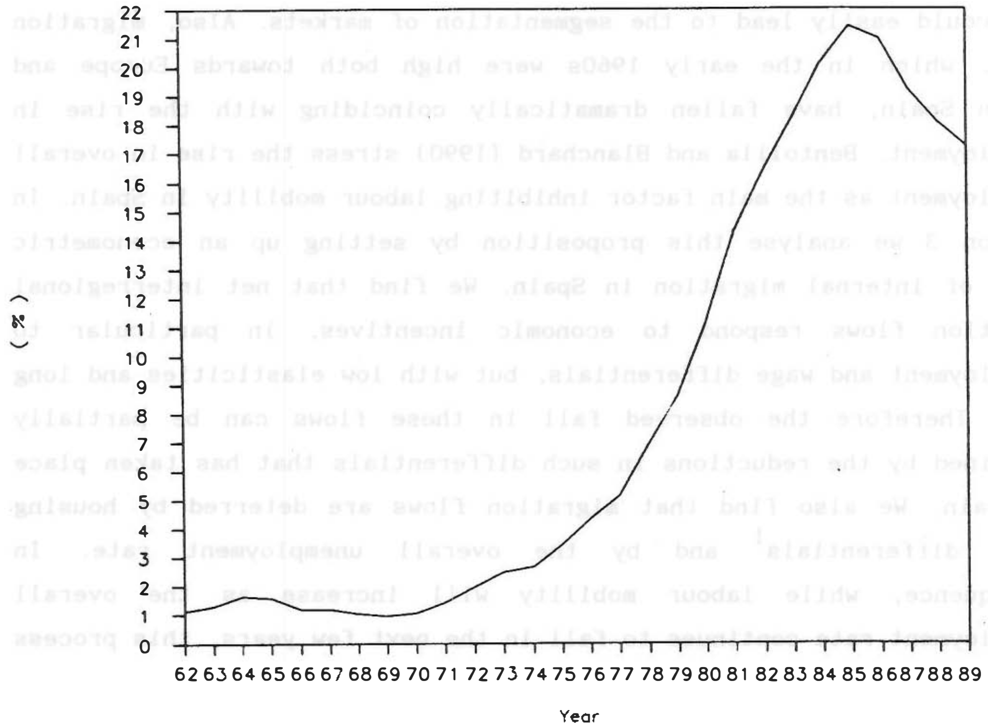
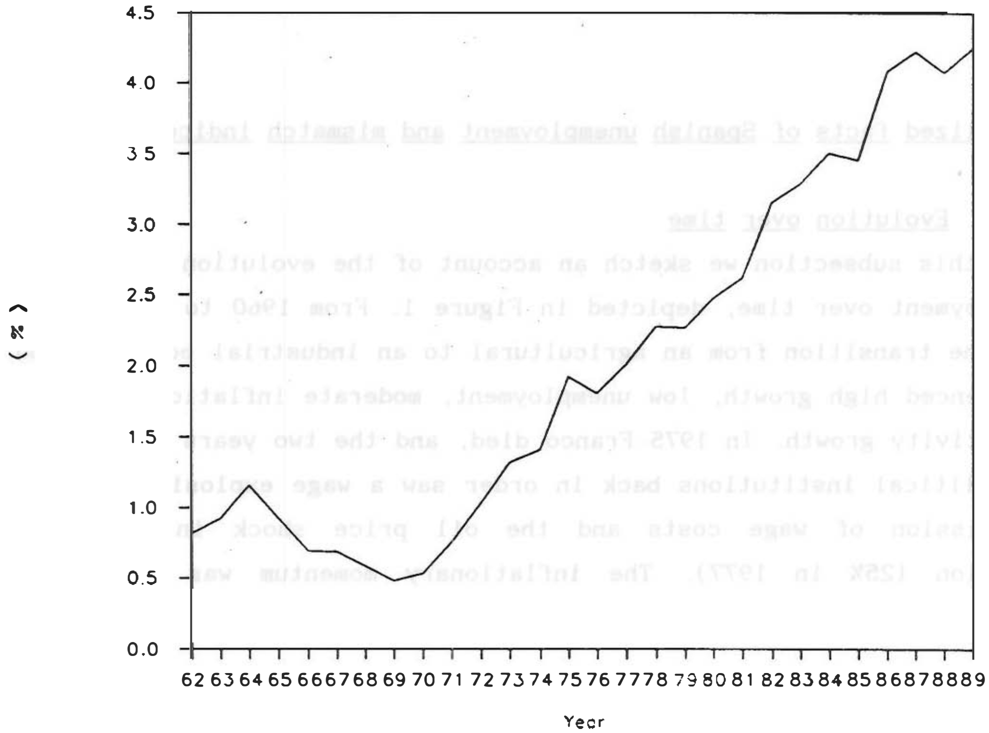


FIGURE 2: Regional Unemp. Ineq. Index



wage growth, whereby wage moderation was progressively attained.

The reduction of inflation implied the cost of a sharp slowdown and a large increase in unemployment, which reached a staggering 21.5% in 1985. The causes of the increase have been extensively analysed (see Dolado *et al.* (1986), Fina (1987), Andrés *et al.* (1990) or Bentolila and Blanchard, (1990)). The initial rise in unemployment can be attributed to the large increases in real wages and the contractionary monetary policies that ensued. There is much less consensus about the continued rise in the 1980s. Bentolila and Blanchard (1990) single out three factors in explaining such rise: a profit squeeze which led many firms to bankruptcy and the remaining ones to curtail investment, a productive reorganisation effort which caused massive labour shedding, and hysteresis effects, by which equilibrium unemployment rose with actual unemployment. In this paper we are specially concerned with the latter.

The literature on hysteresis, which originally stressed the lack of concern of *insider* wage setters for the interests of unemployed *outsiders*, has recently shifted emphasis towards the determinants of the pressure from the unemployed on the wage setting process (Layard and Nickell (1987)). Hysteresis arises if depressed labour markets lead to less downward wage pressure from outsiders. This may happen if, for instance, the long-term unemployed lose skills or get discouraged and stop searching.

As in most European countries, equilibrium unemployment -i.e. the rate compatible with steady inflation or NAIRU- has risen in Spain, although not as much as actual unemployment, since inflation has fallen after 1977; and it is a common finding that unemployment puts little downward pressure on wages in Spain (e.g. Dolado and Malo de Molina (1985), Dolado *et al.* (1986), or Andrés and García (1989)). Moreover, in 1986-88, a boom period in which output and employment grew respectively by 4.7% and 2.7% annually, unemployment fell only by two percentage points, from 21.5% in 1985 to 19.5% in 1988. In contrast, during 1989, unemployment has fallen more quickly, to 17.2%, but inflation has accelerated by two percentage points, to 6.8%, an overheating typically arising near equilibrium unemployment.

The long period of high unemployment in Spain has likely induced changes in the behaviour of the unemployed, towards reducing their search effort and so decreasing downward pressure on wage bargains. There are several features consistent with this: the stigma of being unemployed is mostly gone, the pool of the unemployed has been relatively stagnant until recently, unemployment falls mostly on spouses and the youth -who can survive while being out of work- and, finally, migration flows, which were high in the early 1960s, are now quite small.

How much of the rise in equilibrium unemployment can be attributed to mismatch? We lack a fully worked out model of the relationship between the NAIRU and mismatch, and do not develop one in this paper. Instead, we follow an eclectic approach, by computing two sets of measures which, under two different models, proxy for mismatch. If we were to get similar answers from both measures, we would the feel confident to have found a robust stylised fact. The two models we are referring to are the Jackman-Layard-Savouri (JLS) multisectoral model of determination of the NAIRU, and the Sneesens-Drèze disequilibrium model. In the next two subsections we take each in turn.

2.2. Mismatch as unemployment rate dispersion

In their chapter, JLS derive a relation between mismatch and the NAIRU in a multisectoral model, using two building blocks: the factor price frontier arising from a Cobb-Douglas production function, and a double-logarithmic wage function whereby sectoral wages depend on their own unemployment rate. The combination of both elements gives rise to an unemployment frontier, i.e. the locus of all combinations of sectoral unemployment rates consistent with the absence of inflationary pressure. In this setup, the movement of unemployed workers from relatively high to relatively low unemployment rate sectors can reduce the national unemployment rate without raising inflation. As a consequence, unemployment rate dispersion is a sign of mismatch, a natural measure of the latter turning out to be (half) the squared coefficient of variation of sectoral unemployment rates, i.e.:

$$MM = \frac{1}{2} \text{Var} (u_i/u_N) \qquad [1]$$

where u_i denotes the unemployment rate of sector (group) i and u_N the national rate (hereafter the subindex N denotes the national value of a variable).

In the application of this model, before reporting computed MM indices, we document that, in fact, important unemployment rate imbalances have developed in Spain as unemployment was rising. We present, in Table 1, the breakdown of unemployment by age, sex and skill in (the fourth quarter of) three selected years. Its composition has shifted towards the 20-to-54 year old group, and by 1989 more than half of the unemployed were in their prime age. Still, since the latter is the largest group in the labour force, their unemployment rate is always much lower than the youngest workers' rate, which reached 55% in 1985. Turning now to the sex composition of unemployment, Women have gone from being a third to a half of the unemployed, with their unemployment rate almost doubling the male rate by 1989.

Table 2 shows the regional structure of unemployment. Unemployment rate divergence has greatly increased over time, as evidenced by Figure 2, which plots the sum of the absolute differences in unemployment rates across regions, weighted by their labour force share.

Do these imbalances translate into growing mismatch, as measured by the MM index? Table 3 presents, and Figure 3 plots, the index from 1977 to 1989,[2] by sex, age, education, skill and sector.[3] The age, sector and education dimensions all show decreasing values over time; the sex and skill indices are stable up to 1985 and increase afterwards. The latter two are, therefore, the only ones consistent with the widespread perception of increasing mismatch in Spain in the last few years. The fall in the former three indices reveals that, although absolute differences in unemployment rates have increased over time, relative

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[2] The first complete year for which we have homogeneous data is 1977.

[3] Not all characteristics are equally interesting: change of sex is (almost) impossible, change of age is exogenous, and the sectoral classification is not very informative, since workers do not necessarily find jobs in the same sector where they worked last.

TABLE 1: COMPOSITION OF UNEMPLOYMENT

(Fourth quarter, %)

	Percentage structure			Unemployment rate		
	1976	1985	1989	1976	1985	1989
A) AGE:						
16-19 years old	31.5	19.8	12.8	17.1	54.9	36.6
20-24 years old	19.6	28.3	27.7	10.6	42.5	32.3
25-54 years old	43.0	46.6	52.8	3.6	15.7	13.5
55 years old +	5.9	5.4	6.7	2.1	8.0	8.3
Total	100.0	100.0	100.0	4.9	21.2	16.9
B) SEX:						
Male	65.7	59.2	49.2	5.2	18.6	12.7
Female	34.3	40.8	50.8	6.8	27.6	24.8
Total	100.0	100.0	100.0	4.9	21.2	18.5
C) SKILL:						
Professional and managerial	1.9	2.4	3.1	1.2	6.0	4.9
Clerical	4.8	5.2	6.3	2.4	11.0	9.3
Other non manual	3.7	4.6	4.3	1.9	10.3	7.5
Unskilled	89.6	87.8	86.2	6.0	26.1	21.5
Total	100.0	100.0	100.0	4.9	21.2	16.9

Sources: See Appendix 2.

TABLE 2: REGIONAL UNEMPLOYMENT RATES

(annual averages, %)

	1962	1976	1985	1989
Andalucía	3.4	9.9	29.2	27.0
Aragón	0.2	2.5	17.2	12.1
Asturias	0.3	3.4	18.0	17.8
Baleares	0.4	3.9	13.5	10.7
Canarias	1.1	8.3	25.7	21.5
Cantabria	0.5	2.9	15.5	17.8
Castilla-La Mancha	0.5	4.5	15.5	14.1
Castilla-León	0.3	4.2	17.6	16.7
Cataluña	0.9	3.4	21.7	14.3
País Vasco	0.2	3.7	22.7	19.6
Extremadura	1.7	8.0	27.0	26.4
Galicia	0.4	1.7	13.1	12.1
Madrid	1.1	5.1	21.1	13.2
Murcia	1.4	4.8	18.9	16.2
Navarra	0.1	3.9	18.8	12.8
La Rioja	0.3	1.8	16.5	10.1
Valencia	1.2	3.6	19.9	15.4
National average	1.1	4.3	21.4	17.2

Sources: See Appendix 2.

TABLE 3: MM MISMATCH INDICES (%)

	Sex	Age	Education	Skill	Sector
1977	0.6	28.5	2.8	7.1	27.2
1978	1.3	32.6	4.9	6.6	24.8
1979	1.1	31.4	6.4	6.9	25.8
1980	1.0	30.9	5.1	7.0	24.2
1981	1.1	29.5	5.7	6.8	21.7
1982	1.4	27.2	5.8	6.5	18.1
1983	1.4	25.3	5.7	6.4	16.1
1984	1.4	23.3	4.9	7.0	15.7
1985	1.6	20.3	4.3	6.7	11.7
1986	2.2	19.2	3.5	7.6	8.4
1987	3.2	18.3	3.4	8.8	6.4
1988	4.6	25.0	3.5	8.6	2.9
1989	5.8	25.0	2.3	8.7	2.5

Sources: See Appendix 2.

Note: The level of disaggregation is: (a) Age, 3 groups (16 to 24, 25 to 54 and 55 or more years old); (b) Education, 5 groups (illiterate, primary school, high school, vocational training and university); (c) Skill, 4 groups (as in Table 1); (d) Sector, 4 groups (agriculture and fishing, manufacturing, construction and services.

FIGURE 3: Mismatch indices

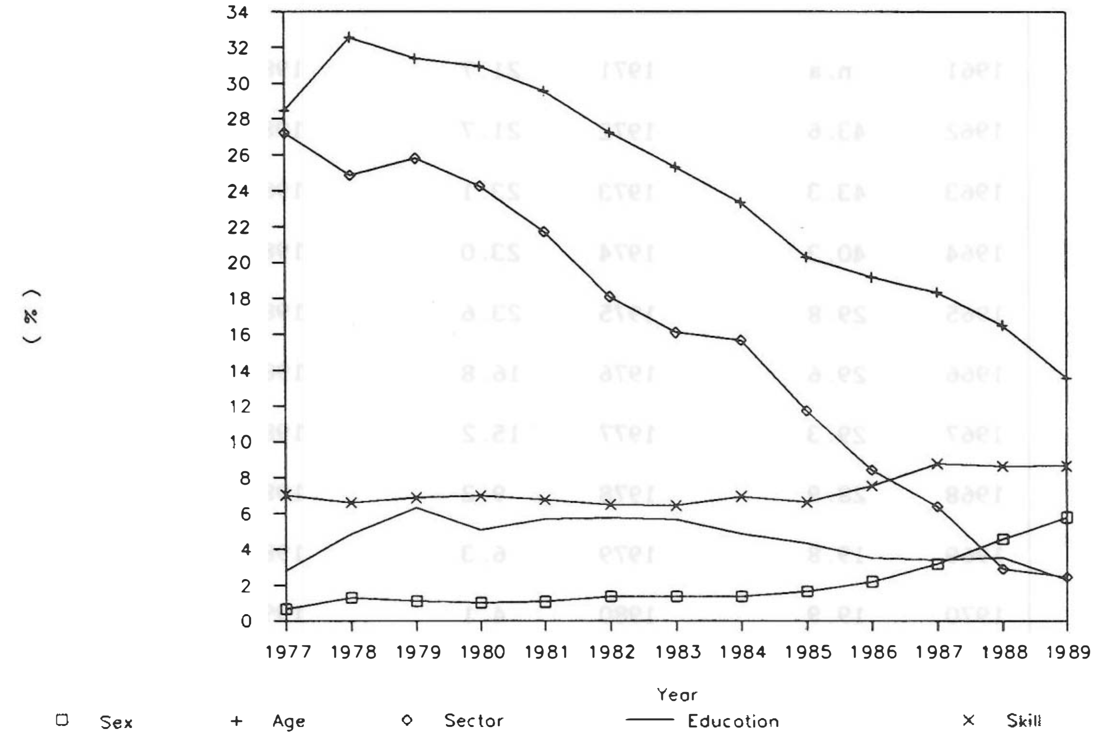
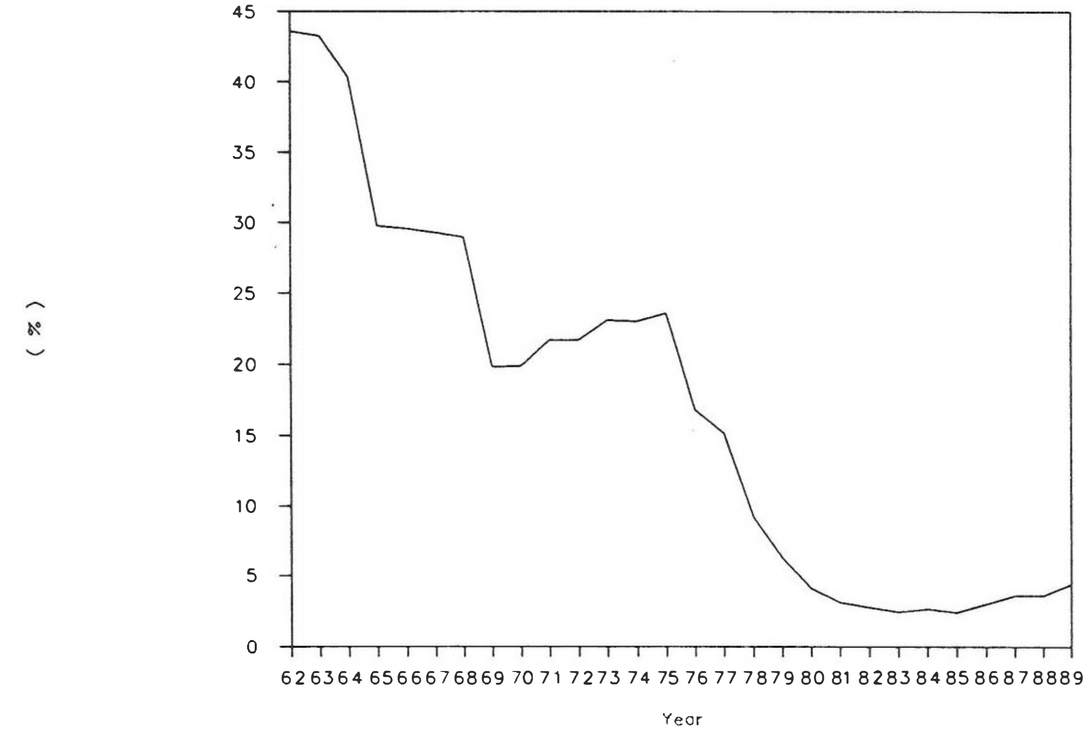


FIGURE 4: Regional mismatch index



differences have fallen. For instance, in 1976Q4 the difference between the prime-age and the 16-to-19 year old unemployment rates was 15 percentage points, rising to 23 p.p. by 1989Q4. However the former rate quadrupled from 1976 to 1989, while the latter *only* doubled.

The case of regions, for which we compute a series starting in 1962 and report the results in Table 4 and Figure 4, is quite striking. In 1962, when national unemployment was very low, geographical unemployment dispersion was very high; but it fell dramatically afterwards, bottoming out in 1985, when national unemployment reached its maximum. The reason is the same as before: absolute differences in unemployment rates increased substantially, but the denominator of the MM index, the national rate, rose so much more than the numerator that it dwarfed the latter's increase. Since 1985 regional dispersion has been increasing again, as the national rate was falling.

2.3. Mismatch as micro-market constraint heterogeneity

A different measure of mismatch results from the model common to all papers in the aforementioned Chelwood Gate II conference. The model is explained in detail by Drèze (1990), so here we only review its main features. It is a disequilibrium model, in which rationing arises from wage and price stickiness as well as from short-run rigidity of technical coefficients of production. As is usual in these models, employment can be constrained by lack of demand, i.e. the keynesian regime, lack of productive capacity, i.e. the classical regime, or lack of labour, i.e. the repressed inflation regime. If every sector in the economy were constrained in an identical way, aggregate employment would equal the minimum of the three employment levels above. Such a situation being highly unlikely, aggregation is performed in the model allowing for heterogeneity of constraints. In particular, aggregation is done via a constant elasticity of substitution (CES) function whose parameter, $1/\rho$, turns out to measure the degree of disparity in the rationing regime of different sectors: the latter goes to zero as ρ goes to infinity. Mismatch is in this way identified with regime disparity across sectors, and is revealed by actual employment being lower than the minimum of labour supply, classical and keynesian employment.

The variable $1/\rho$ measures frictional (structural) unemployment. To understand why we need to recall the CES equation giving aggregate employment, $\tilde{L}$:

$$\tilde{L} = (LK^{-\rho} + LP^{-\rho} + LS^{-\rho})^{1/\rho} \qquad [2]$$

where LK and LP are Keynesian and classical employment and LS is labour supply. Now if LK=LP=LS, then $\tilde{L} = 3^{-\rho}$ LS, so that:

$$1/\rho = (\log 3)^{-1}\ \tilde{u} \qquad [3]$$

where $\tilde{u} = 1 - (\tilde{L}/LS)$ is the frictional unemployment rate, and we have used the approximation $\log(1+x) \simeq x$. This expression also serves to show that $1/\rho$ is in the same (rescaled) units as the unemployment rate.

Figure 5 shows $1/\hat{\rho}$ as estimated for Spain, in Andrés *et al.* (1990), by one of us and other colleagues. It increases slowly up to 1973, quickly from 1974 to 1985, and falls slowly afterwards; i.e. it follows a path similar to that of the unemployment rate. Can we account for the behaviour of $1/\hat{\rho}$ using economic variables related to mismatch? Andrés *et al.* regressed $\hat{\rho}$ on a trend (t), a proxy of mismatch (PMM), and (the logarithm of) an index of the relative price of energy inputs (LPRM). PMM is Layard and Nickell's (1986) turbulence index, i.e. the sum of absolute changes in sectoral shares of employment,[4] though Andrés *et al.* use dependent employment instead of total employment. This variable should capture the need for relocation of labour across sectors as the composition of employment changes, being therefore positively correlated with mismatch; LPRM should also increase mismatch, by changing the relative price of inputs.

Estimating the equation for $1/\hat{\rho}$, instead of $\hat{\rho}$, for 1965-87, we get: (t-ratios in parenthesis)

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[4] Formally: $PMM = \sum_i |\Delta(N_i/N)|$, where N_i is employment in sector i and N is total employment. The sectors are: agriculture, industry, construction and services.

FIGURE 5: Estimated mismatch

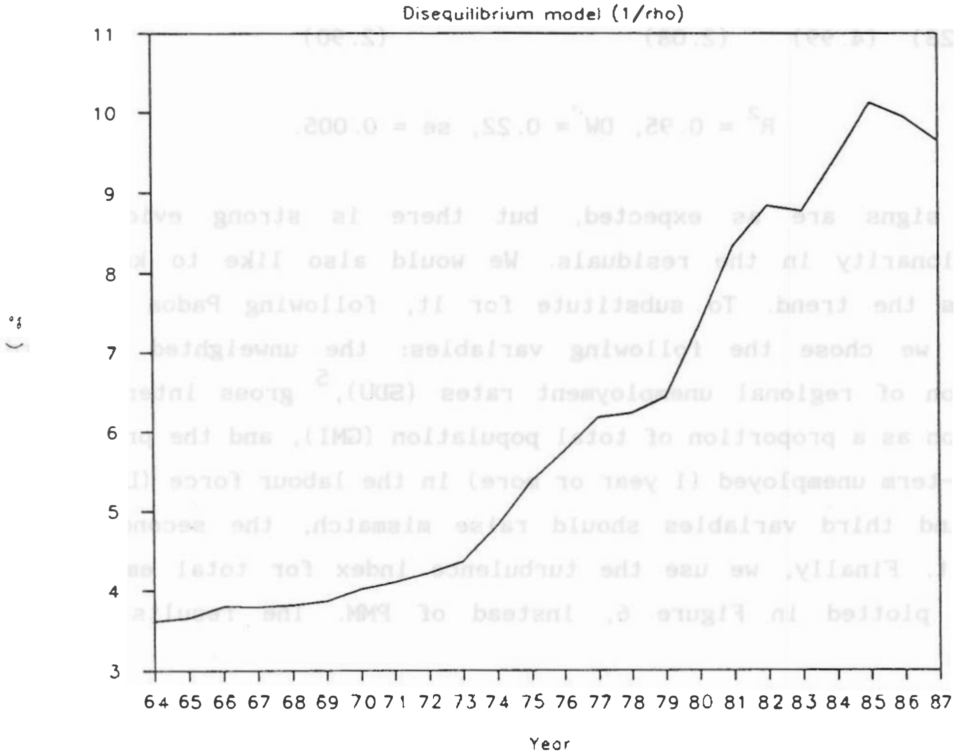
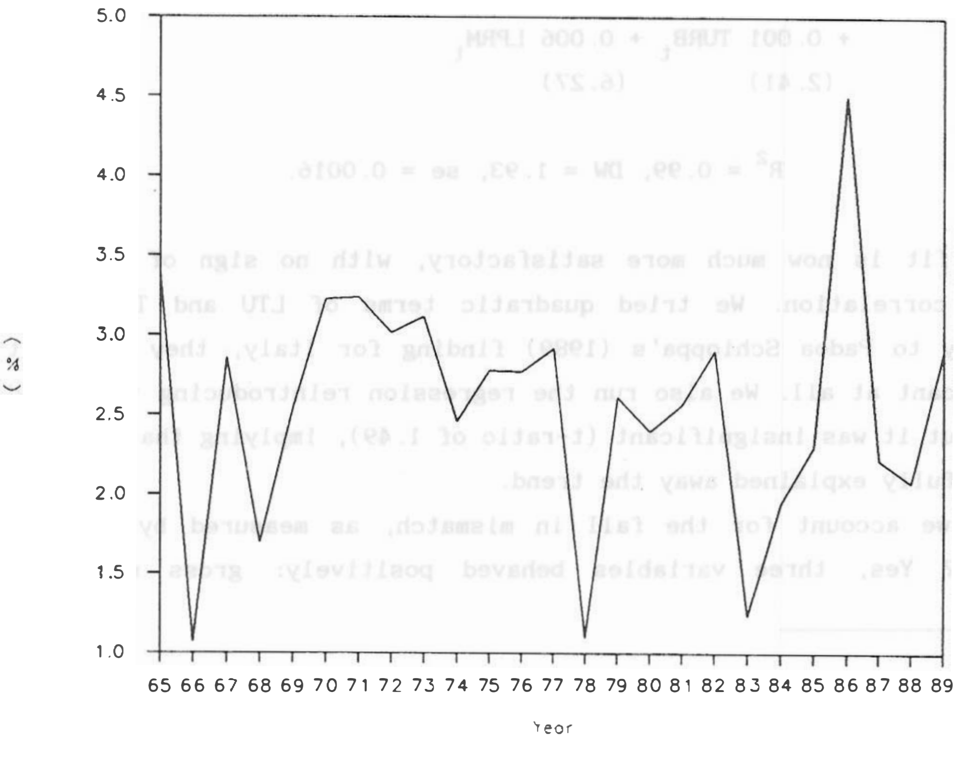


FIGURE 6: Turbulence



$$1/\hat{\rho}_t = \underset{(4.28)}{0.031} + \underset{(4.99)}{0.002}\ t + \underset{(2.08)}{0.059}\ (PMM_t + PMM_{t-1}) + \underset{(2.90)}{0.005}\ (LPRM_t + LPRM_{t-1}) \quad [4]$$

$$R^2 = 0.95,\ DW = 0.22,\ se = 0.005.$$

The signs are as expected, but there is strong evidence of nonstationarity in the residuals. We would also like to know what explains the trend. To substitute for it, following Padoa Schioppa (1989), we chose the following variables: the unweighted standard deviation of regional unemployment rates (SDU),[5] gross interregional migration as a proportion of total population (GMI), and the proportion of long-term unemployed (1 year or more) in the labour force (LTU). The first and third variables should raise mismatch, the second should lower it. Finally, we use the turbulence index for total employment (TURB), plotted in Figure 6, instead of PMM. The results are as follows:

$$1/\hat{\rho}_t = \underset{(20.80)}{0.054} + \underset{(7.30)}{0.007}\ SDU_{t-1} - \underset{(5.38)}{0.023}\ GMI_t + \underset{(8.72)}{0.002}\ LTU_t$$

$$+ \underset{(2.41)}{0.001}\ TURB_t + \underset{(6.27)}{0.006}\ LPRM_t \quad [5]$$

$$R^2 = 0.99,\ DW = 1.93,\ se = 0.0016.$$

The fit is now much more satisfactory, with no sign of residual serial correlation. We tried quadratic terms of LTU and TURB but, contrary to Padoa Schioppa's (1989) finding for Italy, they were not significant at all. We also run the regression reintroducing the trend term, but it was insignificant (t-ratio of 1.49), implying that we have successfully explained away the trend.

Can we account for the fall in mismatch, as measured by $1/\hat{\rho}$, in 1986-87? Yes, three variables behaved positively: gross migration

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5 SDU is the square root of JLS's MM index, but unweighted.

increased in those two years, the proportion of long-term unemployed fell, and so did the price of energy imports. They overcame the continued rise in regional unemployment rate dispersion and the 1986 turbulence blip.

This measure of mismatch does, unfortunately, give a different answer from JLS's MM index. The latter falls according to most characteristics of the labour force, while $1/\hat{\rho}$, which is an overall measure, steadily increases over time.

The units of measurement help explain the different behavior of the two indices. MM is a squared coefficient of variation, i.e. a (squared) standard deviation divided by a (squared) mean, it is therefore dimensionless. But $1/\rho$ is in the same units as the unemployment rate (see equation [3]), i.e. in the square root of the units of the numerator of MM. As we mentioned before, the latter does rise steadily over time (as does the index of absolute unemployment rate differences plotted in Figure 2), but its rise is dwarfed by that of the denominator, the national unemployment rate.

The matter thus boils down to the following question: is it only relative unemployment rate dispersion that matters or do absolute unemployment rates and differentials also provide independently valuable information about mismatch? To put it more bluntly: suppose that an economy has two types of labour with equal size labour forces and with respective unemployment rates of 2% and 4%. Is mismatch the same if the latter are 10% and 20%? The MM index says so, but this is hard to accept. Higher unemployment is usually associated with a higher proportion of long-term unemployment, for example, which is itself likely to produce mismatch. The catch is that mismatch may be both a cause for the persistence of high unemployment and a consequence of it. In summary, we think the numerator of MM to be interesting on its own.

In the case of Spain, two reasons induce us to believe that $1/\hat{\rho}$ might be giving a more accurate picture, in the sense that mismatch is nowadays in Spain at historically high levels. First, we find no evidence in Section 3 below of double-logarithmic wage (i.e., concave) equations at the regional level, a specification which is clearly rejected in favour of a semi-logarithmic form. Second, if mismatch had really been falling according to most dimensions (and dramatically in some of them), it would be very difficult to understand the overheating currently taking place in Spain.

We end this section by noting that two regional variables help explain $1/\hat{\rho}$; (unweighted) unemployment rate dispersion and migration flows. The latter have been falling steadily since the early 1960s, a fact which we would need to explain in order to improve our understanding of mismatch in Spain. Further insight requires, however, formal modelling of the role of unemployment rate disparities, and their interaction with wage setting, in the regional labour allocation process. We undertake this task in the following section.

3. Analysis of migration flows

In this section we analyse the behaviour of internal migration flows in Spain since 1962. We review a few stylised facts, set up a framework of analysis and comment on the time path of our explanatory variables. Finally, we report our estimation results and derive policy implications from them.

3.1. Stylized facts

Even though Spain is a relatively small country, there are genuine differences between regions in terms of weather, language, traditions, etc., which have played an important role in Spanish history. The traditional administrative division is into 50 provinces, but these have always been grouped into regions, whose limits have changed over time. The current structure, dating from 1978, consists of 17 so-called *Autonomous Communities*, which have their own parliaments and governments, with wide-ranging political powers.

Spanish population has increased from 31 million in 1962 to 38.5 million in 1986.[6] Migration flows were very high in Spain in the 1960s and early 1970s. Gross outflows to other, mainly European, countries

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[6] Due to lack of data for most variables after 1986, in the remainder of the paper we restrict ourselves to the period 1962-86.

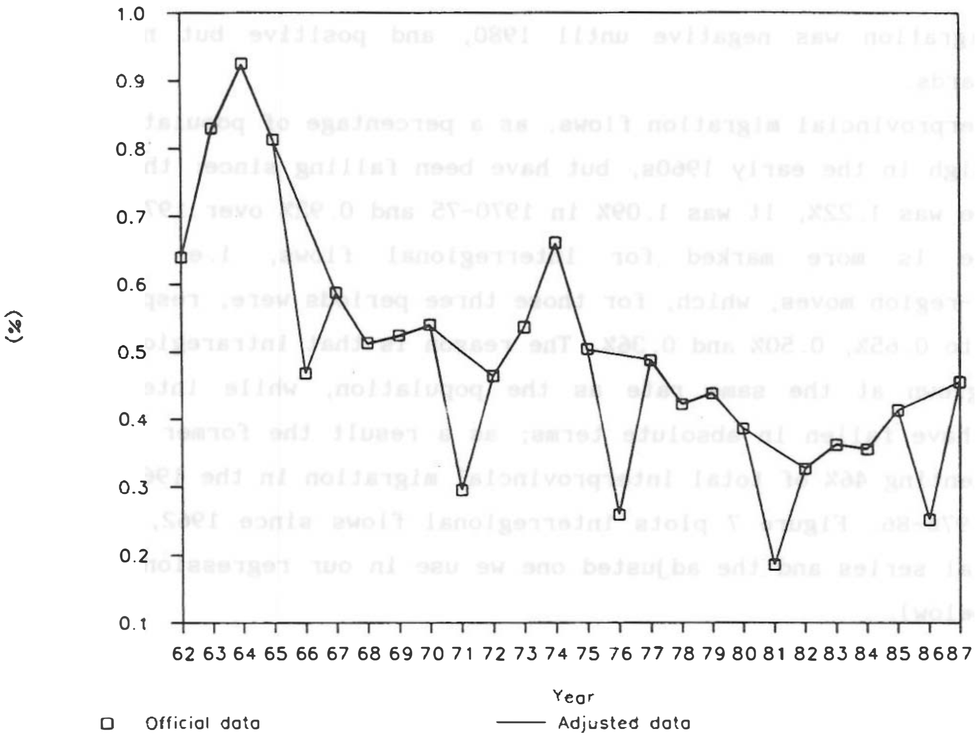
averaged 0.3% of the population in 1960-74. They then fell dramatically, to 0.06% over 1975-86, being overcome by returns, so that net migration was negative until 1980, and positive but negligible afterwards.

Interprovincial migration flows, as a percentage of population, were very high in the early 1960s, but have been falling since: the 1962-69 average was 1.22%, it was 1.09% in 1970-75 and 0.92% over 1976-86. The decline is more marked for interregional flows, i.e. excluding within-region moves, which, for those three periods were, respectively, equal to 0.65%, 0.50% and 0.36%. The reason is that intraregional flows have grown at the same rate as the population, while interregional flows have fallen in absolute terms; as a result the former went from representing 46% of total interprovincial migration in the 1960s to 61% over 1976-86. Figure 7 plots interregional flows since 1962, both the official series and the adjusted one we use in our regression analysis (see below).

Some information on the characteristics of migrants is available. As expected, young people -under 25 years old- are more likely to move: over 1962-86 they represented almost half of all migrants, but just to 42% of the population. In the 1960s the proportions of young people in both the population and migration were increasing, but the 1970s and 1980s show an ageing of both the population -42% were young in 1970, 40% in 1986- and migration -51% of migrants were young in 1970, only 46% in 1986-. In terms of sex, even though the participation of women in the population has increased, from 48.5% in 1960 to 49% in 1986, their share of migration has fallen, from 53% in 1962 to 51% in 1986. Finally, the available data on labour force status are quite useless. Most migrants (around 60%) are classified as *non-active*, but it is unclear if this only considers out-of-the-labour-force migrants or if some or all of the unemployed workers are counted in. Within those labelled as *active*, around three-quarters are manual workers in the 1960s, with the proportion declining in the 1970s and 1980s in favour of skilled non-manual workers, which amount to almost 40% of all migrants in 1986.

In order to highlight other features of migration, we have grouped, looking at per-capita income and gross migration flows, the 17

FIGURE 7: Interregional Migration Rate



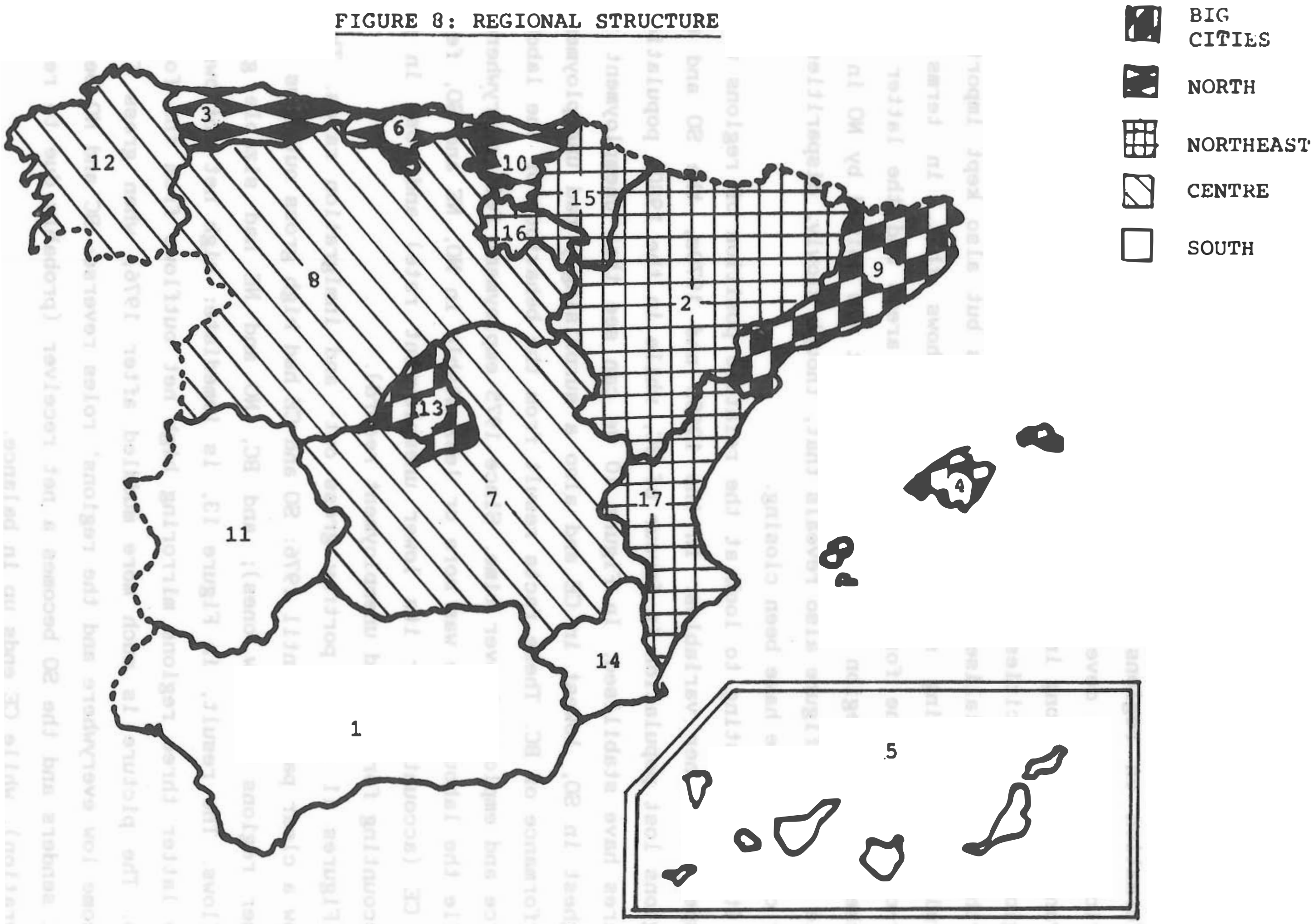
administrative regions into more aggregate regions (in the spirit of Attanasio and Padoa Schioppa in one of the chapters in this volume). We distinguish five super-regions: Big Cities (BC), North (NO), Northeast (NE), Centre (CE) and South (SO); Figure 8 and Appendix 1 provide information on regions and their grouping. The South is the most agricultural, less developed region, followed by the Centre. The most advanced, with strong industrial and service sectors, is the Big Cities region (the big cities are Madrid and Barcelona), followed by the North, which specialised in basic industries but also kept important farming and fishing sectors. Figure 9 shows that, in terms of per-capita GDP, the former are the poorest areas and the latter the richest. The NE region is in between, and it was joined by NO in the late 1970s. The figure also reveals that, though slowly, disparities in per capita income have been closing.

It is interesting to look at the relative position of regions with respect to other variables. In the 1960s and 1970s, the SO and CE regions lost population in favour of BC, while in the 1980s population shares have stabilised. In Figure 10 we can see that unemployment is highest in SO, lowest in CE and also a surprisingly bad unemployment performance of BC. These facts result from the behaviour of the labour force and employment over time. Since 1975 employment fell everywhere, while the labour force was more or less stable in NO, NE and SO, fell in CE (accounting for its lower unemployment rate) and rose in BC (accounting for its bad unemployment record).

Figures 11 and 12 portray gross out- and inmigration rates. They show a clear pattern until 1976: SO and CE had high gross outflows (the other regions had low ones); and BC, NO and NE had sizable gross inflows. The result, in Figure 13, is immediate: high net inflows to the latter three regions mirroring high net outflows from the former two. The picture is much more muddled after 1976, when gross flows become low everywhere and the regions' roles reverse: BC and NO become net senders and the SO becomes a net receiver (probably due to return migration), while CE ends up in balance.

Turning to wages and prices, the real (consumption) wages per employee follow a similar pattern as real per capita GDP (Figure 9), except that wages in NO remain closer to BC's than to NE wages, while

FIGURE 8: REGIONAL STRUCTURE



Note: Region numbers correspond to those in Appendix 1.

FIGURE 9: Per capita real GDP

(1980 million pesetas)

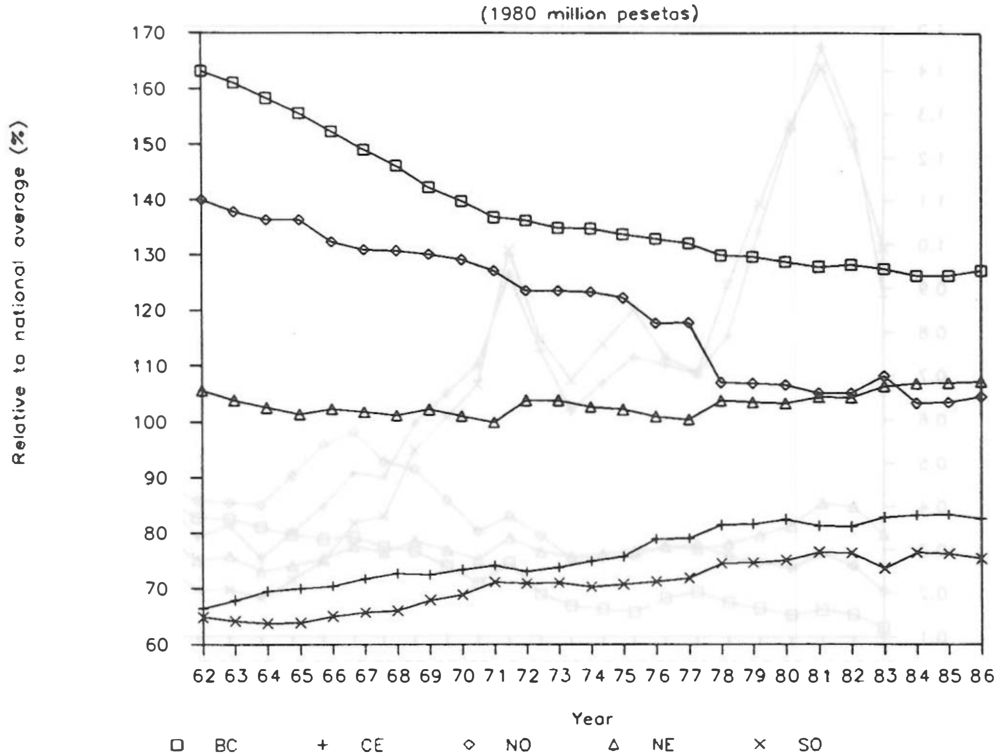


FIGURE 10: Regional Unemployment Rates

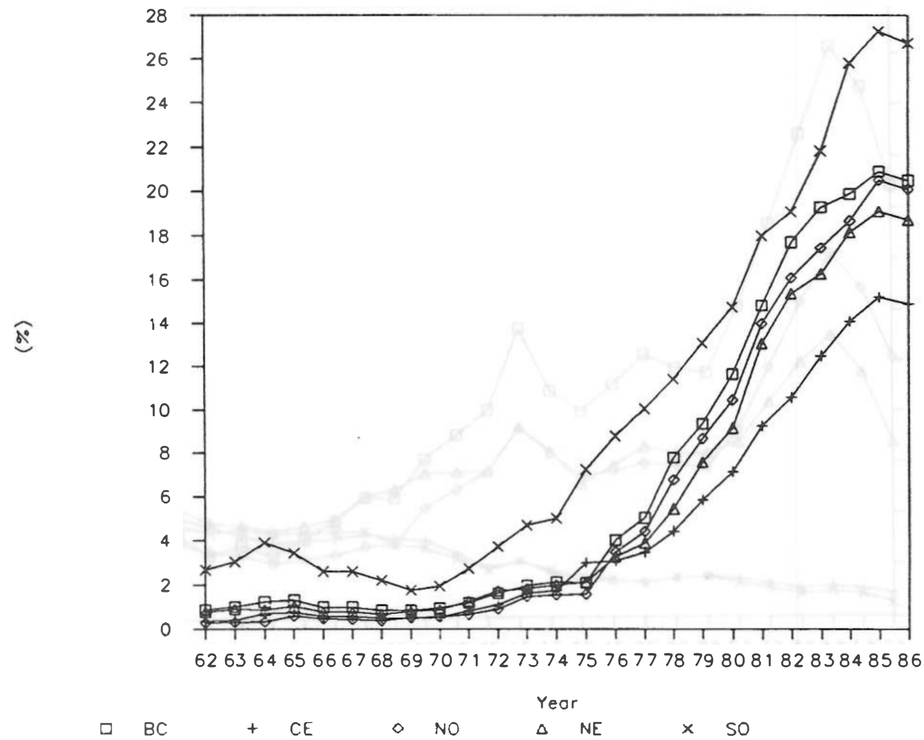


FIGURE 11: Gross Outmigration Rates

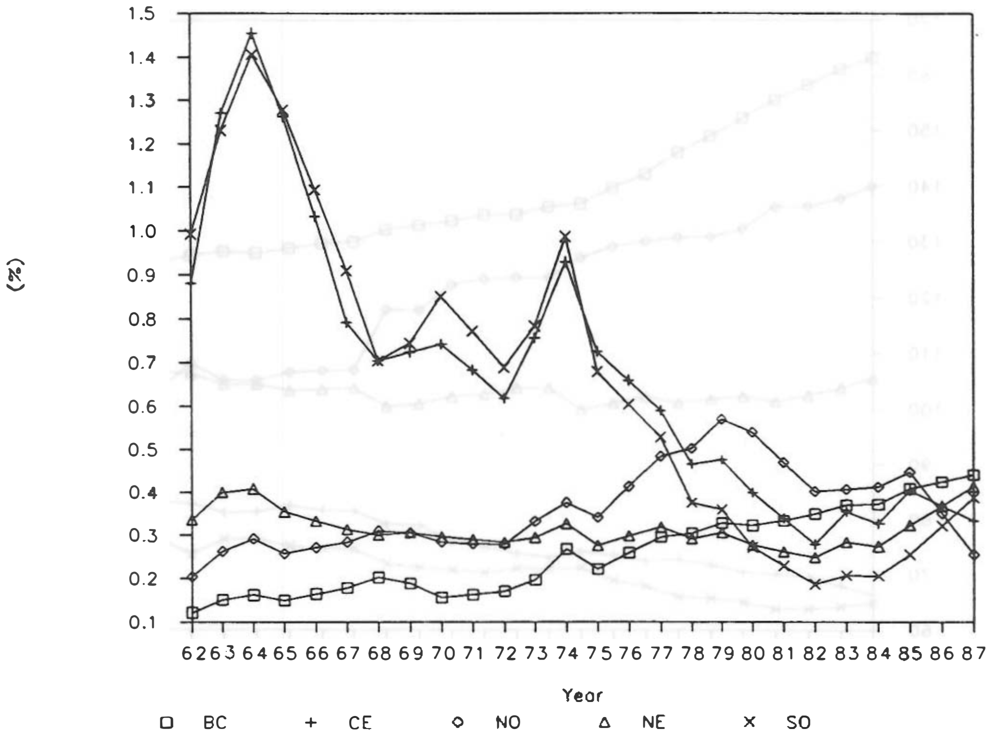


FIGURE 12: Gross Inmigration Rates

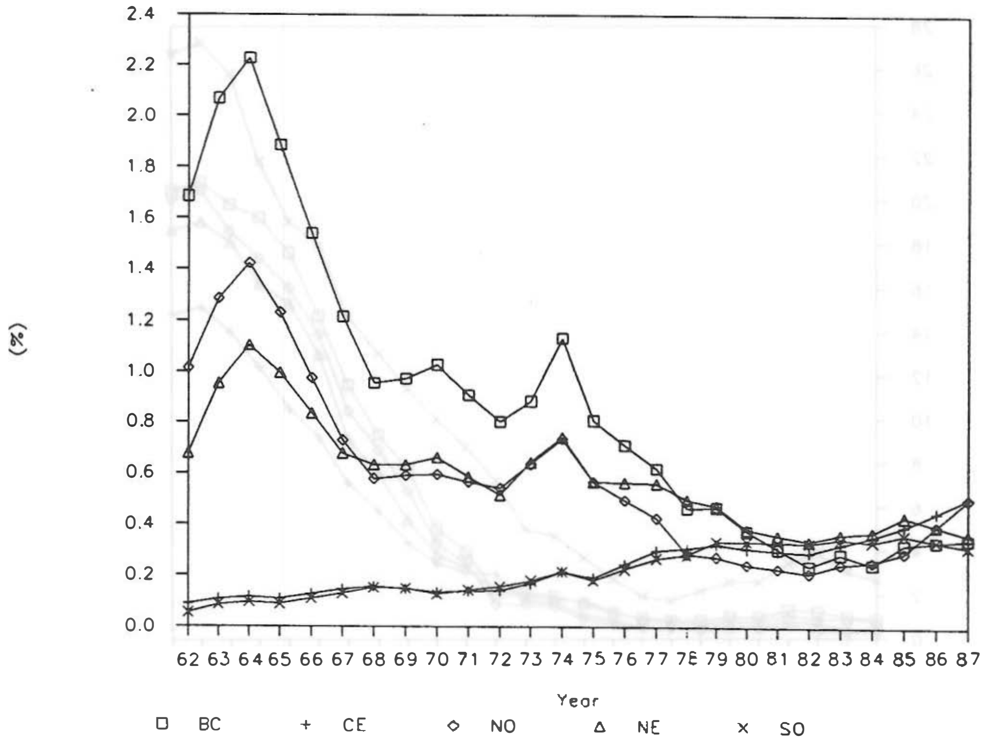


FIGURE 13: Net Inmigration Rates

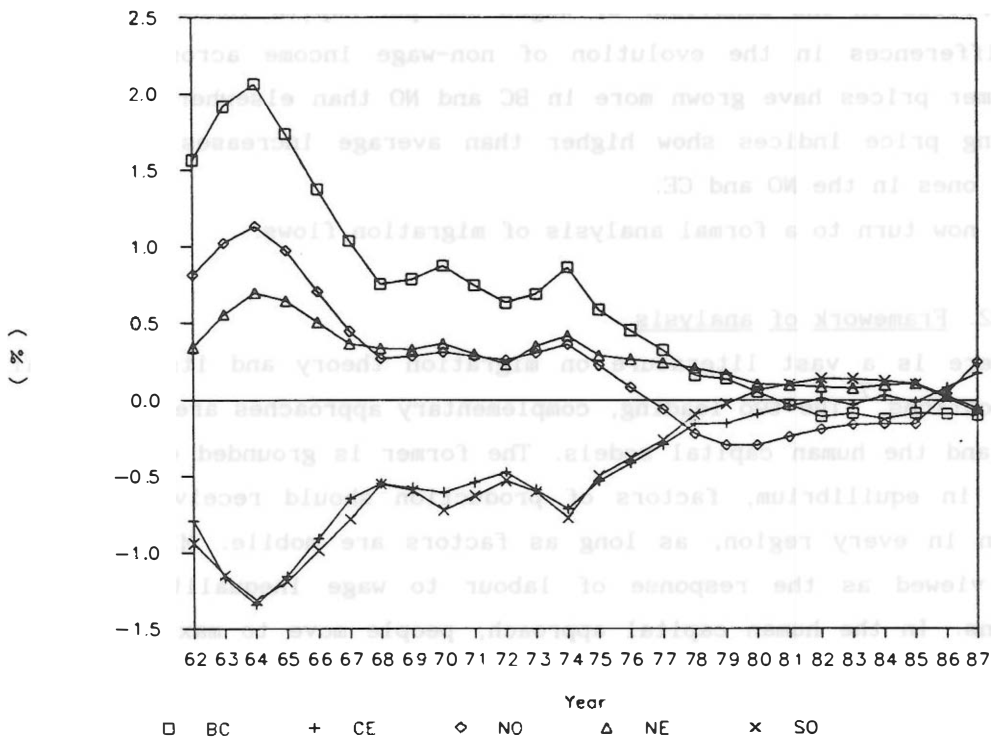
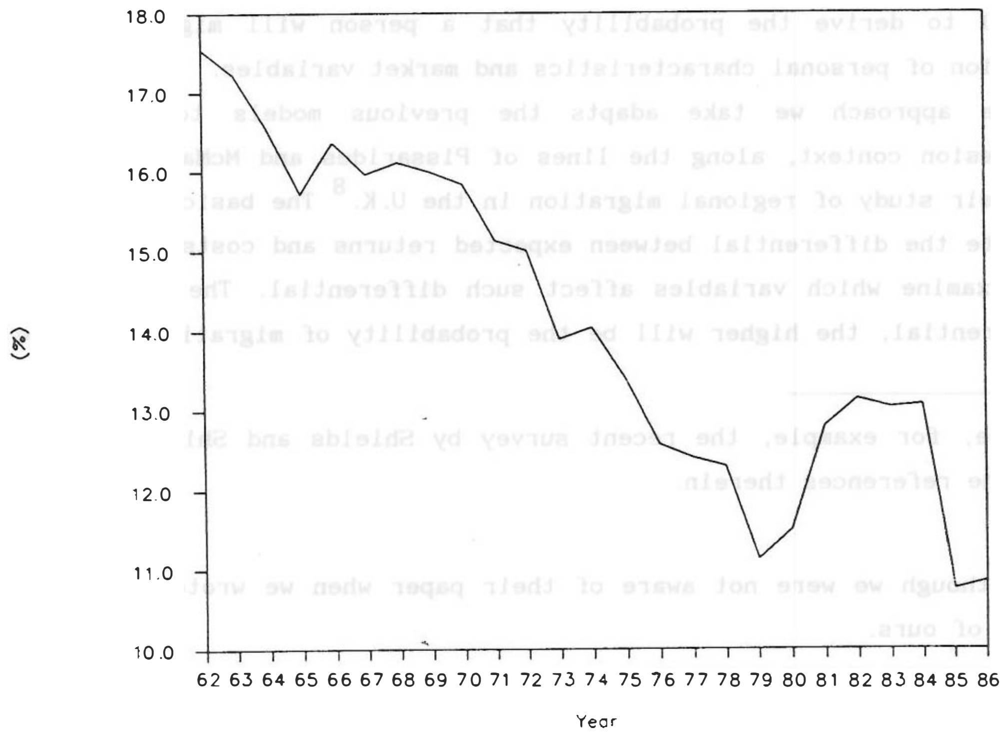


FIGURE 14: Real Wage Inequality Index



wages in CE are more similar to the latter than to those in SO. These disparities in the behaviour of wages and per-capita income point out to differences in the evolution of non-wage income across regions. Consumer prices have grown more in BC and NO than elsewhere. Finally, housing price indices show higher than average increases in SO and lower ones in the NO and CE.

We now turn to a formal analysis of migration flows.

3.2. Framework of analysis

There is a vast literature on migration theory and its empirical applications.[7] The two leading, complementary approaches are the labour flow and the human capital models. The former is grounded on the idea that, in equilibrium, factors of production should receive the same return in every region, as long as factors are mobile. Migration is thus viewed as the response of labour to wage inequalities across regions. In the human capital approach, people move to maximise their expected lifetime utility across regions, which takes into account all features of life in a given place, both economic and non-economic. The model we use is a generalisation of the well-known Harris-Todaro (1970) model, which has been recently adapted by Pissarides and Wadsworth (1987) to derive the probability that a person will migrate, as a function of personal characteristics and market variables.

The approach we take adapts the previous models to a pooled regression context, along the lines of Pissarides and McMaster (1984) in their study of regional migration in the U.K.[8] The basic idea is to compute the differential between expected returns and costs of moving, and examine which variables affect such differential. The higher the differential, the higher will be the probability of migrating. Suppose

---

[7] See, for example, the recent survey by Shields and Shields (1990) and the references therein.

[8] Although we were not aware of their paper when we wrote the first draft of ours.

that there are two regions, 1 and 2, and we want to model the net inflow of migrants to region 1. Each region has an unemployment rate u_i and a wage rate w_i (i=1,2), and there is an unemployment benefit, available to all the unemployed, which is a given proportion b of the wage rate. There are also costs of moving from one region to another, including travel fares, rental or purchase of a house, etc. For analytical convenience we write the costs of living in a given region as a proportion c_i of the respective wage, so that the relevant wage rate is the net wage, $\tilde{w}_i \equiv w_i(1-c_i)$. In this form the cost of moving from region 2 to region 1 can be expressed as c_1-c_2. The expected return for a worker in region i is:

$$R_i = u_i b w_i(1-c_i) + (1-u_i)w_i(1-c_i) = \tilde{w}_i[1-(1-b)u_i] \qquad (i=1,2) \qquad [6]$$

where $0 < b, c_i < 1$.

Denoting by M_1 the net migration into region 1, then the probability of migrating is given by:

$$\text{Prob}\ [M_1 \geq 0] = \text{Prob}\ [R_1-R_2 \geq 0] = \text{Prob}\ [DR_1 \geq 0] \qquad [7]$$

where the return differential is:

$$DR_1 = (\tilde{w}_1 - \tilde{w}_2) - (1-b)\ (u_1\tilde{w}_1 - u_2\tilde{w}_2) \qquad [8]$$

To examine the effects of moving costs and the unemployment and wage rates on the probability of moving, it is useful to write:

$$w_1 = w_2 + dw_1; \quad w_2 = w \qquad [9a]$$
$$u_1 = u_2 + du_1; \quad u_2 = u \qquad [9b]$$
$$c_1 = c_2 + dc_1; \quad c_2 = c \qquad [9c]$$

where dj_1 (j=w,u,c) represents the respective differential, and w, u and c capture the respective overall levels effects. Substituting [9a-c] into [8] we get:

$$DR_1 = dw_1(1-c) - wdc_1 - dw_1dc_1 - (1-b)\ [(u+du_1)(w+dw_1)(1-c-dc_1) - uw(1-c)] \qquad [10]$$

which, by assuming that the cross-differential terms are of smaller order of magnitude, can be approximated by:

$$DR_1 \cong dw_1(1-c) - wdc_1 - (1-b)\,[(udw_1)(1-c) + du_1w(1-c) - uwdc] \quad [11]$$

The effects of the differentials on DR_1 are given by:

$$\partial DR_1/\partial dw_1 = (1-c)\,[1-(1-b)u] > 0 \quad [12a]$$
$$\partial DR_1/\partial du_1 = -\,(1-b)\,w\,(1-c) < 0 \quad [12b]$$
$$\partial DR_1/\partial dc_1 = -\,w\,[1-(1-b)u] < 0 \quad [12c]$$

Therefore net migration into region 1, M_1, depends positively on the wage differential and negatively on the unemployment and moving cost differentials, all with respect to region 2.

It is also easy to derive the effects of the overall levels on DR_1:

$$\partial DR_1/\partial w = -\,dc_1[1-(1-b)u] - (1-b)du_1(1-c) \quad [13a]$$
$$\partial DR_1/\partial u = -\,(1-b)\,[dw_1(1-c) - wdc_1] \quad [13b]$$
$$\partial DR_1/\partial c = (1-b)du_1w - dw_1[1-(1-b)u] \quad [13c]$$
$$\partial DR_1/\partial b = (1-c)\,[udw_1+wdu_1] - uwdc_1 \quad [13d]$$

The sign of the four previous effects are, in general, ambiguous, but can be determined for particular cases. For example, [13a] implies that a higher overall wage reduces migration into regions with higher unemployment and higher cost of moving, and [13b] says that higher overall unemployment discourages inflows into regions with higher wages and lower costs of moving. In any event, the sign of all the level effects have to be determined empirically.

In summary, we can represent the probability of net inflows to region i by:

$$\text{Prob}\,[M_i \geq 0] = f\,[dw_i, du_i, dc_i, w, u, c, b, Z^i] \quad [14]$$

where f is a function satisfying $f_1>0$, $f_2<0$ and $f_3<0$, and the remaining signs are ambiguous. We reserve c_i for the monetary costs of living in region i and Z^i for other relative characteristics, including the weather, public services like education or health care, and amenities.

3.3 Econometric analysis of migration

We now present the results of estimating a formal model of migration. We are concerned with three basic issues. First, we examine how do regional migration rates respond to regional real wage and unemployment levels and differentials, and to moving costs. Second, we analyse the extent to which these differentials persist in the face of migration. Third, we draw some implications from those adjustment processes for regional economic policy. To deal with the first two issues we estimate two pooled cross section-time series regression equations, one for net interregional migration rates and another for regional wage differentials. We close the model by introducing a third equation in an *ad hoc* manner, linking the absolute variation in unemployment with net migration rates, which is derived as a *pseudo-identity* and hence it is not estimated. We take each equation in turn.

Net migration equation

For estimation purposes we use an empirical specification of the probability of migration, similar to the one used by Pissarides and McMaster (1984) for the U.K. and also by the closest piece of work on Spain we know of, Santillana (1978). The latter estimates several cross-section models of bilateral migration flows between Spanish provinces, for selected years between 1960 and 1973. This study finds a significant response of migration to economic variables, mainly to earnings differentials, as well as an important positive effect of the stock of migrants in the province of destination, which the author interprets as proxying for a reduction in informational and settling down costs for new migrants. In contrast, we analyse the more aggregate interregional flows while substantially enlarging the period, and we use a pooling approach to estimation.

The ideal dependent variable would be the movements of workers between regions, but we only have data on population flows. Our dependent variable, m_{it}, is equal to the net flows into a region divided by the region's population in the previous year, which approximates the probability concept.[9] The alternative of estimating separate equations for gross flows was discarded because the unidirectionality of flows makes gross and net migration behave quite alike for most of the period. As shown in Figure 7, the official data on gross migration present extremely low peaks every five years, coinciding with census years. We believe this is an anomaly due to the collection of the new census in those years and have proceeded to smooth out the peaks, by linearly interpolating each region's gross flows, as in the adjusted series in Figure 7.[10]

The regressors, X_i, are wage and unemployment differentials with respect to nationwide levels, i.e. $du_i=u_i-u_N$ and $dw_i=w_i-w_N$, and a set of extra variables, Z^i, proxying for costs of moving, risk aversion, employment structure, influence of external migration, etc., which we discuss below. In order to capture level effects of overall unemployment and wage rates, we allow for regression coefficients to depend inversely on such levels,[11] so that the estimated model is:

$$m_{it} = \alpha_i{}'X_{it} + \varepsilon_{it} \qquad (i=1,\dots,N; \quad t=1,\dots,T) \qquad [15]$$

where $\alpha_i = \alpha_{i0} + \alpha_{i1}\,(1/u_N) + \alpha_{i2}\,(1/w_N)$.

This implies, for example, the presence of absolute differences, du_i and dw_i, and relative differences, defined as $ru_i=u_i/u_N$ and $rw_i=w_i/w_N$, the latter approximated in logarithmic form by $\omega_i=\log(w_i/w_N)$, to be measured in percentage points, as well as cross-product terms like, e.g. ω_i/u_N. Lagged dependent variables are also included, since

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9 See Appendix 2 for sources and definitions of all variables.

10 When we did not make this correction the econometric results -available from us on request- were not very different from those reported below.

11 This procedure was suggested to us by Richard Layard and Stephen Nickell.

migration flows exhibit considerable inertia,[12] and because these variables might proxy for the stock of migrants from other regions.

As a first check of the chance that unemployment and wage differentials have in explaining the downward path of migration, we should recall that relative unemployment rates have in fact converged over time, as revealed by the regional mismatch index computed above. As stated before, the same is not true for absolute differentials of unemployment rates, as is evident from Table 5, which reports du_i for 1962-86 and two subsamples. As to real wage differentials, the table also reports the values of ω_i. A synthetic answer about their evolution is given by the following index of wage inequality:

$$\sum_i (N_i/N) \left[\log (w_i/w_N)\right]^2 \qquad [16]$$

The larger the index, the more different are regional wages, and if all wages are the same the index is equal to zero. Figure 14 plots the index across regions. Inequality was relatively high in the early 1960s, but has been coming down since, except for 1980-84. This is partly explained by a generalised fall in wage dispersion since the early 1970s -as Dolado and Malo de Molina (1985) stress for industry- due, in the second half of the 1970s, to a new system of nationwide wage agreements. García (1988) reports a 27% fall in dispersion across sectors and a 17% fall across occupations.[13]

In order to provide more descriptive measures, Table 6 contains the sample correlation coefficients among the previous variables in the short run (changes) and the long run (levels). We observe small correlations in the short run and larger ones in the long run, but in both cases the signs are the expected ones according to the model.

Regional dummy variables have been included in the set of regressors, to allow for different *fixed effects* on migration rates,

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12 Probably exacerbated by our adjustment of the official data.

13 Fina (1987) remarks that these numbers should be taken with caution, due to a methodological break in the series.

TABLE 5
MIGRATION, WAGES AND UNEMPLOYMENT BY REGION
(Averages in percentages)

Region	Year	m_i	du_i	ru_i	ω_i
AND	1962-86	-0.49	4.11	2.11	-20.62
	1962-75	-0.83	2.47	4.12	-23.30
	1976-86	-0.02	6.18	1.59	-17.19
ARA	1962-86	-0.13	-1.88	0.62	-2.57
	1962-75	-0.26	-0.76	0.89	-1.06
	1976-86	0.03	-3.31	0.73	-4.49
AST	1962-86	-0.07	-1.47	0.69	11.22
	1962-75	-0.07	-0.64	1.01	14.35
	1976-86	-0.07	-2.53	0.81	7.24
BAL	1962-86	0.13	-2.54	0.45	-1.68
	1962-75	0.15	-1.24	0.41	1.43
	1976-86	0.09	-4.19	0.70	-5.64
CAN	1962-86	0.13	1.11	1.03	-11.14
	1962-75	0.12	-0.32	1.33	-9.42
	1976-86	0.15	2.93	1.31	-13.34
CNT	1962-86	-0.06	-2.06	0.59	5.57
	1962-75	0.11	-0.74	0.91	8.24
	1976-86	0.00	-3.73	0.72	2.17
CLM	1962-86	-0.82	-1.33	0.67	-25.55
	1962-75	-1.33	-0.65	1.00	-24.62
	1976-86	-0.16	-2.19	0.88	-26.73
CLE	1962-86	-0.49	-1.29	0.71	-4.88
	1962-75	-0.79	-0.61	1.04	-3.00
	1976-86	-0.11	-2.15	0.85	7.27
CAT	1962-86	0.72	0.00	0.86	6.17
	1962-75	1.32	-0.52	1.13	6.53
	1976-86	-0.04	0.67	1.01	5.73
PVA	1962-86	0.47	-0.25	0.63	10.90
	1962-75	1.03	-1.14	0.51	15.95
	1976-86	-0.25	0.88	1.06	4.48
EXT	1962-86	-0.90	1.88	1.36	-30.32
	1962-75	-1.58	0.57	2.22	-33.07
	1976-86	-0.05	3.56	1.36	-26.83
GAL	1962-86	-0.13	-3.34	0.47	-3.41
	1962-75	-0.20	-0.80	0.85	-5.57
	1976-86	-0.02	-6.58	0.47	-0.66
MAD	1962-86	0.52	0.21	1.05	16.32
	1962-75	0.83	0.05	1.70	17.94
	1976-86	0.13	0.42	1.09	14.26
MUR	1962-86	-0.05	0.10	1.37	-18.44
	1962-75	-0.21	1.17	2.82	-19.83
	1976-86	0.13	-1.25	0.93	-16.67
NAV	1962-86	0.22	-1.26	0.60	-4.55
	1962-75	0.30	-0.92	0.73	-2.44
	1976-86	0.13	-1.69	0.88	-7.25
LRJ	1962-86	0.03	-2.61	0.38	-8.50
	1962-75	-0.14	-1.30	0.35	-10.13
	1976-86	0.24	-4.28	0.60	-6.42
VAL	1962-86	0.45	-0.65	0.90	-8.01
	1962-75	0.69	-0.20	1.46	-9.00
	1976-86	0.14	-1.23	0.89	-6.77

TABLE 6

Sample Correlation Coefficients

(1964-1986)

Short-Run (Changes)

Δm_i	$\Delta\omega_i$	$\Delta\omega_{i-1}$	Δdu_i	Δdu_{i-1}	Δru_i	Δru_{i-1}	Δpa_i
	0.04	−0.13	−0.04	−0.02	−0.09	−0.03	−0.00

$\Delta\omega_i$	$\Delta\omega_{i-1}$	Δdu_i	Δdu_{i-1}	Δpa_i
	0.05	0.01	0.13	−0.01

Long-Run (Levels)

m_i	ω_i	ω_{i-1}	du_i	du_{i-1}	ru_i	ru_{i-1}	pa_i
	0.52	0.51	−0.13	−0.13	−0.26	−0.24	−0.10

ω_i	ω_{i-1}	du_i	du_{i-1}	pa_i
	0.98	0.04	0.10	0.15

introducing a different constant for each region. Through these variables we expect to capture all the subset of variables in Z^i which have not changed systematically during the sample, like the weather or the relative degree of urbanisation.

To take into account the different moving and living costs in different regions, we deflate the nominal wage in each region by the regional retail price index, so that relative wages are real ones. Since finding housing in the region of destination seems of paramount importance in the migration decision, we have also included the relative differential between each region's housing (own imputed and rental) price index (PA_i) and the nationwide index (PA_N), in the logarithmic form $pa_i=\log(PA_i/PA_N)$. A significant coefficient on this variable would signal that migrants give it an importance exceeding its share in the overall retail price index. We expect a negative effect of this variable.

To consider potential effects of the employment composition in the migration decision we included a variable measuring the proportion represented by the building sector in each region's total employment. It has been noted (e.g. Fina (1987)) that construction provided the transition for migrants from agricultural regions when they moved to the cities. A drop in this variable would make migration harder, so we expect a positive sign on this variable. Some researchers suggest that employment growth captures employment opportunities better than the unemployment rate, so we also tried to include in the regression the relative employment growth in a region.

The possibility of having a non-linear utility function in [6], such that $\tilde{w}_i$ and $b(1-c_i)$ are substituted by $V[\tilde{w}_i]$ and $V[b(1-c_i)]$, with $V'>0$ and $V''<0$, has been considered by adding a 3-year moving average of the standard deviations of regional and unemployment rates, which would have a negative effect if risk aversion exists.

Finally, we have used two time dummy drift effects, to capture specific events. The first one, DEUR, taking on a zero value before 1976 and unity afterwards, tries to capture the effects of economic opportunities of migration to European countries. If internal migration was an intermediate step towards external migration, we would expect a negative sign on this variable. The second dummy variable, DPV, is defined as DEUR but just for the Basque Country. It tries to capture political events in that region, which may have induced an outflow of migrants above that explained by purely economic factors.

Since the regressors are potentially endogenous we estimated both by instrumental variables -using two lags of each variable as instruments- and by three-stage least squares. The corresponding Hausman tests for endogeneity of the contemporaneous regressors were never significant and the results were very similar to ordinary least squares (OLS) estimates, so the latter are presented in order to simplify the computation of tests. Due to the use of two lags, the effective estimation period is 1964-1986 and the corresponding number of observations is 391.

The selected specification of the pooled regression, whose encompassing test of the unrestricted equation (i.e. allowing for 3 coefficients per variable, plus cross-terms) is F(30,341)=0.98 (5% critical value=1.46) , appears in the first block of Table 7. Only 10 of the 17 shift dummies were significant at the 10% level, the remaining were excluded from the final specification. The regression fits very well ($\bar{R}^2$=0.98 in the equivalent regression with m_{it} as the dependent variable) and the pooling restrictions (i.e. that all regions have the same coefficients) are accepted at conventional significance levels (FP test in the table).

However, some of the individual equations, reported in Appendix 3, do not fit very well (in particular, AST and CAN) or have wrong signs, revealing some heterogeneity across regions. However, most of the wrong signs are associated to insignificant variables and the low R^2s are associated with regions where little migration flows have occurred. Moreover, of the 51 level coefficients corresponding to relative unemployment and wages, and housing differentials (i.e. β_6, β_7 and β_8), 43 were correctly signed, so we feel reasonably confident about the homogeneity of the chosen specification across regions. There is no indication of first order autocorrelation, and slight signs of heteroskedasticity, as evidenced by the value of White's (1982) test. We report both ordinary t-ratios (in parenthesis) and heteroskedasticity-consistent t-ratios (in brackets); although their levels vary, the implications of both sets of t-ratios are similar. We

TABLE 7

REGRESSIONS

(Pooling)

1. Net Migrations

$$
\begin{array}{lllllll}
\Delta m_{it} = & 0.02 & +0.42\,\Delta m_{it-1} & -0.16\, m_{it-1} & +0.20\,\Delta \omega_{it} & -0.59\,\Delta \omega_{it-1} & -0.06\,\Delta ru_{it} \\
 & (3.07) & (16.31) & (16.30) & (1.61) & (4.25) & (2.73) \\
 & [3.00] & [12.03] & [14.08] & [1.47] & [5.24] & [2.67]
\end{array}
$$

$$
\begin{array}{lllll}
-0.14\, ru_{it-1} & +0.04(\omega_{it}/u_{Nt})_{-1} & -0.02(pa_{it}/u_{Nt}) & -0.077\ DPV & -4.14\ AST \\
(10.04) & (4.46) & (2.33) & (2.90) & (2.05) \\
[15.86] & [4.07] & [2.40] & [3.46] & [3.01]
\end{array}
$$

$$
\begin{array}{llllll}
-4.98\ BAL & +2.34\ CAN & -3.06\ CNT & -5.83\ CLM & -6.61\ CLE & -2.60\ GAL \\
(2.48) & (1.16) & (1.60) & (2.90) & (3.35) & (1.35) \\
[2.61] & [1.43] & [2.00] & [2.40] & [2.81] & [2.57]
\end{array}
$$

$$
\begin{array}{lll}
+3.20\ MUR & +3.50\ NAV & +3.77\ VAL \\
(1.65) & (1.90) & (1.83) \\
[2.25] & [1.68] & [2.87]
\end{array}
$$

$N = 391$, $\bar{R}^2 = 0.72$, $DW = 1.96$, $se = 0.083$, $SS = 2.58$, $SK(2) = 4.28$

$FP\ (134, 237) = 0.96\ [1.20]$, $FHET\ (45, 334) = 1.46\ [1.37]$

2. Wages

$$
\begin{array}{lllllll}
\Delta\omega_{it} = & 0.03 & +0.14\,\Delta \omega_{it-1} & +0.28\,\Delta du_{it-1} & +0.06\, du_{it-1} & -0.23\,\omega_{it-1} & +0.02\, pa_{it-1} \\
 & (6.19) & (4.62) & (2.72) & (12.21) & (10.18) & (2.26) \\
 & [6.80] & [4.40] & [2.68] & [9.44] & [9.43] & [2.36]
\end{array}
$$

$$
\begin{array}{lllllll}
-8.32\ AND & -4.08\ ARA & -1.22\ AST & -3.19\ BAL & -5.85\ CAN & -2.10\ CNT & -9.49\ CLM \\
(8.24) & (5.94) & (2.30) & (4.32) & (6.65) & (3.61) & (8.52) \\
[7.83] & [6.33] & [2.37] & [4.09] & [6.51] & [3.65] & [8.25]
\end{array}
$$

$$
\begin{array}{lllllll}
-4.62\ CLE & -1.78\ CAT & -0.96\ PVA & -11.06\ EXT & -3.55\ GAL & -7.69\ MUR & -4.90\ NAV \\
(6.24) & (3.00) & (1.76) & (9.11) & (4.86) & (7.90) & (6.86) \\
[6.50] & [4.10] & [2.44] & [8.74] & [5.13] & [7.93] & [7.08]
\end{array}
$$

$$
\begin{array}{ll}
-5.05\ LRJ & -4.20\ VAL \\
(6.24) & (5.05) \\
[5.78] & [5.37]
\end{array}
$$

$N = 391$, $\bar{R}^2 = 0.44$, $DW = 2.17$, $se = 0.017$, $SS = 0.114$, $SK(2) = 2.57$
$FP\ (80,289) = 0.91\ [1.27]$, $FHET\ (15,359) = 1.61\ [1.67]$

<u>Note</u>: N : no. of observations

$\bar{R}^2$: coefficient of multiple correlation (corrected by d.f.)

DW: Durbin-Watson statistic

se: Standard deviation of the residuals

SS: Sum of squared residuals

SK: Jarque & Bera's test of normality (distributed as chi-square with 2 d.f.)

FP(.,.): F-Pooling Encompasing test (in brackets 5% c.v.)

FHET(.,.): White's F-Heteroskedasticy test (in brackets 5% c.v.)

The ordinary t-ratios appear in parenthesis whilst White's heteroskedasticity consistent t-ratios appear in brackets.

discuss below the possible origin of a non-constant variance.

In order to avoid perfect collinearity with the constant term, the dummy variable corresponding to Madrid (MAD) has been excluded, so that the fixed effects are interpreted as deviations from the inflow rate in Madrid. There is some association between the dummy coefficients and the weather: regions with warm weather (CAN, MUR, VAL) tend to have significant positive constants above MAD (except BAL). A relationship with the degree of urbanisation is also present, capturing a marginally significant positive constant in the case of NAV and VAL, among the regions with big cities, while PVA and CAT do not differ from MAD.

The response of net migration to changes in wage and unemployment differentials reveal some interesting patterns. Interregional migration rates respond significantly to wage and unemployment differentials in a permanent way, but the response is small, rather slow and affected by the general level of unemployment, the latter confirming the hypothesis advanced by Bentolila and Blanchard (1990). In particular, the (ω_i/u_N) and (pa_i/u_N) terms completely dominated the ω_i and pa_i terms, respectively. The estimated wage effect implies that if in a given region wages rise by 1% above the national average, annual migration into the region rises by 0.002 of 1% of the region's population in the short run,[14] and by $0.0025/u_N$ of 1% in the long run. Thus, the influence of changes in the region's relative wage depends in the long run on the overall level of unemployment, the response being larger in times of low unemployment than in times of high unemployment. So, for example, at 10% national unemployment, the inflow in response to a 1% favourable relative wage is in the long run 0.025 of 1% of the region's population, while at 20% national unemployment the effect is halved. The discouraging effect of the level of national unemployment in the long run but not the short run may be interpreted as follows: when taking a migration decision agents only consider their region's relative wage, while in the long run, once they get information from their predecessors about the difficulties of getting a job in other

---

14 In the estimation, m_{it} is in percentage terms while the unemployment and wage variables are proportions; we do this in order to write coefficients with two digits.

regions, when the nationwide unemployment rate is high, they discount the relative wage rate and their response is consequently lowered.

With respect to unemployment effects, we again find confirmation of the depressing effect of general unemployment, since both in the short and the long run it is relative unemployment what seems to matter. Since relative unemployment rates may be written as:

$$u_i / u_N = 1 + (1/u_N)(u_i - u_N) \qquad [17]$$

having (u_i/u_N) as a regressor is equivalent to assuming that the coefficient on the difference (u_i-u_N) is inversely proportional to u_N: when national unemployment rises, the response of net migration to unemployment differences falls. Numerically, at 10% national unemployment, a 1 percentage point rise in a region's unemployment rate leads to an outflow of 0.006 of 1% of the region's population while in the long run the effect is 0.09 of 1%. As in the effect of relative wages, when the national unemployment rises to 20%, the effect is halved. Comparing the long-run responses of net migration to wages and unemployment differentials, we find evidence that the response is (about four times) larger to unemployment, which has important implications for the design of a sensible regional policy, as discussed below.

As for the remaining explanatory variables, we were only able to find a small effect with respect to relative housing prices, scaled by u_N. A 1% increase in such a variable for a region leads to a net outflow of $0.0002/u_N$ of 1% of the region's population in the short run, and of $0.00125/u_N$ of 1% in the long run. None of the variables proxying for risk aversion, employment structure, employment growth or European opportunities were significant. Contemporaneous and lagged effects were tried, with the corresponding tests being non significant in every case: $F(2,369)=0.65$, 0.46, 0.83 for the variance of wages and employment, and the building sector variable, respectively, and $F(1,370)=1.26$ for DEUR, all well below their respective 5% critical values. Only the DPV dummy was significant, accounting for a relative fall of 7.7% of annual migration in the Basque Country starting from 1976.

Since, as stated before, the behaviour of net migration changes after 1975, it is important to check the stability of the equation. We thus reestimate the equation for two subsamples, 1964-75 and 1976-86, and test for coefficient instability. The two subsample regressions are reported in Table 8. The corresponding Chow test does not reject stability at the 5% level, though it is noticeable that two coefficients suffer shifts. The coefficient on (ω_i/u_N) is not significant in the second subsample, implying the absence of a permanent effect on net migrations after 1975, with unemployment taking the sole role as explanatory variable. This is in agreement with the diminishing role of relative wages as resource allocation mechanisms, as mentioned above. The pa_i coefficient also shifts, reflecting an increasing effect on migration of the widening housing price differentials. The last column in Table 8 presents the estimation of the migration equation without the lagged wage level, with the remaining coefficients hardly changing, which is used for simulation exercises below.

There are also strong downward shifts in the CLM and VAL dummies, and all excluded dummy variables were again not significant in either subsample. Since heteroskedasticity tests do not reveal any problems in either subsample we conclude that the heteroskedasticity present in the full sample is mostly due to the change in the wage and housing price coefficients.

<u>Regional wage equation</u>

The second question that we analyse is the extent to which regional wage and unemployment differentials persist. We want to emphasise from the start that we do not pretend to have a complete model of regional wage determination. The latter issue has been studied in depth by several authors (e.g. Cowling and Metcalf (1967) for the UK and Rodríguez (1988) and references therein for Spain). The common view is that local supply and demand conditions only play a limited role in the determination of regional wages. Our being able to explain just over 40% of the variance of changes in wages -conditioning only on shift dummies and unemployment and housing price differentials- seems to agree with that view. In particular, we do not control for employment

TABLE 8
MIGRATION
Subsamples

Variable	1964 - 75	1976 - 86	1976 - 86
const	0.02 (2.00)	0.03 (2.78)	0.02 (2.72)
$\Delta\ m_{it-1}$	0.43 (11.97)	0.34 (6.47)	0.31 (6.77)
m_{it-1}	-0.17 (12.61)	-0.21 (7.38)	-0.20 (8.26)
$\Delta\ \omega_{it}$	0.11 (0 40)	0.32 (2.12)	0.30 (2.00)
$\Delta\ \omega_{it-1}$	-0.58 (2.15)	-0.58 (3.80)	-0.55 (3.73)
$\Delta\ ru_{it}$	-0.05 (1.80)	-0.07 (1.88)	-0.07 (1.88)
ru_{it-1}	-0.14 (5.78)	-0.13 (7.17)	-0.13 (8.26)
ω_{it-1}/u_{Nt-1}	0.04 (3.50)	0.02 (0.75)	-
pa_{it}/u_{Nt}	-0.02 (1.74)	-0.07 (1.70)	-0.07 (1.84)
AST	-6.62 (1.86)	-3.00 (1.19)	-2.11 (0.94)
BAL	-4.42 (1.36)	-6.45 (2.54)	-6.51 (2.57)
CAN	1.23 (0.38)	4 64 (2.03)	4.34 (1.92)
CNT	-5.70 (1.77)	-2.00 (0.85)	-1.47 (0.66)
CLM	-10.08 (3.05)	-3 00 (1.27)	-3.37 (1.46)
CLE	-12.18 (3.68)	-2.66 (1.15)	-2.34 (1 03)
PVA	-0.97 (0.30)	-9.67 (3.73)	-8.99 (3.75)
GAL	-3.50 (1.13)	-2.58 (1.13)	-2.26 (1.01)
MUR	3.14 (1.09)	4.11 (1.70)	3.51 (1.55)
NAV	4.70 (1.60)	2.74 (1.16)	2.50 (1.11)
VAL	7.57 (2.10)	1.02 (0.44)	0.67 (0.30)
N	204	187	187
se	0.094	0.068	0.068
R^2	0.74	0.68	0.68
DW	0.85	2.07	2.05
SK(2)	3.86	3.02	2.99
FHET	1.36 [1.35]	1.16 [1.35]	

FCH(22,351) = 1.24 [1.57]
Note: As in Table 7. FCH: Chow's F-Stability test.

composition across regions, while in his study Rodríguez (1988) concludes that interprovincial nominal wage differentials in Spain are to a large extent due to the specialisation of southern regions in low-wage sectors.

The OLS regression model results for the change in relative wages is reported in the second block of Table 7. Contrary to the case of migration, most shift dummies are very significant, with AND, CLM, EXT and MUR well below the relative growth in MAD and AST, CNT and PVA around the same growth as in MAD, which is the highest. This is possibly related to the composition of employment across regions: the regions corresponding to the first four dummies are all in the South of Spain, those associated to the latter three dummies are in the North. The pooling and the remaining parameter restrictions pass at the 5% level, and there are no signs of either first-order autocorrelation or heteroskedasticity in the residuals.

There are two non-competing hypothesis to test. The first is a short-run one, the *Phillips curve* view by which as desired employment increases employers raise wages to attract more labour, i.e. a negative relationship between unemployment and changes in wages, in our case between variables which are relative to the national mean. We were not able to capture this effect in neither the full sample nor the second subsample, with only small favorable evidence in the first subsample (see below). The positive coefficients on the level and the differences of relative unemployment seemingly imply that the mere rise in actual employment, in spite of the fall in wages, is sufficient for employers in order to attract labour. Though the simple correlations in Table 6 point in the same direction, we believe this not to be a satisfactory feature but, since data on the many other variables which may influence regional wages were not available to us -and the statistical performance of the equation is fairly correct- we keep it as is.

The other hypothesis concerns expected wage equalisation. In the long run, with perfect labour mobility and in the absence of risk aversion or other variables, expected returns would equalise across regions, i.e. $R_i = R_N$, with R_i as in [6]. Since $R_i/R_N=1$, $\ln(R_i/R_N)=0$, and the long-run condition reads:

$$\ln(w_i/w_N) + \ln\{ [1-(1-b)u_i]/[1-(1-b)u_N] \} + \ln[(1-c_i)/(1-c_N)] = 0 \quad [18]$$

Making use of the approximation $\ln [1-(1-b)u_i] \cong -(1-b)u_i$ and the functional assumption $(1-c_i)/(1-c_N) = (PA_i/PA_N)^{-\vartheta}$, we get:

$$\ln (w_i/w_N) = (1-b) (u_i - u_N) + \vartheta\, pa_i \quad [19]$$

If there were no unemployment benefits, expected wages would equalise across regions, whereas if the unemployment compensation were 100% of the wage rate, then the expected return, apart from moving costs, would be equal to the wage rate, hence it would be wage rates that tend to be equal: $w_i=w_N$. In between there is a continuum of cases indexed by b, giving rise to a positive relation between the region's relative wage (w_i/w_N) and the relative unemployment level, u_i-u_N, with a coefficient possibly different from unity.

Another possibility of getting a coefficient different from unity arises if factors such as risk aversion play a role, causing some non-linearity in the expected returns from moving or staying. It is, however, possible to show under simple assumptions, like having a concave constant relative risk aversion (CRRA) utility function, that the tradeoff between the unemployment differential u_i-u_N on the relative wage $\ln(w_i/w_N)$ would be larger than unity. In this sense, we expect to be able to discriminate between the linear utility *cum* replacement ratio hypothesis and the concave CRRA utility hypothesis, by examining the size of the estimated slope coefficient.

In our estimates there is a well defined, albeit small, positive relationship between a region's relative wage and its unemployment differential. Regions with above-average unemployment will tend to have, in the long run, above-average wages. So wages eventually compensate for differences in unemployment, but at the rate 0.3:1. The restriction that the coefficients of the two level terms are equal in absolute value is clearly rejected (t-ratio=8.03). To be precise, we find that in a long-run steady state relative wages, in deviation of their drifts and their relative moving costs, are a fixed proportion of relative unemployment rates. Making all first difference terms equal to zero, the long-run solution yields:

$$\ln (w_i/w_N) = D_i + 0.09\, pa_i + 0.3\, (u_i - u_N) \qquad (i= 1, \ldots, 17) \qquad [20]$$

where D_i is the long-run value of the i-th dummy and we have used the same approximation as in [19].

Hence our estimates imply that in the short run wages do not respond to above-average unemployment (recall that the contemporaneous value of u_i does not appear in the regression) while in the long run, if a region has, on average, 1 percentage point of unemployment above the national level, its wage rate would be 0.3% above the average rate for the country as a whole. According to our discussion of [19] this implies an estimated replacement ratio of 0.7, which fits reasonably well with the scanty evidence we have about the sample average of this variable. For example, Dolado *et al.* (1986) construct its time series for industrial workers, obtaining a sample mean of 0.67 for 1966-86, with a range going from 0.61 at the beginning of the sample to 0.75 at its end. To examine the hypothesis of risk aversion we included as extra regressors the standard deviations of regional wage and employment rates, but both were insignificant. This gives us some confidence that the replacement ratio hypothesis provides a better interpretation of the evidence.[15]

It should be noted that, as in the migration equation, convergence to the long-run equilibrium is slow. For example, if a given region's relative wage is 1% above the national average, *ceteris paribus*, this will lead, in the following year to a 0.28 of 1% decline in the region's relative wage rate, 0.20 of 1% in the second year, 0.17 of 1% in the third year and so on, until the original deviations are eliminated. Finally, we find a positive effect of relative rental housing prices on relative wages, such that in the short run, if a region's housing price index is 1% above the national average, its real wage rate would tend to be 0.02 of 1% above the nationwide average, whilst such effect is 0.1 of 1% in the long run.

---

[15] We decided against including the replacement ratio in the wage or the migration equation as a regressor, because we feel the quality of the constructed measure is too low.

As with the migration equation, we estimate the wage equation in the two subsamples 1964-75 and 1976-86. The results are presented in Table 9. Apart from the coefficients on the lagged changes of wages and of unemployment differentials the remaining ones do not change much, and the Chow test of parameter stability is easily passed at the 5% level. The pooling restrictions are again not rejected, but we found that in the first subsample the contemporaneous value of the unemployment differential was negative with a t-ratio of 1.62 (not reported). This may point to a short-run Phillips curve effect being present in that subsample, an effect not found at all in the second subsample. The individual equations (see Appendix 3) show the same picture, with only AND and NAV showing some signs of a Phillips effect.

Regional unemployment equation

To close the model we need an equation for regional unemployment rates. We made several attempts using migration rates and wages, but all of them failed. At the end, in order to be operative at the simulation stage, we reached a compromise solution by which the equation was not estimated but derived as a *pseudo-identity*, following the same approach as in Pissarides and McMaster (1984).

Just to see what is involved assume that there are no outflows and all new immigrants, M_{it}, are unemployed during the first year of arrival. Then the absolute number of unemployed in region i would be $U_{it} = U_{it-1} + M_{it}$. Dividing by the population (POP) at t-1 and after some simple algebra we can express the unemployment rate as:

$$u_{it}=\frac{U_{it}}{L_{it}}=\frac{L_{it-1}}{L_{it}}\frac{U_{it-1}}{L_{it-1}}+\frac{POP_{it-1}}{L_{it}}\frac{M_{it}}{POP_{it-1}}=\frac{L_{it-1}}{L_{it}}u_{it-1}+\frac{POP_{it-1}}{L_{it}}m_{it} \qquad [21]$$

With a basically constant labour force -which one gets for Spain comparing 1986 with 1962- and assuming a constant population-labour force ratio this could approximately be written as:

$$\Delta u_{it} \cong \alpha\, m_{it} \qquad [22]$$

where α^{-1} is the labour force-population ratio. Guided by the sample

TABLE 9
WAGES
Subsamples

Variable	1964 – 75	1976 – 86
const	0.04 (4.68)	0.05 (5.16)
$\Delta\ w_{it-1}$	0.25 (4.90)	0.14 (3.06)
$\Delta\ du_{it-1}$	0.51 (1.52)	0.28 (2.49)
ω_{it-1}	-0.27 (6.51)	-0.35 (6.91)
du_{it-1}	0.05 (6.75)	0.07 (8.16)
pa_{it-1}	0.03 (1.60)	0 01 (1.15)
AND	-10.94 (5.84)	-10.74 (5.92)
ARA	-4 98 (4.66)	-6.15 (4.95)
AST	-1.08 (1.50)	-2.26 (2.72)
BAL	-4.33 (3.95)	-5.43 (4.00)
CAN	-7.11 (4.80)	-9.42 (6.02)
CNT	-2.78 (3.38)	-3.10 (2.97)
CLM	-10.88 (5.47)	-14.37 (6.49)
CLE	-5 58 (4.77)	-7.16 (5.23)
CAT	-2.70 (3.11)	-2.65 (2.84)
PVA	-0.83 (1.17)	-2.40 (2.58)
EXT	-13.85 (6.05)	-14.73 (6 49)
GAL	-5.43 (4.20)	-4.61 (3.98)
MUR	-9.14 (4.91)	-11.00 (6.22)
NAV	-5.68 (5.15)	-7.21 (5.39)
LRJ	-6.74 (4.52)	-6.90 (5.03)
VAL	-5.77 (3.88)	-6.73 (4.93)
N	204	187
SS	0.053	0.052
se	0.017	0.018
R^2	0.42	0.48
DW	2.16	2.26
FHET	0.83 [1.83]	1.02 [1.83]

FCH(22,347) = 1.35 [1.55]

Note: As in Table 7 and Table 8.

value of this ratio, around 0.4, and realizing that assuming full unemployment for the migrants is an extreme, we chose for α the value of 1.5 (in fact 0.015, recall that we use m_i in percentage terms but not u_i), which would correspond to 60% of migrants being unemployed during the first year in the simplified case above. Given that the wage changes in the simulations below are small, we feel some confidence about the order of magnitude chosen. Experimentation with values around 0.01 in the simulations below only changed the results slightly, giving in any case a stable path of the system towards the long-run compensating equilibrium between wages and unemployment rates. For example, taking the long-run coefficients derived from the estimated migration and wage equations for the total sample with u_N=15% we would have, abstracting from constants and housing prices:

$$m_i = 1.67\, w_i - 5.83\, u_i \quad (+ \text{ terms in } \Delta w_i, \Delta u_i, \Delta m_i) \qquad [23a]$$
$$w_i = 0.21\, u_i \quad (+ \text{ terms in } \Delta w_i, \Delta u_i) \qquad [23b]$$
$$\Delta u_i = 0.015\, m_i \qquad [23c]$$

Ignoring short-term changes and substituting the first two equations into the third the evolution of u_i is be guided by an AR(1) process with a root of 0.92, whereas if the parameter in the unemployment equation was 0.01 the coefficient would be 0.95, also stable.[16] Notice that the process by which the regional unemployment rates tend to be equalised (in deviations from drifts and housing prices) is quicker the lower the level of overall unemployment. For instance, if u_N was 7.5%, then the root of the AR(1) process would be 0.89, greatly reducing the length of the adjustment. In order to analyse more precisely the pattern of adjustment, and the concomitant role of regional policy, we devote the next section to a properly dynamic simulation of the model.

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[16] If short-term differences were included we would get a more general AR(p) process, whose gain at L=1 would be identical to the previous values, giving a similar picture of the adjustment process; see Wickens and Breusch (1988).

3.4 Simulations and regional employment policy implications

Our previous empirical findings support the view that:

(i) net regional migration rates respond nonlinearly to wage and unemployment differentials, and

(ii) there is a tendency for regional unemployment and wage differentials to approach a long-run compensating equilibrium in terms of deviations of their respective drifts and housing prices.

If the process by which wage and unemployment differentials disappear was fast, and there were no differences in drifts and moving costs, these findings would cast doubt on the effectiveness of a regional employment policy. The latter could then be defined only in a neutral sense: moving jobs to people or the opposite.

This conclusion is certainly true in the long run. But according to our empirical findings there are significant drifts in wage differentials and also the adjustment of migration and wages is slow. Regional employment policy might then be able to speed up the process of equalisation. We would advocate several types of policies.

First, measures could be taken to encourage jobs to move before people do. Some of these policies should give firms more incentives to hire more labour in depressed regions, like marginal employment subsidies or the lowering of payroll taxes.[17] This would apply to regions like the Northeast, with relatively -in national terms- high unemployment rates and real wages. But labour is already cheapest in the South, which reveals the lack of other conditions. In particular, to be effective, these policies should be complemented by higher investment in infrastructure in general, and communications in particular, which has been very low in Spain in the recent past (see Viñals *et al.* (1990)). In this way, less favoured areas could become an attractive alternative for the location of firms, which is not the case now.

A second set of policies would make it economically more profitable for people to move, for instance by stepping up the program of subsidies for migration established in 1986 and by temporarily subsidising housing in target areas, so that housing price

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[17] Similar measures have been proposed for the UK by C. Huhne (1990).

differentials would have less of a deterring effect on migration. Finally, a third type of policy should aim at variables which in our econometric model are the regional dummy variables, i.e. public goods and amenities which make it more pleasant to live in a given area: education, health care, cultural events, etc. The distribution of these public goods is also quite uneven in Spain, and its improvement could provide an important motivation for migration.

Because geographical disparities have always been sizable, regional development policy is an old-age subject in Spain. The classic study is Richardson (1975), and a more recent survey is provided by Martín (1988). Regional policies have been implemented in the past, which helps account for the convergence of regional GDP per capita over time (as shown in Figure 9). However, the long period of recession, 1975-85, brought a large part of the official development policy to a halt. What is needed is to take advantage of the economic growth taking place today to divert a larger part of those resources into regional development policies. In fact, some of the policies we advocate are considered in the medium-term regional development plan of the Spanish Government for 1989-1993 (Ministerio de Economía y Hacienda (1989a,b)), but it remains to be seen whether they will be implemented or not. One should be aware that on these issues there is a nontrivial coordination problem between the central government and the 17 regional governments, which depends on a stable agreement on the financing of the latter, which has still not been reached.

Can the efficiency of these policy proposals be defended analytically? JLS show that a rigorous case for expected tax differentials in favour of high unemployment groups can be made and that such differentials should be higher the less responsive is the labour force to wage inequalities. Since we find low responsiveness and slow adjustment to economic incentives, our proposals would meet the conditions required by the theory.

In order to illustrate the role of these policies, we have used the estimates of the regional system formed by the migration and wage equations to compute the following simulation. We take as the initial state one in which a given region's unemployment rate is 1 percentage point above the national rate -which is kept unchanged at 15%, not an unlikely forecast for 1990- assuming zero drifts and housing price differentials. This will encourage net outmigration until the differential is eliminated. The model is closed by using the proposed relation between changes in regional unemployment rates and net migration rates. Since the migration equation shows some signs of instability, we have performed the calculation for the total sample and the two subsamples. The results are in Table 10 and in Figures 15,16 and 17.

It is clear that unemployment differentials only vanish very slowly, as roughly illustrated above by the high root that its autoregresive representation would have. In the first subsample, it takes 9 years after the initial shock for half of the initial shock to disappear, while the corresponding length for the second subsample and the total sample is between 11 and 9 years, respectively. During the adjustment there is continuous outmigration, which reaches a peak 5 or 6 years after the initial shock and then declines slowly. Wages also adjust, initially increasing due to the effect of lagged wages and larger unemployment, but falling afterwards, as unemployment falls. They reach a peak in the second year (first year for the second subsample) and then fall monotonically.

The role of regional employment policy in this example can be evaluated by the *saving* in unemployment points that would happen if the initial regional unemployment rate is increased by 1 percentage point. This is measured by the sum of the elements in each of the columns labelled *du* in Table 10. The estimated *person-years* of unemployment would be 10 points for the first subsample, where the adjustment is quicker, and between 13 and 16 points for the second subsample and the total sample. Given the homogeneity of the system in terms of the initial level of unemployment, an initial shock of x percentage points to the region's unemployment differentials will take between 10x and 16x percentage points of unemployment, depending on the case, before the differential is eliminated. Calculated gains depend negatively on the national unemployment rate. For example, if instead of 15% the national rate was assumed to be 10%, the preceding gains would be divided by a factor around 1.5. In other words, at times of high unemployment, like the present situation, people are less likely to

FIGURE 15: Regional System

Adjustment Path (Total Subsample)

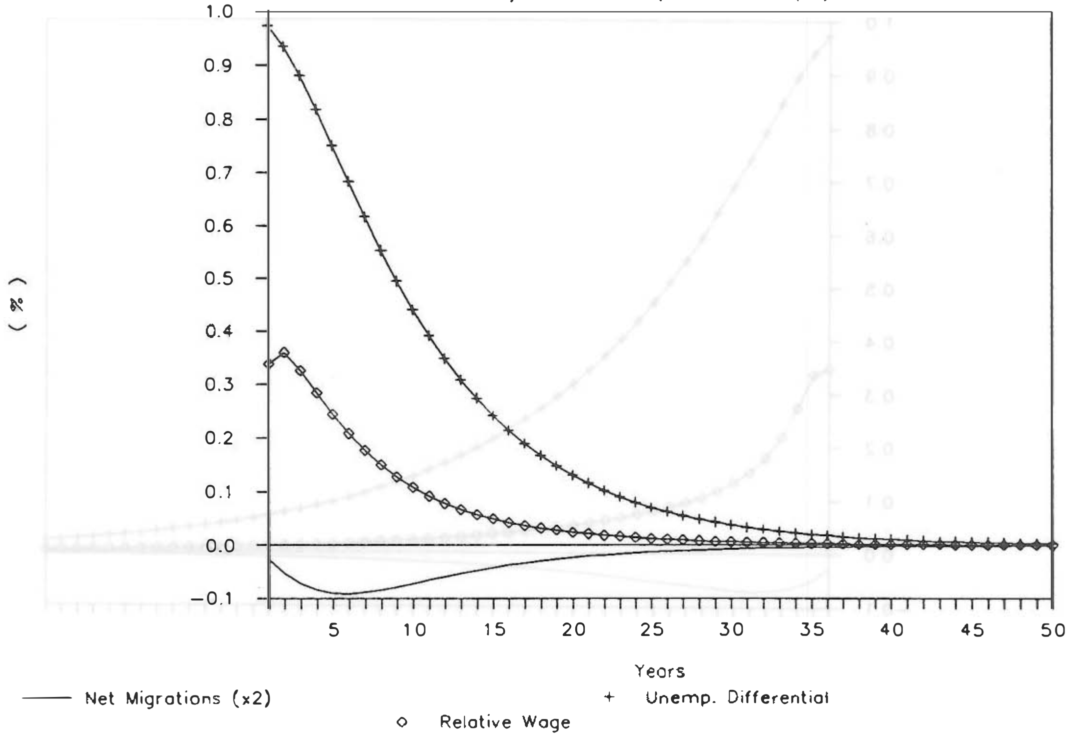
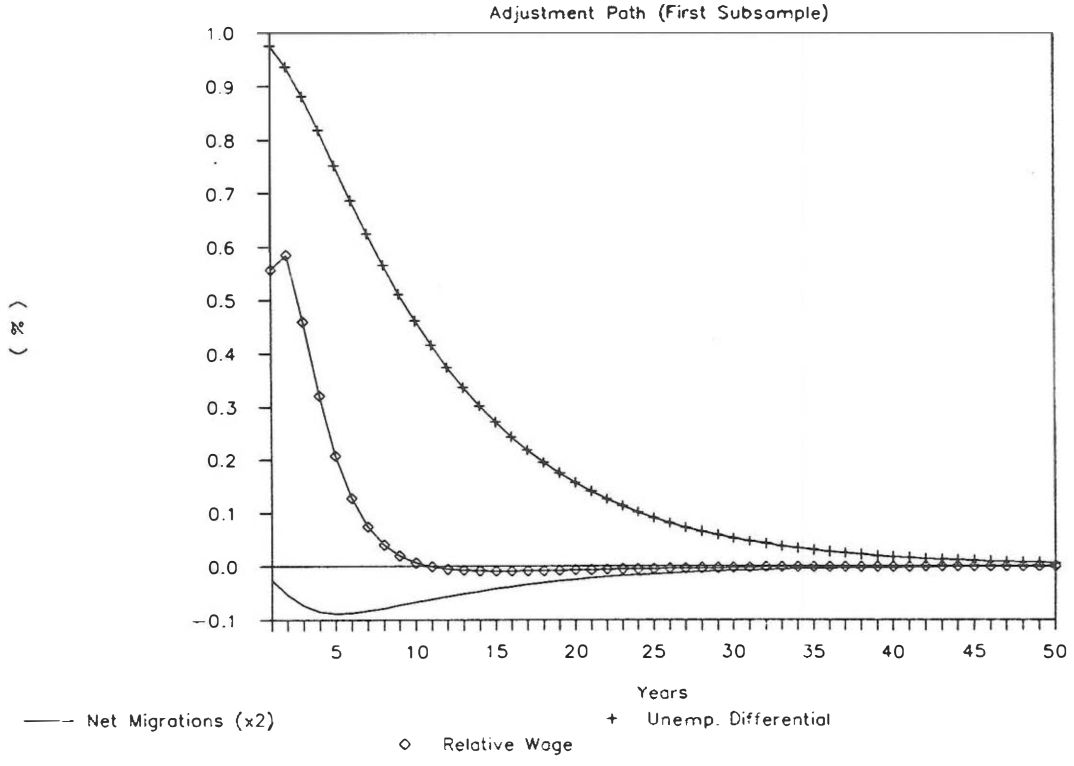


FIGURE 16: Regional System

Adjustment Path (First Subsample)



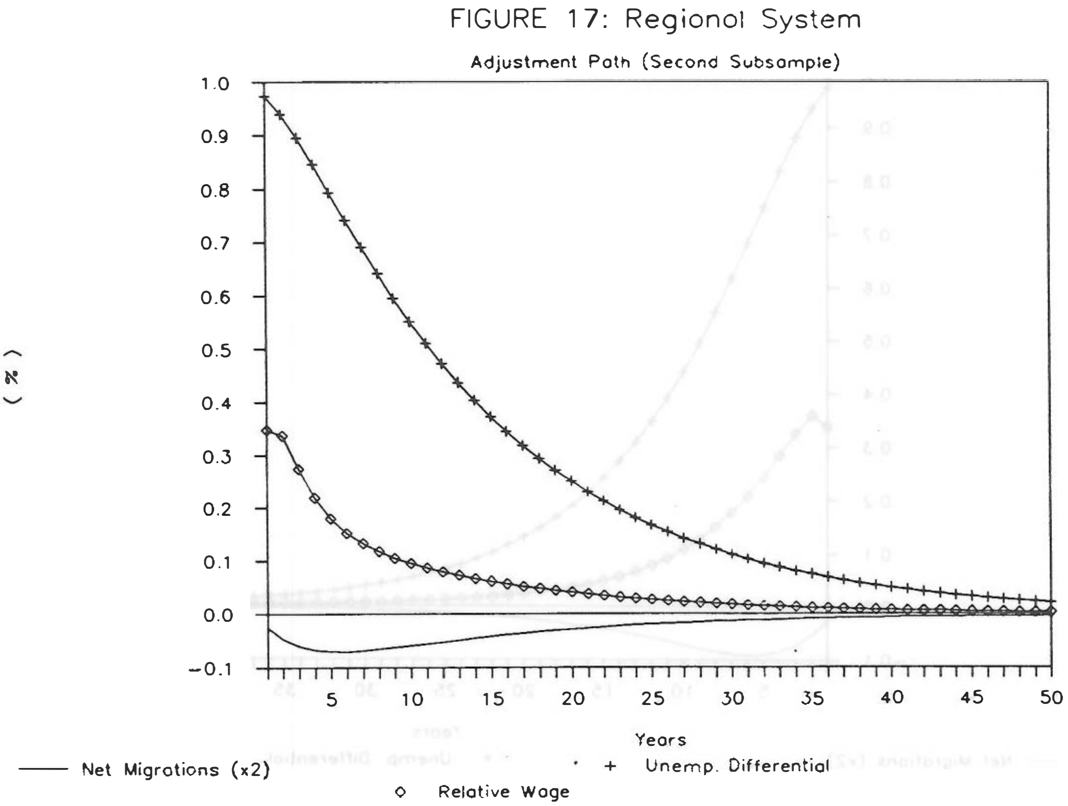


FIGURE 17: Regional System
Adjustment Path (Second Subsample)
1.0
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0.0
−0.1
(%)
5
10
15
20
25
30
35
40
45
50
Years
Net Migrations (x2)
+ Unemp. Differential
◇ Relative Wage

TABLE 10: REGIONAL SYSTEM DYNAMIC ADJUSTMENT PATH

	Whole sample			First subsample			Second subsample		
Year	m	du	w	m	du	w	m	du	w
1	-0.031	0.933	0.333	-0.039	0.918	0.556	-0.034	0.926	0.342
2	-0.031	0.886	0.336	-0.041	0.857	0.561	-0.036	0.872	0.321
3	-0.030	0.841	0.287	-0.040	0.797	0.411	-0.036	0.818	0.248
4	-0.029	0.797	0.239	-0.038	0.740	0.260	-0.034	0.767	0.190
5	-0.028	0.756	0.202	-0.037	0.684	0.151	-0.033	0.717	0.152
6	-0.027	0.716	0.174	-0.036	0.630	0.083	-0.031	0.671	0.128
7	-0.026	0.677	0.152	-0.035	0.578	0.044	-0.030	0.626	0.112
8	-0.024	0.641	0.135	-0.034	0.527	0.022	-0.028	0.584	0.101
9	-0.023	0.606	0.121	-0.032	0.479	0.010	-0.026	0.545	0.092
10	-0.022	0.572	0.110	-0.030	0.434	0.003	-0.025	0.508	0.085
11	-0.021	0.540	0.101	-0.029	0.391	-0.002	-0.023	0.473	0.078
12	-0.020	0.510	0.093	-0.027	0.351	-0.005	-0.022	0.440	0.072
13	-0.019	0.481	0.086	-0.025	0.314	-0.008	-0.020	0.410	0.067
14	-0.018	0.453	0.080	-0.023	0.279	-0.010	-0.019	0.381	0.062
15	-0.017	0.427	0.075	-0.021	0.248	-0.012	-0.018	0.355	0.058
16	-0.016	0.402	0.070	-0.019	0.219	-0.013	-0.017	0.330	0.054
17	-0.016	0.379	0.065	-0.017	0.193	-0.013	-0.015	0.307	0.050
18	-0.015	0.357	0.061	-0.016	0.170	-0.013	-0.014	0.285	0.046
19	-0.014	0.336	0.057	-0.014	0.148	-0.013	-0.013	0.265	0.043
20	-0.013	0.316	0.054	-0.013	0.129	-0.013	-0.012	0.246	0.040
21	-0.012	0.297	0.050	-0.011	0.112	-0.013	-0.012	0.229	0.037
22	-0.012	0.280	0.047	-0.010	0.097	-0.012	-0.011	0.213	0.035
23	-0.011	0.263	0.044	-0.009	0.084	-0.011	-0.010	0.198	0.032
24	-0.010	0.247	0.041	-0.008	0.072	-0.011	-0.009	0.184	0.030

25	-0.010	0.232	0.039	-0.007	0.061	-0.010	-0.009	0.171	0.028
26	-0.009	0.219	0.037	-0.006	0.052	-0.009	-0.008	0.159	0.026
27	-0.009	0.205	0.034	-0.005	0.044	-0.009	-0.007	0.148	0.024
28	-0.008	0.193	0.032	-0.005	0.037	-0.008	-0.007	0.137	0.022
29	-0.008	0.182	0.030	-0.004	0.031	-0.007	-0.006	0.128	0.021
30	-0.007	0.171	0.028	-0.004	0.026	-0.007	-0.006	0.119	0.019
31	-0.007	0.160	0.027	-0.003	0.021	-0.006	-0.006	0.110	0.018
32	-0.006	0.151	0.025	-0.003	0.017	-0.005	-0.005	0.102	0.017
33	-0.006	0.142	0.024	-0.002	0.014	-0.005	-0.005	0.095	0.015
34	-0.006	0.133	0.022	-0.002	0.011	-0.004	-0.004	0.088	0.014
35	-0.005	0.125	0.021	-0.002	0.009	-0.004	-0.004	0.082	0.013
36	-0.005	0.117	0.019	-0.001	0.007	-0.003	-0.004	0.076	0.012
37	-0.005	0.110	0.018	-0.001	0.005	-0.003	-0.004	0.071	0.011
38	-0.004	0.104	0.017	-0.001	0.004	-0.003	-0.003	0.066	0.011
39	-0.004	0.097	0.016	-0.001	0.003	-0.002	-0.003	0.061	0.010
40	-0.004	0.092	0.015	-0.001	0.002	-0.002	-0.003	0.057	0.009
41	-0.004	0.086	0.014	-0.001	0.001	-0.002	-0.003	0.053	0.009
42	-0.003	0.081	0.013	0.000	0.000	-0.002	-0.002	0.049	0.008
43	-0.003	0.076	0.013	0.000	0.000	-0.001	-0.002	0.046	0.007
44	-0.003	0.071	0.012	0.000	0.000	-0.001	-0.002	0.042	0.007
45	-0.003	0.067	0.011	0.000	-0.001	-0.001	-0.002	0.039	0.006
46	-0.003	0.063	0.010	0.000	-0.001	-0.001	-0.002	0.037	0.006
47	-0.003	0.059	0.010	0.000	-0.001	-0.001	-0.002	0.034	0.006
48	-0.002	0.055	0.009	0.000	-0.001	-0.001	-0.002	0.032	0.005
49	-0.002	0.052	0.009	0.000	-0.001	-0.001	-0.001	0.029	0.005
50	-0.002	0.049	0.008	0.000	-0.001	0.000	-0.001	0.027	0.004

move, and, thus, the role of regional employment policy can be substantial.

4. Conclusions

In this paper we continue the search for the causes of the rise in unemployment in Spain. We start by analysing the behaviour of mismatch, from two different angles. First we follow the approach that considers relative unemployment rate dispersion as an indication of mismatch. We document that absolute differentials in unemployment rates across categories such as sex, age or region have increased as the overall rate has risen. Relative unemployment rate dispersion indices have, however, fallen over time in most of these categories; i.e., mismatch measured in this way seems to have been reduced.

In contrast, the estimation for Spain of a disequilibrium model provides a measure of mismatch, understood as the heterogeneity of constraints on firms in different markets, which steadily increases over time. We find we can statistically explain the path followed by this index by regressing it on a set of variables related to mismatch. On account of this and the casual evidence provided by the current overheating situation in Spain we reluctantly conclude that the disequilibrium measure seems to provide a more accurate picture of the behaviour of mismatch in Spain.

Part of the rise in mismatch, as measured by the latter index, comes from the evolution of the regional distribution of economic variables. To gain a better understanding on this aspect of mismatch, and given that internal migration has steadily decreased in Spain since the early 1960s, we set up and estimate an econometric system modelling internal migration flows and regional wage differentials. We find, on the one hand, that interregional migration responds significantly to economic variables such as real wage and unemployment differentials, but with a relative small value and also with long lags. On the other hand, the overall unemployment rate and housing price differentials are also found to deter migration. We then simulate the dynamic response of the system to an exogenous increase in the unemployment rate of a region, finding that the convergence of the process to a long-run equilibrium with compensating wage differentials is very slow. We infer that a regional policy targeted at moving jobs to people -in contrast to relying on the movement of people to jobs- could save a sizable amount of unemployment during the short- and medium run, specially starting from a high national unemployment rate. Other policy recommendations, related to tax schemes and regional housing price measures are also drawn.

A lot remains to be done. Our econometric model can still be improved by including omitted variables, unavailable at this time, but which are probably important, like those approximating the availability of public goods and amenities, or the demographic characteristics of the population by region. Moreover, cross-section variation could be gained by repeating the exercise for provinces instead of the more aggregate regions.

Similarly, this kind of dynamic system could be estimated for the case of mobility across economic sectors. Also, the issue of skill scarcity, the only interesting dimension where the unemployment dispersion mismatch index is increasing in the recent past, should be explored. The construction sector, where shortages of manpower in general -and skills in particular- have been reported in the last few years, would be an extremely interesting sector to analyse. We have to leave these issues for further research.

APPENDIX 1: GROUPING OF REGIONS INTO 5 AGGREGATE REGIONS

1. BIG CITIES (BC):	4. Baleares (BAL)
	9. Cataluña (CAT)
	13. Madrid (MAD)
2. NORTH (NO):	3. Asturias (AST)
	6. Cantabria (CNT)
	10. País Vasco (PVA)
3. NORTHEAST (NE):	2. Aragón (ARA)
	15. Navarra (NAV)
	16. La Rioja (LRJ)
	17. Valencia (VAL)
4. CENTRE (CE):	7. Castilla La Mancha (CLM)
	8. Castilla León (CLE)
	12. Galicia (GAL)
5. SOUTH (SO):	1. Andalucía (AND)
	5. Canarias (CAN)
	11. Extremadura (EXT)
	14. Murcia (MUR)

APPENDIX 2: SOURCES AND DEFINITIONS

A) Migration and Population:

The source is: Instituto Nacional de Estadística (INE), *Anuario Estadístico de España*, years 1960 to 1987.

B) Labor force, employment and unemployment data:

For national data, the source is: INE, *Encuesta de Población Activa*, years 1964 to 1989. In the available disaggregate classifications, we have used the homogenisation of these data in Ministerio de Economía y Hacienda (MEH) (1987). For regional employment and unemployment the source is: Banco de Bilbao, *Renta Nacional de España y su Distribución Provincial*, issues of 1962 to 1985 (available every other year except for 1966). An interpolation to get annual data was done using the profile of the aggregate unemployment series from the first source in the previous paragraph. The levels were rescaled to that of the homogenised data in MEH (1987).

C) Wages:

The source is as for regional unemployment data. The series is nominal compensation of employees in the province divided by the number of wage earners. The interpolation to get annual data was done using the profile of the wage series in: INE, *Encuesta de Salarios*, 1962 to 1986. Nominal wages were deflated by the consumer price index for each province from: INE, *Indice de Precios al Consumo*, years 1962 to 1986. All provincial CPI series were set to 100 in 1962. Housing and rental prices are taken from the appropriate item in the retail price index.

Regional data were calculated by weighting each province's wages by its share of dependent employment in the region.

The aggregation of provinces into regions is the official one, established in 1978, extended backwards to 1962.

APPENDIX 3
MIGRATION EQUATIONS

$$\Delta m_{it} = \beta_0 + \beta_1 \Delta m_{it-1} + \beta_2 m_{it-1} + \beta_3 \Delta\omega_{it} + \beta_4 \Delta\omega_{it-1} + \beta_5 \Delta ru_{it} + \beta_6 ru_{it-1} + \beta_7 (\omega_{it}/u_{Nt})_{-1} + \beta_8 (pa_{it}/u_{Nt})$$

Region	β_0	β_1	β_2	β_3	β_4	β_5	β_6	β_7	β_8	R^2	DW
AND	0.03	0.57	-0.11	0.10	0.42	-0.11	-0.20	-0.21	-1.30	0.36	2.06
	(0.77)	(2.69)	(1.23)	(0.06)	(0.25)	(0.57)	(3.04)	(0.25)	(1.14)		
ARA	0.09	-0.13	-0.02	-0.40	-1.12	-0.12	-0.67	0.97	0.12	0.16	2.28
	(0.76)	(0.63)	(0.13)	(0.80)	(2.28)	(0.95)	(0.62)	(0.56)	(0.20)		
AST	-0.13	0.00	-0.36	0.25	-0.47	0.08	1.24	0.24	-0.18	0.05	2.01
	(1.36)	(0.02)	(1.29)	(0.52)	(1.23)	(1.01)	(1.04)	(0.78)	(0.80)		
BAL	-0.06	0.76	-0.14	0.55	-0.56	-0.07	-0.16	0.24	-0.61	0.87	1.98
	(1.29)	(6.10)	(1.13)	(1.36)	(1.26)	(0.47)	(6.68)	(1.24)	(1.53)		
CAN	0.20	0.17	-0.51	0.57	-0.58	-0.08	-0.67	0.61	0.33	0.05	1.38
	(1.27)	(0.40)	(1.16)	(0.99)	(1.10)	(1.04)	(0.91)	(1.05)	(0.70)		
CNT	0.00	0.26	-0.55	0.18	-0.41	0.01	-0.15	0.63	-0.33	0.20	1.93
	(0.01)	(0.90)	(2.37)	(0.91)	(0.47)	(0.13)	(0.13)	(1.57)	(0.69)		
CLM	0.01	0.50	-0.18	0.11	-1.85	0.16	-0.09	0.98	-0.35	0.77	1.95
	(0.15)	(4.64)	(2.00)	(1.01)	(1.82)	(1.10)	(1.80)	(1.86)	(0.25)		
CLE	-0.29	0.50	-0.24	0.50	-1.23	-0.10	-0.27	0.54	-0.27	0.25	2 21
	(0.73)	(2.16)	(2.40)	(0.72)	(0.72)	(0.26)	(0.67)	(0.58)	(0.09)		
CAT	0.01	0.26	-0.10	0.37	-1.25	-0.27	-0.13	-0.20	-0.16	0.84	2.40
	(0.36)	(2.64)	(1.51)	(2.17)	(1.83)	(1.09)	(6.01)	(0.96)	(1.46)		
PVA	0.19	0.53	-0.34	0.38	-0.77	0.16	-0.22	0.47	-0.12	0 84	2.32
	(1.42)	(4.81)	(6.35)	(1.05)	(0.53)	(0.83)	(2.92)	(0.49)	(0.80)		
EXT	0.03	0.40	-0.15	0.18	-0.70	0.00	-0.14	-0.20	-0.15	0.89	2.37
	(0.70)	(5.72)	(4.30)	(0.92)	(1.07)	(0.06)	(10.8)	(0.44)	(1.00)		
GAL	0.05	0.16	-0.19	-0.02	-0.71	-0.20	-0.13	1.41	-0.95	0.46	2.49
	(1.34)	(0.72)	(2.17)	(0.04)	(1.42)	(2.49)	(1.69)	(2.47)	(2.03)		
MAD	0.03	0.40	-0.20	1.25	1.82	0.10	-0.15	0.56	0.10	0.74	2.43
	(0.92)	(3.32)	(2.66)	(2.67)	(2.17)	(1.05)	(3.52)	(1.06)	(0.82)		
MUR	0.05	0.11	-0.19	0.19	-0.73	-0.01	0.28	0.99	-0.65	0.27	1.84
	(1.04)	(0.51)	(2.10)	(0.53)	(2.06)	(0.22)	(0.72)	(2.20)	(1.40)		
NAV	0.40	-0.21	-0.35	0.45	0.10	-0.25	-2.16	0.63	-1.87	0.76	1.79
	(4.57)	(1.74)	(4.24)	(1.00)	(0.16)	(5.46)	(5.17)	(1.75)	(1.56)		
LRJ	0.03	-0.17	-0.17	0.42	-0.49	-0.29	-0.15	0.54	-0.16	0.55	1.88
	(0.83)	(0.92)	(1.33)	(1.44)	(1.62)	(1.51)	(3.77)	(1.13)	(0 38)		
VAL	0.02	0.71	-0.23	0.87	-0.72	-0.01	-0.12	0.19	-0.13	0.52	2 30
	(0.07)	(3.81)	(1.63)	(2.42)	(0.85)	(0.06)	(0.04)	(1.14)	(2.26)		

Note: The coefficients β_7 and β_8 are multiplied by 10.

APPENDIX 3

WAGE EQUATIONS

$$\Delta\omega_{it}=\beta_0+\beta_1+\beta_2\Delta\ du_{it-1}+\beta_3\ du_{it-1}+\beta_4\ \omega_{it-1}+\beta_5\ pa_{it-1}$$

Region	β_0	β_1	β_2	β_3	β_4	β_5	R^2	DW
AND	-0.39 (4.24)	0.36 (1.51)	-0.64 (1.79)	0.20 (3.88)	-0.43 (4.27)	0.11 (1.90)	0.44	2.15
ARA	-0.01 (1.56)	-0.18 (0.71)	0.52 (0.93)	-0.42 (0.68)	-0.07 (0.27)	0.10 (1.20)	0.04	2.00
AST	0.02 (1.03)	0.27 (1.08)	0.07 (0.10)	0.05 (0.96)	-0.23 (1.64)	0.00 (0.08)	0.08	1.85
BAL	0.01 (0.69)	0.22 (1.56)	0.34 (0.71)	0.04 (3.04)	-0.39 (2.30)	0.05 (0.99)	0.58	2.00
CAN	-0.02 (1.13)	0.15 (0.79)	0.38 (0.83)	0.04 (1.97)	-0.27 (2.21)	0.04 (0.60)	0.24	2.22
CNT	0.02 (2.52)	0.33 (2.45)	-0.34 (0.37)	0.06 (2.38)	-0.38 (3.84)	0.03 (0.06)	0.48	2.33
CLM	-0.05 (0.86)	0.12 (0.64)	-0.18 (0.49)	0.08 (2.48)	-0.22 (0.90)	-0.11 (1.68)	0.31	2.34
CLE	-0.02 (0.22)	0.17 (0.57)	0.81 (0.97)	0.21 (0.02)	-0.31 (1.30)	0.00 (0.00)	0.11	2.10
CAT	0.05 (3.03)	0.10 (0.76)	-0.09 (1.13)	0.56 (3.07)	-0.46 (3.21)	0.06 (1.80)	0.38	2.12
PVA	0.03 (4.24)	0.28 (2.86)	-0.18 (0.25)	0.05 (3.55)	-0.27 (5.13)	0.19 (2.93)	0.65	2.15
EXT	-0.08 (2.70)	0.52 (0.46)	0.24 (1.03)	0.05 (2.62)	-0.20 (2.40)	0.12 (2.14)	0.60	2.30
GAL	-0.03 (1.50)	0.10 (0.28)	0.64 (0.87)	-0.82 (0.41)	-0.53 (1.83)	-0.04 (1.04)	0.03	1.80
MAD	0.07 (2.16)	0.31 (1.72)	0.08 (0.16)	0.68 (0.65)	-0.40 (2.42)	0.10 (1.28)	0.26	1.87
MUR	-0.08 (2.73)	0.24 (1.59)	0.40 (1.03)	0.04 (2.73)	-0.45 (2.90)	0.02 (1.00)	0.53	2.26
NAV	-0.01 (1.76)	0.37 (3.00)	-0.48 (1.33)	0.06 (3.20)	-0.22 (1.84)	-0.02 (0.33)	0.64	1.70
LRJ	0.00 (0.10)	0.05 (0.58)	0.85 (2.00)	0.06 (5.47)	-0.16 (1.70)	0.11 (1.90)	0.70	1.78
VAL	0.00 (0.00)	0.09 (0.44)	0.26 (0.45)	0.07 (2.15)	-0.12 (1.63)	0.02 (0.18)	0.17	1.52

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