Labour market dynamics in Spanish regions: evaluating asymmetries in troublesome times*

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31 January 2014

Abstract

The Spanish labour market disproportionately booms in expansions and bursts in recessions; meanwhile, its regions' relative position persists: those with the highest unemployment rates in 1996 were also in the worse position in 2012. To examine this twofold feature, we apply Blanchard and Katz's (1992) methodology and evaluate how the Spanish labour market reacts to regional employment shocks in a variety of cases. Shock responses are channelled via changes in unemployment, labour market participation, and spatial mobility. Our results provide evidence of asymmetric responses across business cycle phases (1996-2007 and 2008-2012). While changes in participation rates are the main adjustment mechanism in expansion, unemployment and spatial mobility become the central ones in recession. We also provide evidence of real wage rigidities in both periods, but strong asymmetries in house prices, which are sticky in recession but notably reactive in expansion. We conclude with a cluster analysis showing that high and low unemployment regions have similar responses in the short-run while, in the long-run, the former are more reactive in terms of spatial mobility. Overall, we provide evidence that people are more willing to migrate when a regional shock occurs in relatively worse economic contexts.

JEL Codes: J20, E24, J61, R11.

Keywords: Employment, Unemployment, Regional labour markets, Spain.

* **Acknowledgments:** We are grateful to the Spanish Ministry of Economy and Competitiveness for financial support through grant ECO2012-13081.

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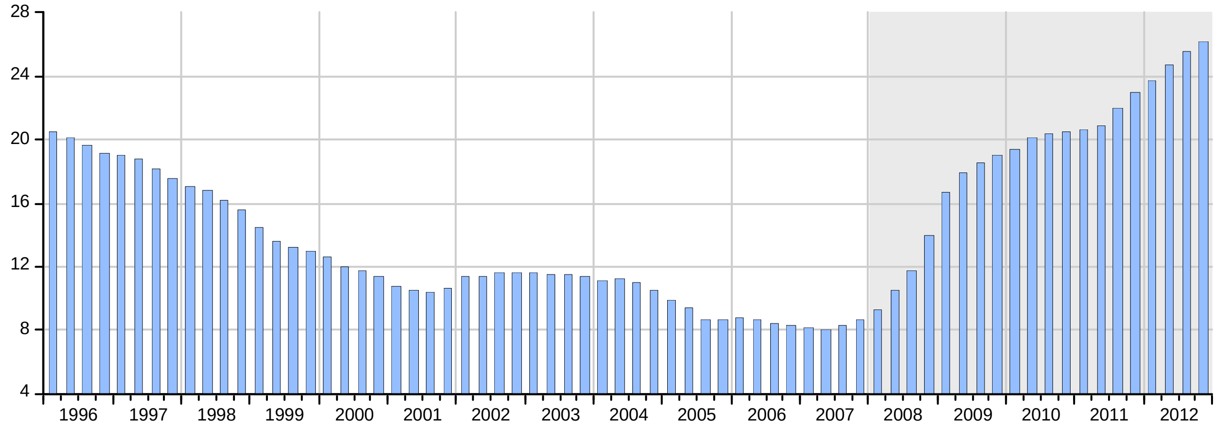
1 Introduction

The Great Recession has severely hit Spain in many dimensions. No more than other economies in most of them (economic (de)growth, sovereign debt crisis, banking system collapse), but disproportionately hard on unemployment. After more than a decade trending downwards and converging to the European average, the rate of unemployment reached 8.0% in 2007 –falling from a peak of 24.5% in 1994 and values above 20% still in 1996. In 2012, however, after five years of steep rise, the historical maximum was surpassed reaching a massive 26.0%.

Figure 1. Quarterly unemployment rate in Spain. 1996-2012.



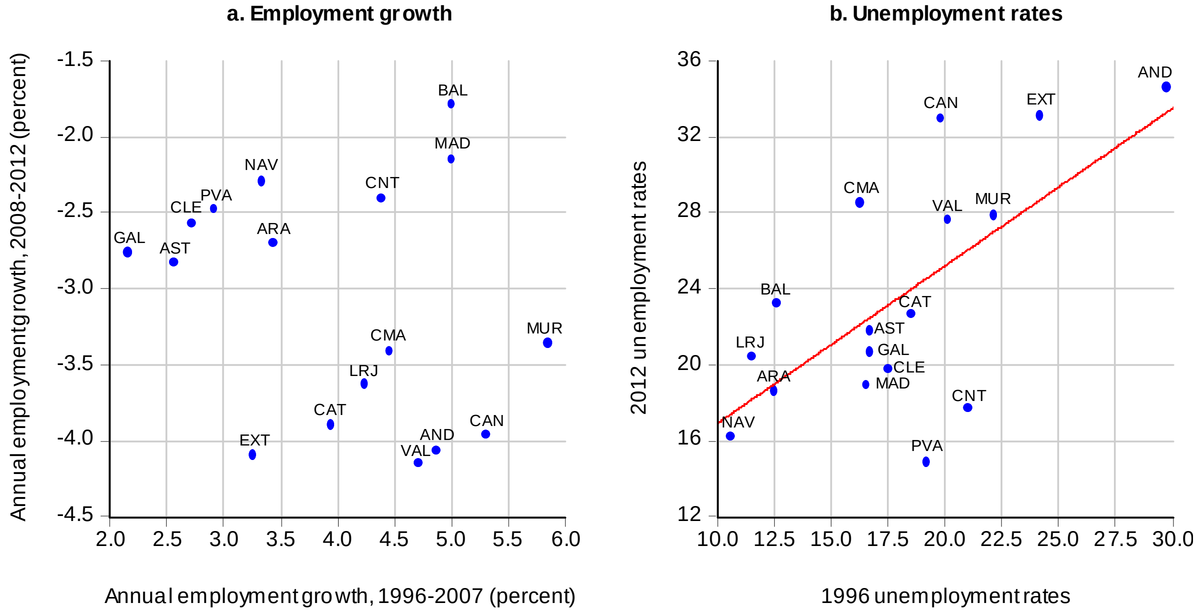
Source: Spanish Labour Force Survey (EPA).

The intense progress first, and deterioration afterwards, of the Spanish labour market goes in parallel with an extreme degree of regional persistence in labour outcomes. This is illustrated in Figure 2. Figure 2a shows two groups of regions. The first one with employment growth rates around -2.5% in 2008-2012 (with the Balearic Islands and Madrid close to -2.0%), and a second one between -3.3% and -4.1%.[1] The difference between the two groups points to the existence of less responsive regions in the North and North-West of Spain (Galicia, Asturias, Castile and Leon, Basque Country, Navarre, Aragon), and more volatile ones in the South and East part of Spain. Madrid (also the Balearic Islands and to some extent Cantabria) would be a salient exception with top employment performance simultaneously in good and bad times. Figure 2b, in contrast, gives a much homogeneous picture in terms of unemployment rates, with a regression slope of 0.83 and adjusted R^2 of 0.39. When combined, the information supplied by Figures 2a and 2b discloses two

[1] Given that a simple regression line takes a misleading downward slope, it is not drawn. We should rather think on two upward sloping lines, one per group, indicating that well-performing regions coincide in booms and busts.

main stylised facts: (i) changes in employment provide just a partial explanation of the evolution of unemployment, and (ii) there is a great persistence in regional unemployment over the years.

Figure 2. Labour market performance of Spanish regions. 1996-2012.



Source: Spanish Labour Force Survey (EPA).

AND=Andalusia; ARA=Aragon; AST=Asturias; BAL=Balearic Islands; CAN=Canary Islands; CNT=Cantabria; CLE=Castile and Leon; CMA=Castile-La Mancha; CAT=Catalonia; VAL=Valencian Community; EXT=Extremadura; GAL=Galicia; MAD=Community of Madrid; MUR=Region of Murcia; NAV=Navarre; PVA=Basque Country; LRJ=La Rioja.

These facts and the regional specificity of the Spanish labour market may be studied from a variety of perspectives, taking into account, along the lines of Marston (1985), that changes in regional (un)employment may be the outcome of both national and regional driving forces. Elhorst (2003) distinguishes four types of approaches including single-equation models, implicit models (where he places the Blanchard and Katz model), accounting identity models, and simultaneous-equation models dealing with interactions. The strength of the implicit models are their solid theoretical basis, while simultaneous equation models should be chosen from an empirical viewpoint (Elhorst, 2003, p. 741).

Multi-equation structural models have been used in Bande and Karanassou (2009, 2013a and 2013b) to assess to what extent the evolution of differences in Spanish regional unemployment can be attributed to disparities in the respective regional equilibrium unemployment rates or to the evolution of other key variables such as, for example, capital accumulation.[2] Our aim, however, is to analyse the regional labour market from a regional specific point of view. It would be too demanding, in our context, to conduct a detailed analysis using their Chain Reaction Theory methodology. The reason is that we consider small sample periods of study, as deserved by the unprecedented specificities of the recent economic developments, at the same time that we need information highly disaggregated by regions. On one side, this causes severe restrictions in terms of degrees of freedom. On the other side, it constrains the analysis to a relatively small number of variables quarterly available for all Spanish regions, and with up-to-date coverage.[3]

We are interested in answering questions related to the most recent evolution of the Spanish labour market. What has happened regarding the specific regional responses to labour market shocks? Have they changed relative to previous responses, studied for the period up to the mid 1990s? Are these responses similar in good and bad times? What role do prices play? The framework of analysis developed in Blanchard and Katz (1992) allows us to provide answers to these questions. It yields the possibility of evaluating the impact of employment shocks through the responses they cause, not only in terms of the unemployment rate, but also through changes in participation rates and regional mobility.[4] Such analysis will enhance our understanding, from a regional perspective, of the labour market adjustment mechanisms in the different scenarios studied.

The model of Blanchard and Katz has been used to investigate the dynamics of regional labour markets in the US (Blanchard and Katz, 1992), Europe (Decressin and Fatás, 1995), Sweden (Fredriksson, 1999), the Netherlands (Broersma and van Dijk, 2002), Finland (Mäki-Arvela, 2003) and, more recently, for the German East-West disparities (Alecke *et al.*, 2010). It has also been used to analyse the Spanish labour market by uncovering its regional persistence in 1976-1994 (Jimeno and Bentolila, 1998), and to provide specific analyses on the Southern regions (Murillo *et al.*, 2006) and by level of education (Mauro and Spilimbergo, 1999).

Notwithstanding its wide use, it is important to discuss two of its prominent features since it is a model that relies upon (1) the assumption of perfect mobility across regions; and (2) the measurement of regional variables as deviations from the national average, which implies that shocks are regionally idiosyncratic.

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[2] Other significant articles concerned with Spanish labour market regional disparities are López-Bazo and Motellón (2012 and 2013) and, with a specific focus on wage setting, Bande *et al.* (2008 and 2010).

[3] Note that National Accounts data at the regional level are issued with severe delays, and other regional databases, such as the BD-Mores, only cover up to 2007. Moreover, none of them provide data at quarterly frequencies.

[4] "The *feedback* effects of the regional unemployment rate on labour supply, labour demand and regional wage-setting in simultaneous equations models dealing with interactions are comparable to those in Blanchard and Katz" (Elhorst, 2003, p. 723).

Regarding the first assumption, we acknowledge that Spanish regions are not fully characterised by perfect mobility of workers and firms. But, as the literature in this field does, we argue that this methodology can still deliver useful insights. A clear example is the analysis of the European regions conducted by Decressin and Fatás (1995) despite the well known fact that mobility across countries is lower than within countries. Beyond that, it is important to note that a critical value added of our analysis is the comparison across periods. On this account, although the Spanish recent experience has been characterised by an important inflow of workers, such increase has been shared by most regions.

Table 1 shows the average relative internal residential migration rate in each Spanish region (i.e., the internal residential migration within a region as a ratio over the total internal residential migration in Spain). It can be seen that this ratio has remained roughly constant between 1998-2007 and 2008-2012 which are, basically, the two periods of analysis in this work. Our analysis, therefore, is not subject to biases stemming from variations in the regional migration behaviour.[5]

Table 1. Relative internal residential migration (%).[6] 1998-2012.

AND		ARA		AST		BAL		CAN		CNT	
98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12
14.4	14.2	2.2	2.3	1.7	1.7	3.4	3.3	5.5	5.3	1.5	1.4
CLE		**CMA**		**CAT**		**VAL**		**EXT**		**GAL**	
98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12
5.3	5.0	4.6	5.1	20.0	19.2	11.3	11.1	1.6	1.6	5.2	5.2
MAD		**MUR**		**NAV**		**PVA**		**LRJ**			
98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12	98-07	08-12		
14.2	15.2	2.4	2.5	1.6	1.6	4.0	4.0	0.7	0.7		

Note: the complete name of each region is provided in the note below Figure 2.

Source: National Statistics Institute (Variaciones Residenciales Interiores).

Regarding the fact that we are only evaluating region-specific shocks, we acknowledge that nation-wide shocks may also be relevant, as argued by Bande and Karanassou (2009, 2013b) for other periods. As shown below, however, the fact that the estimated

[5] The period averages shown in Table 1 are not hiding relevant information. Yearly examination of this ratio yields the same conclusion.

[6] The sample period is 1998-2012 due to data availability. Data used is internal residential migrations by region of destiny. We have also checked the internal residential migration by region of origin, which yields very similar results.

employment growth equation has an explanatory power below 0.50 in 11 out of 17 regions is an indication of the relevance of our analysis; even if it also leaves space for nation-wide shocks whose effects cannot be examined within our analytical framework. In a sense, therefore, this work should be interpretted as complementary to the existing ones conducted through the estimation of multi-equation models.

In any case, the novelty of our analysis neither lies in the use of Blanchard and Katz's methodology nor is a mere time extension of the work by Jimeno and Bentolila (1998). The paper contributes to the literature in three main dimensions.

One contribution is the specific evaluation of the effects of average regional employment shocks when hitting in expansion and when hitting in recession. For this, we use quarterly data (as Jimeno and Bentolila, 1998) and consider two subsample periods: 1996-2007, covering the expansion; and 2008-2012, covering the crisis. This disaggregation allows an evaluation of the asymmetries in shock responses across business cycle phases (upward and downward).

Another key contribution consists in extending the labour market model to include prices. This extension was already present in Blanchard and Katz (1992), but it has generally been disregarded in subsequent literature. Consideration of prices in Spain is a relevant issue both in 1996-2007 and in 2008-2012. In expansion, it allows us to assess the response of wages and house prices to the improved economic conditions of the workers. In recession, it allows us to examine to what extent price adjustments have followed the intense quantity adjustments characterising the Spanish economy in recent years. Summing up, we offer new information on how prices respond regionally to labour demand shocks, and the potential asymmetries of these responses in good and bad times.

A third key contribution, finally, is the additional disaggregation by groups of regions based on a cluster analysis. The two resulting groups (one including Catalonia, Madrid, Navarre, and the Basque Country, and the other one grouping the rest of the regions) are used to re-estimate the models and conduct the analysis for the two groups.

Our findings are diverse. First, we identify asymmetric labour market responses across business cycle phases. We find that changes in participation rates are the main adjustment mechanism in expansion, while unemployment becomes the central one in recession. Moreover, the long-run employment impact is larger when the shock hits in a recessive period than when it hits in expansion. This result is an indication that net migration –spatial mobility– is more relevant in troublesome than in good times.

We also provide evidence of real wage rigidities in both periods along the lines of Jimeno and Bentolila (1998).[7] In contrast, when house prices are considered, we find

[7] Jimeno and Bentolila (1998) uncovered a high degree of persistence in the Spanish regions, as compared to the US and the EU. This was mainly due to real wage rigidities and low interregional

an asymmetric response across the different phases of the business cycle. During the expansionary years, there is a significant reaction in house prices, which does not take place in recession. During the crisis, house prices are not sensitive and the differences with respect to the analysis with product and labour market prices are not substantial. We interpret these results as evidence of spillover effects taking place in good times going from the real state to the labour market, and vice-versa. This mutually feeding spiral led to a housing bubble and booming rates of employment growth in all Spanish regions.

There is, finally, evidence of similar labour market dynamics across high and low unemployment regions generated by the one-off employment shocks. This is consistent with the large degree of (un)employment persistence affecting all Spanish regions, and is to some extent reassuring in the sense that consideration of an average Spanish region is not flawing the results. Nevertheless, we still find differences in the relative long-run regional employment impact, resulting on larger spatial adjustments in high than in low unemployment regions, which appear as more resilient to the shock. On this account, it seems safe to conclude that people are more willing to migrate not just when regional shocks take place in a recessive period, but also when they impact in places with larger relative unemployment rates.

The remaining of the paper is structured as follows. In Section 2, we outline the analytical framework and its empirical implementation. In Section 3 we present the data used and the econometric methodology. In the following four sections we show our findings related, respectively, to the aggregate analysis, the disaggregation by business cycle phases, the inclusion of price responses, and the consideration of two groups of regions. Section 8 concludes.

2 Analytical framework

To conduct our analysis, we use the framework developed in Blanchard and Katz (1992). This framework is derived from a set of equations representative of the average regional labour market within a given economy, and entails the estimation of the following three reduced-form equations:

$$\begin{aligned}
\Delta n_{it} &= \lambda_{i10} + \lambda_{11}(L)\,\Delta n_{it-1} + \lambda_{12}(L)\,u_{it-1} + \lambda_{13}(L)\,pr_{it-1} + \varepsilon_{int}, && (1)\\
u_{it} &= \lambda_{i20} + \lambda_{21}(L)\,\Delta n_{it} + \lambda_{22}(L)\,u_{it-1} + \lambda_{23}(L)\,pr_{it-1} + \varepsilon_{iut}, && (2)\\
pr_{it} &= \lambda_{i30} + \lambda_{31}(L)\,\Delta n_{it} + \lambda_{32}(L)\,u_{it-1} + \lambda_{33}(L)\,pr_{it-1} + \varepsilon_{ipt}, && (3)
\end{aligned}$$

migration.

where n is relative employment (in logs), u is the relative unemployment rate (in %), and pr is the relative participation rate (in logs); L is the lag operator; i stands for region, t for period; the $\lambda's$ are parameters, and the $\varepsilon's$ are residuals.

The term *relative* affecting n, u, and pr indicates the weighted difference of these variables in region i with respect to the national average. More precisely these three variables are defined as:

$$\begin{aligned} n_{it} &= \log(N_{it}) - \hat{\beta}_i \log(N_t), & (4) \\ u_{it} &= U_{it} - \hat{\gamma}_i U_t, & (5) \\ pr_{it} &= \log(PR_{it}) - \hat{\delta}_i \log(PR_t), & (6) \end{aligned}$$

where N, U and PR denote employment, the unemployment rate and the participation rate (all in absolute values); and $\hat{\beta}_i$, $\hat{\gamma}_i$, $\hat{\delta}_i$ account for the regional sensitivity of these three variables with respect to changes in their national counterpart. These estimated parameters are obtained from the following regional regressions:

$$\begin{aligned} \Delta\log(N_{it}) &= \alpha_{1i} + \beta_i \Delta\log(N_t) + \mu_{1it}, & (7) \\ U_{it} &= \alpha_{2i} + \gamma_i U_t + \mu_{2it}, & (8) \\ \log(PR_{it}) &= \alpha_{3i} + \delta_i \log(PR_t) + \mu_{3it}, & (9) \end{aligned}$$

where the $\alpha's$, $\beta's$, $\gamma's$, and $\delta's$ are the parameters to be estimated, and the $\mu's$ are the corresponding residuals.

In order to compute the values of the relative variables n, u, and pr, we follow Decressin and Fatás (1995) and Broersma and van Dijk (2002), who use the residuals from equations (7)-(9). The alternative would be imposing the coefficients of the $\beta's$, $\gamma's$, and $\delta's$ to be equal to 1, and take the relative variables as the difference between the regional and the national variables. This has been the route followed by Blanchard and Katz (1992) and Jimeno and Bentolila (1998). However, in view that our regional estimated equations deliver estimates of these parameters which are significantly different from unity in a number of cases (see Tables 5 and 7), we abstain from imposing additional constraints and proceed as the former studies do.[8]

Another useful piece of information is delivered by equations (7)-(9) in terms of their adjusted R^2's. More precisely, the value of this coefficient indicates the extent to which the pattern of regional employment growth fits the pattern of employment growth at the national level. For example, low values of the adjusted $R^{2\prime}s$ show that most movements

[8]Note that when the null hypothesis of a unit coefficient cannot be refused, the regional specific variable is roughly the same, independently of the chosen method.

in regional employment are not driven by national changes in labour demand. For the whole sample of our analysis (1996-2012), the average adjusted R^2 is 0.48 implying that half of the shocks are common to all regions.

Summing up, this methodology involves, first, the estimation of equations (7)-(9); second, the use of $\hat{\beta}_i$, $\hat{\gamma}_i$, $\hat{\delta}_i$ to compute n_{it}, u_{it}, and pr_{it}; and, third, the estimation of the system of equations (1)-(3).

Once this process is completed, we shock the system with a one-off shift in the residual of equation (1). Such unexpected and temporary shocks on employment growth are the ones evaluated in our analysis. It is important to note that, when we examine the dynamics of these shocks, we are not evaluating the persistence in the average value of the variable under scrutiny (no matter whether this is the employment growth or the unemployment and participation rates). What we are checking, rather, is the speed of convergence of this variable to the previous equilibrium, wherever this one is. This means that these shocks can be formally introduced as positive shocks in all cases, even if their effects are evaluated for a sample period just containing recessive years (as we do later on).

The interpretation of the effects of the shock is made under two common assumptions in the literature. The first one is that unexpected changes in regional relative employment within a year are due to changes in labour demand. This assumption is considered to be correct when most year-to-year unexpected movements in employment are caused by shifts in labour demand rather than shifts in labour supply. This is arguably the case in Spain in the last business cycle.

The second assumption concerns the identification of the shock. It states that employment growth is independent of current changes in the unemployment and participation rates, whereas these rates respond contemporaneously to changes in employment. Jimeno and Bentolila (1998) note that this assumption is more likely to hold when using quarterly data than when using annual data, which is precisely the case here.

When the shock takes place, it gives rise to three adjustment mechanisms, two of them directly arising from the estimated model, and a third one computed as a residual.[9] The first two arise from equations (2) and (3), where the reactions in terms of unemployment and participation rates take place. By construction of the model, whatever employment stimulus not absorbed by a decrease in unemployment or an increase in participation can be ascribed to interregional spatial mobility. Mobility is thus the third adjustment mechanism, and is computed as the difference between the overall employment response to the shock and the unemployment and participation rates responses.

Blanchard and Katz (1992) consider an extension of the basic labour market model

[9] Alecke *et al.* (2010) augment Blanchard and Katz's (1992) model with a migration equation. Unfortunately, data on internal migrations is not available for Spain at quarterly frequencies.

where price responses are also evaluated. This requires the addition, to the system of equations (1)-(3), of the following price equation:

$$w_{it} = \lambda_{i40} + \lambda_{41}(L)\,\Delta n_{it} + \lambda_{42}(L)\,u_{it-1} + \lambda_{43}(L)\,pr_{it-1} + \lambda_{44}(L)w_{it-1} + \varepsilon_{iwt}, \qquad (10)$$

where w_{it} represents three different price variables depending on the estimated model: nominal wages (in our case, the hourly total labour cost), consumer prices (the standard CPI), and house prices.[10] We consider this addition because it sheds new light on the price behaviour in the aftermath of average regional shocks in good and bad times.

We endeavour to examine price responses in Spain for a twofold reason. First, because of the conspicuous link between labour market prices and quantities. It is on this account that we incorporate nominal wages and a price deflator to the analysis. And, second, because the recent evidence on the connection between real state prices and the labour market –see Rogers and Winkler (2013)– calls for such analysis in Spain.

Jimeno and Bentolila (1998) had already dealt with this issue by examining the wage response to local economic conditions in 1983-1988. They found "a low responsiveness of wages to regional economic conditions" in which nominal wages were less flexible than prices. Here, we retake this issue, but sticking to Blanchard and Katz's (1992) methodology.

3 Data and estimation issues

As noted, we first need to estimate equations (7), (8) and (9), and then equations (1), (2) and (3). The first set of equations consist on a time-series estimation by Ordinary Least Squares (OLS), whereas the second set of equations calls for the use of Panel Vector AutoRegression (PVAR) techniques estimated by system GMM. In this way the well-known bias arising from the inclusion of lagged dependent variables in the right side of the equations can be avoided. Estimation of regional-fixed effects in equations (1), (2) and (3) control for intrinsical regional amenities.

3.1 Data

Information on the labour market variables (employment, participation rate, and unemployment rate) is obtained from the Labour Force Survey (*Encuesta de Población Activa*, EPA, from its Spanish acronym). In turn, information on the labour, product and house prices variables comes, respectively, from the Quarterly Survey of Labour Costs (*Encuesta*

[10] Further to the addicion of equation (10), note that consideration of an aditional variable in the model implies the addition of the term $\lambda_4(L)w_{it-1}$ in each of the equations in model (1)-(3).

Trimestral de Coste Laboral, ETCL), the National Statistics Institute (*Instituto Nacional de Estadística*, INE), and the Ministerio de Fomento. Table 2 presents the notation, sources and available sample periods for each of these variables.

The Labour Force Survey underwent a methodological change, in 2002, affecting the definitions of unemployment and participation rates. Since our study departs from 1996 (coinciding with the previous methodological change) to cover at length the last business cycle ($1996-2012$), we need to construct homogeneous series. We do that using the official link coefficients supplied by INE itself, and use the resulting series in our analysis.

Table 2. Definitions of variables.

	Variables	Sources	Time Period
Labour market			
N_{it}	Employment	EPA	$1996q1-2012q4$
PR_{it}	Participation rate	EPA	$1996q1-2012q4$
U_{it}	Unemployment rate	EPA	$1996q1-2012q4$
Prices			
TC_{it}	Hourly total labour costs	ETCL	$2000q1-2012q3$
CPI_{it}	Consumer prices index	INE	$1996q1-2012q4$
HP_{it}	House prices	M. Fomento	$1996q1-2012q4$

From the ETCL we obtain the effective hourly total labour cost TC, which we think it is the relevant variable when examining how sensitive factor prices are to labour market shocks. The TC is the gross cost paid by the employer taking into account any other cost beyond the wage. Note that this variable has a shorter sample size starting in $2000q1$ and finishing in $2012q3$. The variable for prices, CPI, is the standard consumer price index, while we obtain house prices, HP, as the amount of euros per meter-squared corresponding to free houses.

All our variables are disaggregated regionally, have quarterly frequencies, and are seasonally adjusted by using the US X12 Census Bureau process.

3.2 Estimation methodology

The PVAR econometric model to be estimated takes the following reduced form in matrix notation:[11]

$$\mathbf{y}_{i,t} = \Gamma_0 + \Gamma_1(L)\mathbf{y}_{i,t-1} + \varepsilon_{it}, \tag{11}$$

[11] We estimate a PVAR(2). The lag order of the PVAR is chosen to use the maximum sample period available without neglecting the relevance of dynamics. For robustness, the PVAR has been estimated using different lag orders. The results remain roughly the same and are available upon request.

where $\mathbf{y}_{i,t}$ is the vector of endogenous variables (in our case $\mathbf{y}_{i,t} = \Delta n_{it}, u_{ir}, pr_{it}$), $\Gamma_1(L)$ is a matrix of the reduced form coefficients relating past variable values to current values, Γ_0 is a vector of time invariant, region specific effects, and ε_{it} is a vector of idiosyncratic errors.

However, as we estimate a dynamic system of equations where the lagged dependent variables appear as explanatory variables, the use of a standard fixed-effects estimation could yield inconsistent coefficients even if the residuals were not serially correlated. To avoid this problem, we follow Love and Zicchino (2006) and use the "Helmert procedure", which consists in forward mean-differencing the variables to remove the fixed-effects (i.e. this procedure removes only the forward mean of each variable in each period, see Arellano and Bover (1995)). The advantage of this procedure is that it keeps the orthogonality between the transformed variables and the lagged regressors. This allows us to use lagged regressors as instruments to estimate the coefficients by system GMM.

Once the estimation is performed, we compute impulse-response functions (IRFs) describing the reaction of the dependent variables to changes in the innovation of one particular variable in the estimated system. Following the model of Blanchard and Katz this variable is employment growth. We will thus evaluate the dynamics of the labour market to one-off shocks in regional employment.[12]

3.3 Spatial dependence

Regional variables may be liable to spatial correlation. In our study, however, the potential incidence of this problem should be lessened by the fact that national averages are subtracted from the regional values –by equations (4), (5), and (6). Nevertheless, in order to discard that the regional distribution of each variable in the model is not random, and thus causes biased and inconsistent results, we run Moran's I test.

We construct a binary contiguity weighting matrix $\mathbf{W}$ in which the i, j elements (corresponding to the relative position of region i with respect to region j) take value 1, $\bar{w}_{ij} = 1$, if the involved regions share their borders, at least partially; and take value 0, $\bar{w}_{ij} = 0$, otherwise. We then standardise $\mathbf{W}$ so that the rows add up to unity and regions with a small number of borders do not have excessive weights.

For each quarter in our sample, we conduct the two tail version of the test so that the null hypotheses of randomness (i.e., no spatial dependence) is contrasted against the alternative of no randomness (or spatial correlation). Table 3 reports the results we obtain for all variables of interest at the 1% and 5% critical values. The information shown is

[12] All PVAR models in this paper are estimated by using the package provided by Ryan Decker, which is an update of the original package developed by Inessa Love and used in Love and Zicchino (2006).

the number of periods for which the null of randomness is rejected and the corresponding percentage over the total number of quarters in each subsample period.

Table 3. Results on Moran's I test.

	1% critical value				5% critical value			
	1996-2007		2008-2012		1996-2007		2008-2012	
	# periods	%	# periods	%	# periods	%	# periods	%
Δn_{it}	1/47	2.1%	0/19	0.0%	5/47	10.6%	0/19	0.0%
u_{it}	2/48	4.2%	0/20	0.0%	10/48	20.8%	0/20	0.0%
pr_{it}	1/48	2.1%	0/20	0.0%	5/48	10.4%	2/20	10.0%
tc_{it}	1/32	3.1%	0/19	0.0%	3/32	9.4%	2/19	10.5%
cpi_{it}	6/48	12.5%	1/20	5.0%	14/48	29.2%	1/20	5.0%
hp_{it}	6/48	12.5%	1/20	5.0%	19/48	39.6%	3/20	15.0%

Notes: Detailed test results are available from the authors upon request.

Regarding the labour market variables, there is no indication of serious spatial dependence. In the worst case, the unemployment rate variable examined at the 5% critical value in 1996-2007, we cannot reject the null of randomness in 38 out of 48 quarters. Then, with respect to prices, we can safely discard regional correlation at the 1% critical value (even if at the 5% house prices in 1996-2007 deliver less satisfactory results). It is important to note that these results are consistent with the reported findings in Karanassou and Bande (2013b) for labour market variables using annual data between 1980 and 2000.

3.4 Panel unit root tests

Another important issue is the potential presence of unit roots in the variables. Hence, to check the validity of our estimation we have to prove that stationary panel data techniques are appropriate given the integration order of our variables.

To do that, we conduct a series of panel unit root tests. Although it is well-known that the popular individual unit root tests –Dickey-Fuller (DF), Augmented Dickey-Fuller (ADF), and Phillips-Perron (PP) tests– have limited power in distinguishing the null of a unit root from stationary alternatives with highly persistent deviations from equilibrium, it is also generally accepted that the use of pooled cross-section time series data can generate more powerful unit root tests (Levin *et al.*, 2002).

Taking this into account, we conduct a series of panel unit tests to check if the use of stationary panel data estimation techniques is appropriate in our context. We thus carry out the statistic test proposed by Maddala and Wu (1999), which is an exact nonparametric test based on Fisher (1932):

$$\zeta = -2\sum_{i=1}^{N} \ln \pi_i \sim \chi^2(2N),$$

where π_i is the probability value of the ADF unit root test for the ith unit (region in our case).

This test has the following attractive characteristics: (i) it does not restrict the autoregressive parameter to be homogeneous across i under the alternative of stationarity; and (ii) the choice of the lag length and of the inclusion of a time trend in the individual ADF regressions can be determined separately for each region.

Table 4 shows the results of Maddala and Wu's (1999) unit root tests for our 6 variables of interest. The test statistic, ζ, follows a chi-squared distribution, which in our case has a 5% critical value of approximately 49. It is easy to see that all the panel unit root test statistics are greater than the critical value, and the null of a unit root can therefore be rejected at the 5% significance level. It is thus safe to proceed with the analysis by applying stationary panel data techniques.[13]

Table 4. Panel Unit Root Tests.

	$\zeta(\Delta n_{it})$	$\zeta(u_{it})$	$\zeta(pr_{it})$	$\zeta(cpi_{it})$	$\zeta(hp_{it})$	$\zeta(tc_{it})$
96-07	557.23	104.72	122.08	63.88	50.69	
08-12	175.21	73.81	67.49	86.65	76.64	
00-07	319.66	100.10	111.71			237.93
08-12*	202.44	70.98	58.73			176.03

Notes: $\zeta(.)$ is the test proposed by Maddala and Wu (1999). It follows a chi-square distribution whose 5% critical value is 48.6. Lower case letters denote relative variables.

* The data here runs from 2008q1 to 2012q3 due to data limitations in tc_{it}.

4 Aggregate results

We start by presenting an aggregate picture along the lines of previous literature. It comprises the whole sample period, does not yet consider prices, and all regions are taken

[13] Note that we only present a selection of test results, which correspond to some specific periods and do not cover entirely the estimated models in sections 4, 5, 6 and 7. Results for the whole sample period and all regions, and for the whole sample period and the high and low unemployment groups of regions, follow the same pattern. They are available upon request.

into account. These three issues –splitting the sample, considering prices, and grouping the regions– will be faced in subsequent sections.

Table 5 shows the estimated β, γ and δs, together with the corresponding adjusted R^2s for the regional regressions of equations (7), (8), and (9) corresponding to years 1996-2012.

Table 5. Summary of estimates. 1996-2012.

	Equation (7)		**Equation (8)**		**Equation (9)**	
	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$
AND	0.72	1.15	0.94	1.27*	0.98	1.08*
ARA	0.44	0.72*	0.92	0.84*	0.97	1.17*
AST	0.31	0.90	0.87	0.73*	0.95	1.30*
BAL	0.29	0.88	0.78	1.01	0.96	0.89*
CAN	0.46	1.12	0.83	1.35*	0.97	0.96*
CNT	0.34	0.84	0.74	0.73*	0.96	1.17*
CLE	0.65	0.76*	0.94	0.71*	0.98	1.01
CMA	0.64	1.04	0.91	1.11*	0.98	1.31*
CAT	0.81	1.15*	0.98	0.99	0.98	0.71*
VAL	0.74	1.17*	0.96	1.21*	0.98	0.87*
EXT	0.27	0.93	0.89	0.96	0.96	1.08*
GAL	0.29	0.62*	0.81	0.62*	0.95	0.69*
MAD	0.60	0.95	0.97	0.85*	0.98	1.27*
MUR	0.42	1.16	0.97	1.26*	0.98	1.11*
NAV	0.36	0.77	0.92	0.64*	0.95	0.87*
PVA	0.46	0.81	0.54	0.55*	0.95	0.68*
LRJ	0.28	0.92	0.83	0.82*	0.96	1.32*

Notes: * indicates $\hat{\beta}$, $\hat{\gamma}$ and $\hat{\delta}$ are different from unity at a 5% critical value; for the list of regions, see the notes below Figure 2.

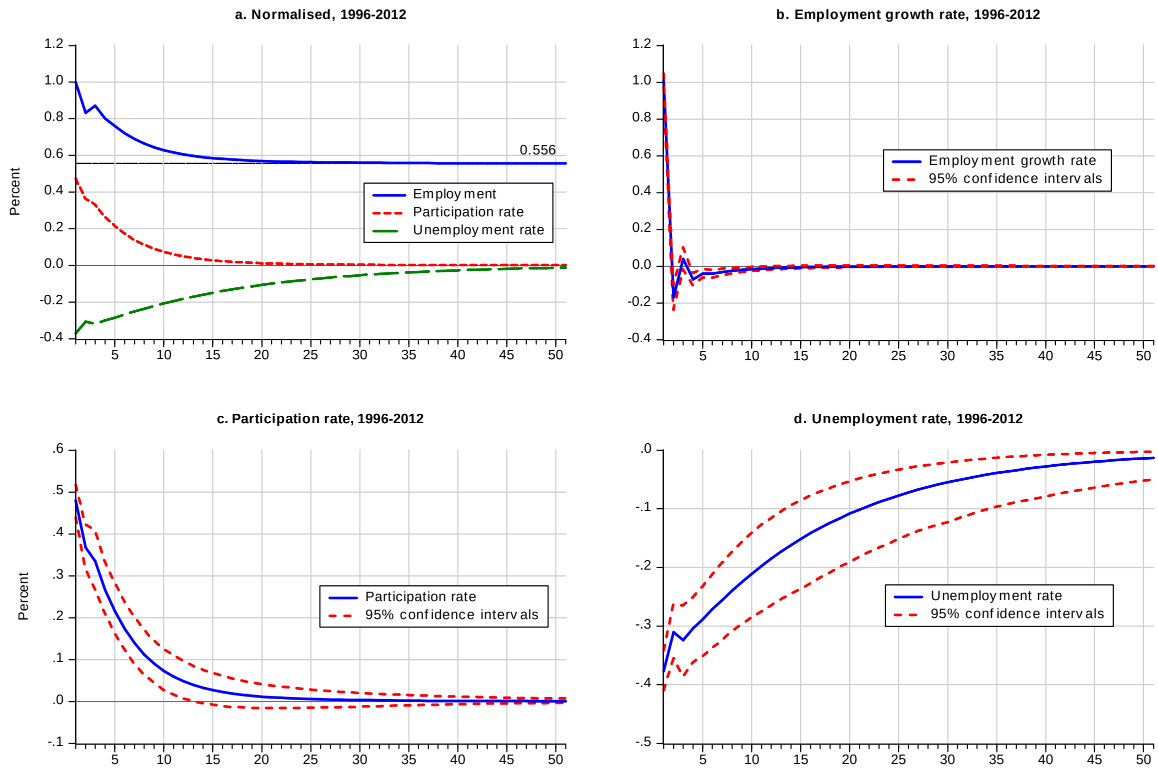
We find the estimates of β to be significantly different from unity (at a 5% critical value) in 5 out of 17 regions. In contrast, the estimates of γ and δ are significantly different from unity in 14 and 16 regions respectively. In view of these results, we choose not to restrict these coefficients.[14] The idea behind this procedure it is not impose extra constraints to the data. We thus create the regional specific variables allowing regions to respond differently to common shocks.

The values of the adjusted R^2s corresponding to equation (7) provide a measure of the relevance of the regional shocks vis-à-vis the national ones. The fact that this value is

[14]Note that Decressin and Fatás (1995) and Broersma and van Dijk (2002) take the same decision with similar results.

below 0.50 in 11 out of 17 territories is an indication that regional-shocks are very relevant for the understanding of the Spanish labour market behaviour.

Figure 3. Aggregate IRFs to a regional employment shock. 1996-2012.



The residuals from these estimated equations are used to examine the aggregate impulse responses to a regional labour demand shock. These are plotted in Figure 3. Figure 3a shows the reaction of the average Spanish region to this shock in terms of employment, and the participation and unemployment rates. Employment converges to 0.556 indicating that 55.6% of the one-off shock is translated into a larger, long-run, relative regional employment level, which is covered by an increase in population (spatial mobility). The rest of the shock is absorbed by the growing participation rates and falling unemployment rates.

It should be noted that Figure 3a provides normalised responses to the shock, while Figures 3b to 3d deliver the original impulse-response functions, together with their standard errors. The reason for normalising is that the IRFs are calculated on one standard deviation shocks and may deliver small divergences from unity. We normalise to ensure comparability across results in next sections. The standard errors are calculated using Monte Carlo simulations with 500 replications.[15]

[15] To conserve space, for the rest of the analysis we only show our results in terms of the normalised

Figure 3a is directly comparable to Figure 2 in Jimeno and Bentolila (1998), and yields a similar picture both in terms of the labour market dynamics generated by the shock, and in terms of its employment effect. They place this effect at 0.40, while Decressin and Fatás (1995) place it at 0.60 for the European Union.

Table 6 provides a detailed comparison of the way a stylised region responds to a regional employment shock today (1996-2012) with its reaction in the past (1976-1994). The main differences observed are the following. First, adjustments via changes in participation rates are much more relevant today than in the past, especially in the short-run (41% versus 23% in the first year). However, there is a much lower persistence today causing the adjustment in the participation rate to fall by 7 percentage points today relative to years 1976-1994. Second, whereas unemployment behaves in a similar way today than in the past (both in magnitude and persistence), spatial adjustments become the most important adjustment mechanism from year 2 onwards. This is due to the low persistence characterising the participation rate response. This change may be related to the increase in the proportion of employment in the service industry and the secular improvement in Spanish educational levels (see Bover and Arellano, 2002).

Table 6. IRFs decomposition to a 1% regional employment shock.

	Spain (1976 − 94)*			Spain (1996 − 2012)		
	Participation	Unemployment	Migration	Participation	Unemployment	Migration
Year 1	23%	36%	41%	41%	37%	22%
Year 2	18%	39%	43%	22%	37%	41%
Year 3	18%	33%	49%	11%	32%	57%

* Results taken from Jimeno and Bentolila (1998), p. 33.

Note: quarterly data aggregated to annual values and normalised by the employment response in the year.

5 Labour market dynamics during the 'wild-ride' and the 'steep-fall' periods

So far we have studied the dynamics of the average region for the whole last business cycle. Some authors, however, have come out with the idea that some Spanish regional unemployment features (e.g. regional disparities) are related with the different phases of the business cycle (Bande *et al.*, 2008). Accordingly, this section aims at examining regional-specific dynamic adjustment mechanisms when a positive shock hits the economy in good and bad times.

impulse-response functions. The original ones can be seen in the Appendix.

Table 7 shows the estimated βs, γs, δs, and the corresponding adjusted $R^2 s$ for the regional regressions of equations (7), (8) and (9) conducted for two subsample periods: 1996-2007, corresponding to the 'wild-ride' of the Spanish economy during those years, and 2008-2012, corresponding to the subsequent steep fall. We acknowledge the fact that these estimates may be sensitive to the sample period length. However, they correspond to such specific and contrasted developments that it is worth examining them individually. Moreover, in order to discard biases, next section provides two robustness checks related to the shortening of the first sub-sample period (to 2000-2007 due to data availability in the total labour costs variable) and to the addition of different price variables to the system. As we will see, the results are robust to these changes.

Table 7. Key estimates for equations (7), (8) and (9).

	1996-2007						2008-2012					
	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$	$\bar{R}^2$	$\hat{\beta}$	$\bar{R}^2$	$\hat{\gamma}$	$\bar{R}^2$	$\hat{\delta}$
AND	0.23	1.39	0.94	1.51*	0.98	0.95*	0.51	0.84	0.99	1.23*	0.47	2.52*
ARA	-0.01	0.24*	0.88	0.62*	0.96	1.16*	0.07	0.46*	0.96	0.87*	-0.04	-0.49
AST	0.06	1.25	0.83	0.74*	0.91	1.17*	0.30	1.18	0.96	0.97	-0.06	-0.04
BAL	0.04	0.98	0.63	0.48*	0.95	0.91*	0.06	0.82	0.93	0.98	0.05	1.20
CAN	0.13	1.45	0.86	0.77*	0.96	0.98	0.28	1.25	0.98	1.16*	0.11	1.94
CNT	-0.003	0.55	0.97	1.18*	0.94	1.24*	0.02	0.35*	0.99	0.77*	-0.05	-0.28
CLE	0.22	0.99	0.98	0.88*	0.97	0.93*	0.41	0.62*	0.99	0.75*	0.24	1.40
CMA	0.03	0.54	0.96	0.71*	0.98	1.21*	0.33	0.92	0.98	1.21*	0.20	1.80
CAT	0.27	1.13	0.97	1.02	0.99	0.77*	0.80	1.46*	0.99	0.99	-0.05	0.20
VAL	0.14	0.90	0.96	1.002	0.99	0.94*	0.56	1.12	0.98	1.14*	0.01	-0.91*
EXT	0.04	1.05	0.89	0.93	0.95	1.04	0.04	0.88	0.90	1.23*	0.11	2.24
GAL	0.03	0.69	0.83	0.71*	0.92	0.65*	0.05	0.44	0.96	0.87*	0.26	1.37
MAD	0.11	0.90	0.95	0.91*	0.98	1.37*	0.38	1.03	0.98	0.76*	0.15	1.11
MUR	-0.002	0.61	0.96	1.15*	0.96	1.08*	0.12	0.87	0.97	1.15*	0.31	1.55
NAV	0.16	1.27	0.86	0.49*	0.94	0.97	0.03	0.59	0.95	0.67*	0.02	-0.72*
PVA	0.11	0.92	0.98	1.10*	0.98	0.79*	0.26	1.04	0.94	0.59*	-0.03	0.42
LRJ	0.07	1.27	0.73	0.46*	0.95	1.43*	-0.06	-0.01	0.91	0.91	-0.03	0.58

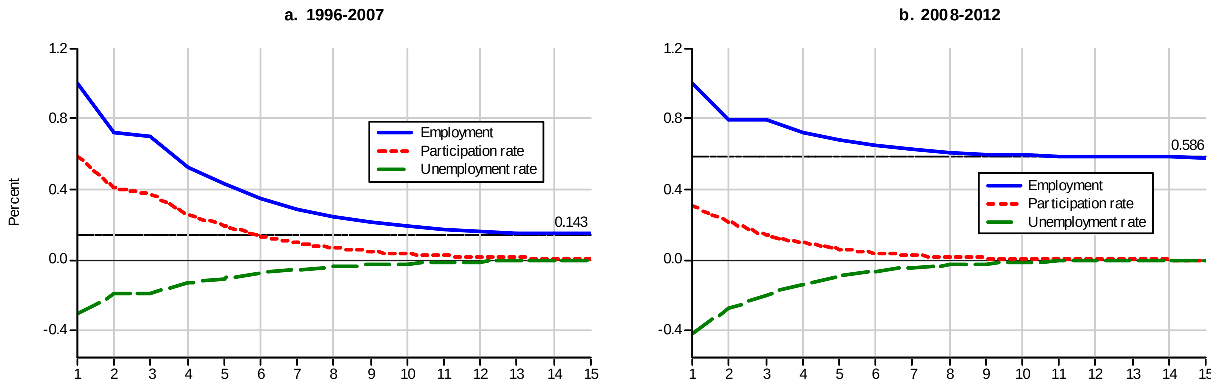
Notes: * indicates $\hat{\beta}$, $\hat{\gamma}$ and $\hat{\delta}$ are different from unity at a 5% critical value; for the list of regions, see the notes below Figure 2.

The first salient finding is the relatively large adjusted R^2s of equation (7) for years 2008-2012, in contrast to the low ones for 1996-2007. This is an indication that nation-wide shocks have become more relevant during the steep-fall years, and is consistent with the fact that Spain was fully caught by the Great Depression, which has driven all regions to an unprecedented slump. The few cases, not infrequent in the literature (see Decressin and Fatás, 1995), in which the adjusted R^2s fall around zero reveal the prominence of regional shocks.

A second noticeable result concerns the change in the participation rate behaviour. There is a stark contrast between the values of the adjusted R^2 in the first period (above 0.90 in all regressions) and the second one (in which they become low). To discard the possibility that this is due to the short sample period of 2008-2012, we shortened the first period to 2003-2007 to take the same length. Since the results remained essentially unchanged, we credit the hypothesis that regional participation rates have radically changed their behaviour. A behaviour that was mainly determined at the national level but has become regionally specific. This finding reinforces our strategy of disaggregating the analysis in subperiods.

As before, we shock the estimated systems and compute the resulting IRFs for each period of analysis. These are shown in Figure 4.

Figure 4. IRFs to a regional employment shock in 1996-2007 and 2008-2012.



The long-run impact of the shock on the relative regional level of employment is smaller in 1996-2007 than in 2008-2012. This reflects the enhanced spatial mobility during the crisis, people being more willing to migrate when there is a regional shock in bad times than when the shock impacts in good times. Despite these differences in the adjustment mechanisms across periods, it is worth noting that the persistence of both the participation and unemployment rates are very similar, taking around 9 quarters for them to converge to the equilibrium. This result contrasts with the larger persistence of these variables (around 15 and 30 quarters) when the whole sample period is examined. This result is mainly due to the analysis of much more homogeneous periods when the full sample is split in expansionary and recessive years, than when both subsamples are considered together (see Altissimo *et al.*, 2009).

Regarding the decomposition of the different components in terms of their influence on the adjustment process, Table 8 shows the large response of relative regional participation rates in 1996-2007. It explains almost 60% of the adjustment in the first quarter, and still accounts for more than 40% of it in quarter 15. In contrast, the immediate response of unemployment dominates in 2008-2012 (more than 40% versus more than 30% the participation rate response), although it vanishes progressively. The picture at the end is one of stark differences across periods. While adjustments during the wild ride are distributed more homogeneously with a dominant participation rate mechanism, during the crisis it is migration, almost exclusively, what leads the adjustment. This reflects people's enhanced willingness to migrate when the shock takes place in bad times.

Table 8. IRFs decomposition to a 1% regional employment shock.

	1996-2007	**2008-2012**
Final employment effect:	0.14%	0.59%
Adjustment in 1st quarter by:		
Participation	58.2%	30.3%
Unemployment	30.9%	42.2%
Migration	10.9%	27.4%
Cumulative adjustment in 15th quarter by:		
Participation	41.3%	9.2%
Unemployment	21.9%	13.2%
Migration	36.8%	77.6%

6 Price responses

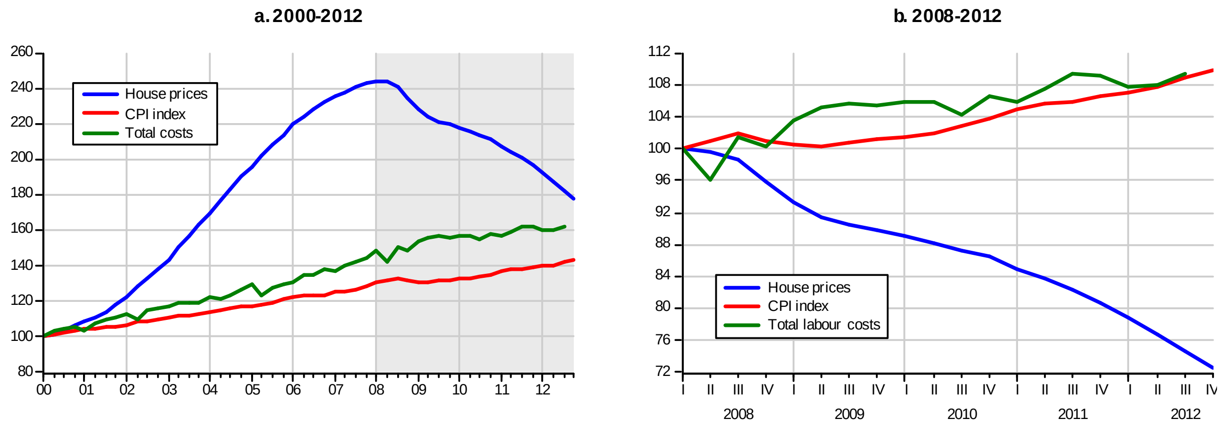
Price rigidities are an important area of interest in macroeconomics and labour economics. In this section we enquire to which extent they are also relevant at the regional level. To outline the most recent developments, Figure 5 shows the average evolution of product market, labour market and house prices in 2000-2012 and 2008-2012.[16]

The first striking feature is the fact that nominal wages (as measured by the hourly total labour cost) have always grown above the CPI with hardly any exception, and even during the most severe crisis years 2011-2012. Between 2000 and 2007, nominal wages

[16] It is worth noting that despite these price evolutions are presented at a national level, their patterns are alike across Spanish regions. The purpose of this figure, therefore, is to inform about the economic context in which the impact of the shock will be studied regarding its price effects. We start in 2000 since the total labour cost series is only available 2000 onwards, and we need full comparability across series.

grew by 40%, consumer prices by 20% and, as a result, real wages grew by around 20%. The second striking feature relates to the evolution of house prices in expansion and recession. During the expansion, their increase amounted to 240% if we take 2000 as the departing year (although considering the full sample, starting in 1996, this rise surpasses 300%). In turn, since 2008 they have decreased by close to 30%. This figure, therefore, shows that nominal wages, consumer prices, and real wages are highly insensitive to the business cycle, while house prices are much more volatile.

Figure 5. Product market, labour market, and house prices. Index 100.

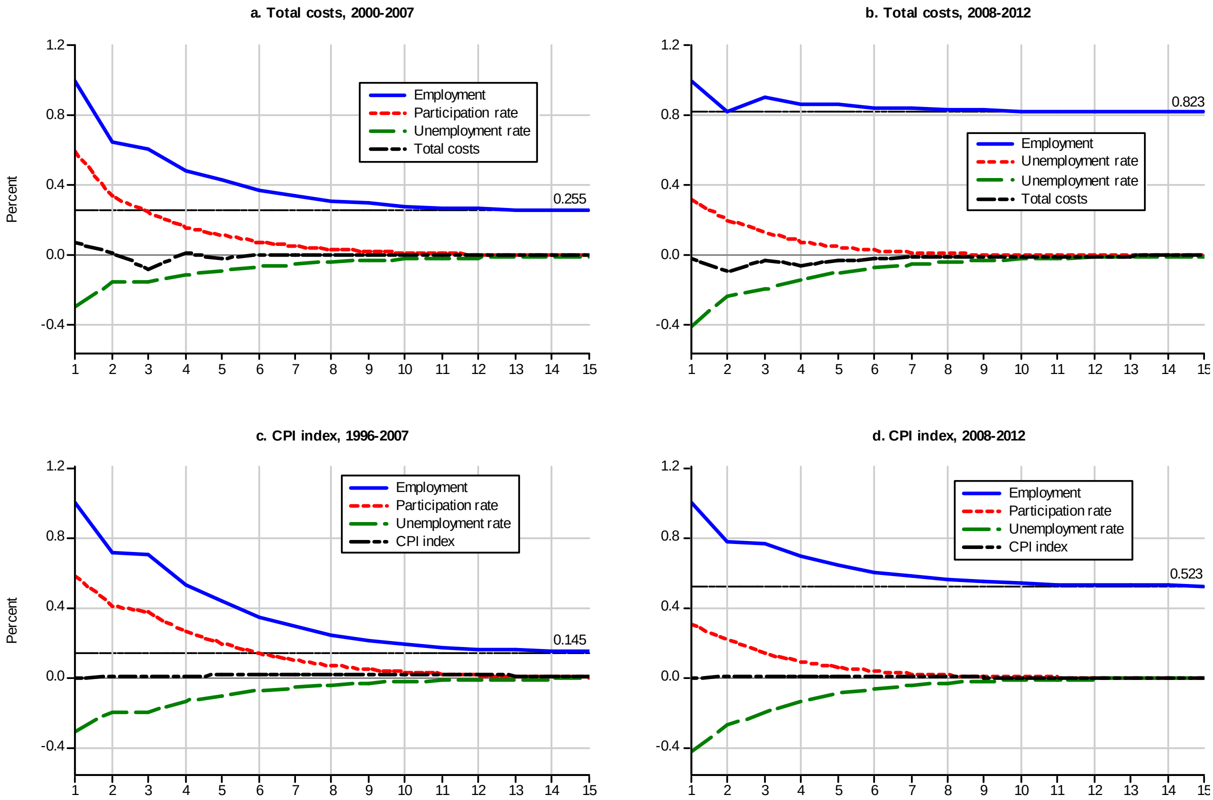


Figures 6 and 7 display the labour market IRFs when the system of equations (1)-(3) is augmented with equation (10) so that prices are included in the analysis.

Figure 6 deals with nominal wages (total labour costs) and consumer prices (CPI index) so that comparison of the two allows to infer the reaction of total compensation in constant terms through the dynamics of its two components. Due to restricted availability of data regarding the total labour costs, we are bound to reduce the system estimation for the first period to years 2000-2007, which implies loosing a third of the sample period (16 quarters of information). This is not the case when using the CPI index and we thus have a natural robustness check. More precisely, comparison of Figures 6a and 6c allows us to discard any significant bias resulting from the shortening of the sample period. This reassures us regarding the results for the second period, which is shorter by definition.

The average regional labour market response across the four different models (Figure 6a to 6d) yields a substantive robust picture. Although consideration of total costs cause larger long-run relative regional employment effects, it is clear that the qualitative response in terms of the three adjustment mechanisms remains very much alike across periods and price variables: asymmetric responses, with participation rates acting as main adjustment channel in expansion, unemployment in recession, and a larger long-run employment level in 2008-2012 pointing to an enhanced willingness to migrate when the shock takes place in troublesome times.

Figure 6. IRFs to a regional employment shock in total costs and product prices.



Regarding prices, we confirm that price rigidities are not just present at the national level, but are also a regional matter. This was outlined in Jimeno and Bentolila (1998) for a sample period running from 1983 to 1988, and for real and nominal wages. Nominal wages have thus been considered as a main source of price rigidities in Spanish regions. Here, we further extend this result by disentangling the real wage in its two components. One of the contributions of our work in this context is the finding that price rigidities cannot be ascribed just to nominal wages, but also to the behaviour of prices.[17]

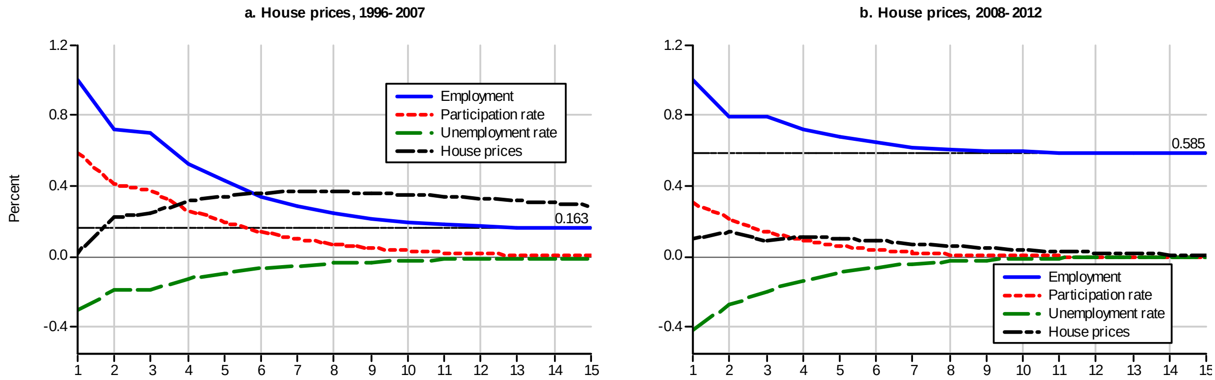
The implication of this finding is that the product and labour market reforms that have been passed since the mid 1990s have been essentially unsuccessful in increasing price flexibility. This result is complementary to the one claiming that labour adjustments in Spain are mainly achieved by adjusting temporary work in a labour market that is neatly segmented since the early 1990s (Dolado *et al.*, 2002) in contrast to other European economies (Sala *et al.*, 2012). And this result, in addition, helps to explain the virtual

[17] For a detailed discussion on how regional specific variables affect prices see Beck *et al.* (2011).

lack of labour market and product market price adjustments, at least relative to those in house prices (as it is next discussed).

The results of a regional labour shock evaluated in the presence of house prices in the estimated system are shown in Figure 7.

Figure 7. IRFs to a regional employment shock in house prices.



When the model includes house prices, the picture is not significantly altered and the labour market adjustment processes remain broadly similar, with all IRFs quickly converging to their equilibrium levels. For employment, this level is 16.3% in the first period, while it reaches 58.5% in the second one. Note that both values fall in between the corresponding ones when the CPI and total labour cost are examined and, hence, the quantitative and qualitative conclusions remain largely unchanged with respect to the labour market response.

Where relevant changes are identified is in the response of prices to the shock. The change in house prices, first of all, is extremely persistent in 1996-2007, and it takes a long period to converge significantly to its equilibrium value (Figure 7a).[18] In contrast, the reaction in the second period is much quicker and full convergence is already achieved within 10 quarters.

This analysis, therefore, discloses a new result for the Spanish economy. In contrast to the extreme stickiness found when total labour costs and consumer prices are considered (in both good and bad times), we now find an asymmetric response across the different phases of the business cycle.

During the expansionary years, there is a significant initial reaction in house prices, which rise continuously during the first 8 quarters after the impact of the shock. This is not the case in the recessive years in which, although there is some reaction, the differences relative to the analysis with product and labour market prices are not significant. In

[18]Figure 7a only shows 15 quarters for consistency with the rest of the figures. Nevertheless, since it displays a very gradual transition to equilibrium, it takes as much as 8 years to reach full convergence.

addition, it can also be observed that the rise witnessed in the boom years lasts for a long time.

Following our findings, the real state sector had a different reaction than the product and labour markets during the wild-ride years. Its developments were not only driven by international and national events (in the form, respectively, of low interest rates and the Land Act of 1998 liberalising land in Spain), but it was also sensitive to specific regional conditions. This could be due to the extraordinary growth path followed by real state activities together with the non-tradable nature of these goods. In any case, this ended up as a housing bubble which bursted during the crisis and caused massive damage to Spanish economic activity.

7 Cluster analysis and regional specificities

Beyond the differences along business cycle phases, a number of studies point out that Spanish regional unemployment could be regionally clustered into high and low unemployment regions –López-Bazo *et al.* (2005), Bande *et al.* (2008), Bande and Karanassou (2009, 2013a, 2013b), and López-Bazo and Motellón (2013). In this spirit, we next test for the existence of significant differences in the regional adjustment mechanisms in high and low unemployment regions when the analysis is performed on regional relative variables (as it is done here, in contrast to the mentioned studies, which focus on absolute regional variables).

7.1 Cluster analysis

We follow Bande *et al.* (2008), and subsequent articles by Bande and Karanassou (2009, 2013a, and 2013b), and proceed in two steps. In the first one we use kernel density function analysis to uncover potential clusters in the Spanish regional relative unemployment rates.[19] In a second step, we perform a k-mean cluster analysis based on exogenous variables to rank each region in one of the groups.

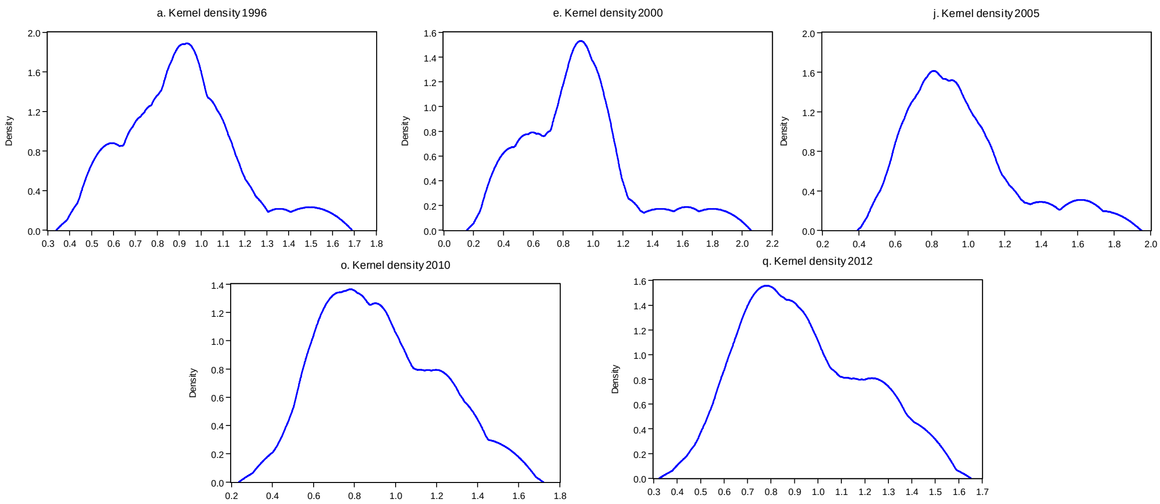
Figure 8 depicts the estimated functions for several years. In 1996, most regional rates were slightly below the national unemployment rate, with two small groups in each extreme of the distribution. Then, as the analysis moves forward in time, two groups become clearly distinguishable. A "low unemployment group", where most of the regions take values around 0.8, and a "high unemployment group" with values around 1.25.

These results are similar to the ones obtained by Bande and Karanassou (2009, 2013a

[19] The regional relative unemployment rates are obtained by dividing each regional unemployment rate over the national one. A unit value, therefore, implies equal regional and national unemployment rates.

and 2013b). Like them, we cluster the Spanish regions in two groups, and proceed to assign each Spanish region to one of them. A tempting procedure would be to allocate the regions according to their regional unemployment rate, given that Figures 2 and 8 would provide immediate allocations. This, however, would generate a purely endogenous classification, which we want to avoid. Hence the second step.

Figure 8. Kernel density functions: relative unemployment rates.



Notes: Densities estimated with an Epanechnikov kernel. Results are robust to different kernel methods.
Source: Spanish Labour Force Survey (EPA).

We conduct a cluster analysis following a k-means procedure.[20] This procedure is based on a selection of exogenous variables, which are used to allocate each region in a single group. We follow the literature and choose two variables closely related to regional social welfare: the participation rate and the relative total labour cost.[21] Table 9 shows the composition of the two groups and the group averages of the two variables examined.

Group 1 is formed by Andalusia, Aragon, Asturias, Balearic Islands, Canary Islands, Cantabria, Castile-Leon, Castile-La Mancha, the Valencian Community, Extremadura, Galicia, Murcia and La Rioja, whereas Group 2 is formed by just four regions: Catalonia, Madrid Community, Navarre, and the Basque Country. It is worth noting that Group 2 comprises regions with larger participation rates and compensations, and lower relative unemployment rates than the regions in Group 1. Thus, from now one we refer to Group 2 as the "low unemployment group" and to Group 1 as the "high unemployment group".

[20] For a detailed description of cluster analysis, see Everitt *et al.* (2001) and Bande *et al.* (2008).
[21] Although it is frequent to choose relative per capita income, our closest measure at hand to a quarterly and recent regional measure of social welfare is total compensation.

Table 9. Composition of groups from the cluster analysis.

Group 1			Group 2	
Andalusia	Castile-La Mancha		Catalonia	
Aragon	Valencian Community		Madrid Community	
Asturias	Extremadura		Navarre	
Balearic Islands	Galicia		Basque Country	
Canary Islands	Murcia			
Cantabria	La Rioja			
Castile-Leon				
	Mean	SD	Mean	SD
Activity rate	0.555	0.049	0.595	0.031
Relative Total labour costs*	0.906	0.065	1.142	0.051
Relative unemployment rate	0.989	0.301	0.715	0.139

Notes: SD=standard deviation; data sources as explained in Table 1.

* Analysis restricted to 2000q1-2012q3 due to data availability.

Once the two groups are identified, the next step is to create regional specific variables for each group in which the reference mean is the one of the group in which each region has been allocated. Then, in order to be consistent with our methodology, we follow Broersma and van Dijk (2002) and regress again equations (7) to (9), but this time considering two average regions: one representing those in Group 1 and another one representing those in Group 2.[22]

7.2 Regional specificities

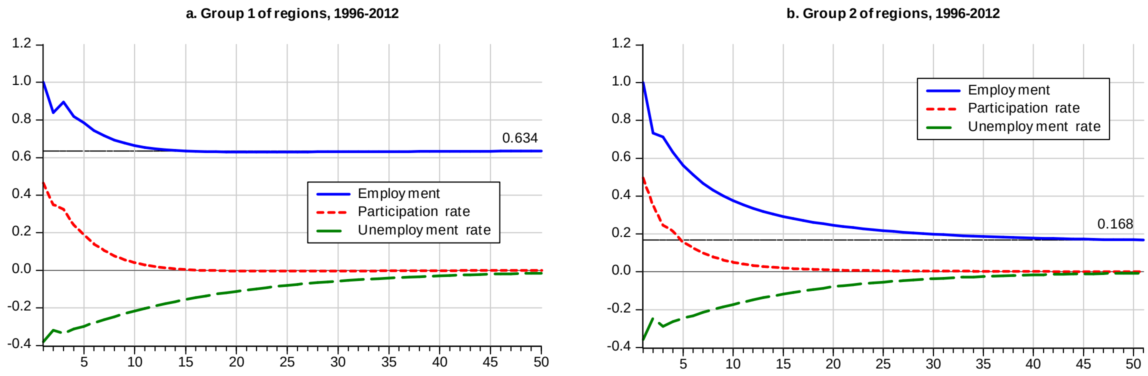
Figure 9 shows the IRFs corresponding to the high and low unemployment rate groups.[23]

The first noticeable result is the similarity in the dynamic response of the unemployment and participation rates in both groups of regions to a regional employment shock. The group of high unemployment regions (Group 1) responds in a similar way than the Spanish regions when considered all together, and achieves a long-run, relative regional employment level that is increased by 63.3% (not so far from the aggregate rise of 55.6%). In turn, the group of low unemployment regions (Group 2) responds in a softer way with a rise of 16.8%.

[22] Note that we do not combine time and regional disaggregation simultaneously. The reason is the tight amount of degrees of freedom in which this analysis is conducted, which would be problematic in case of estimating equations (7) to (9) for Group 2 containing only four regions, for the restricted sample period 2008-2012.

[23] To conserve space, the results of the new estimation of equations (7) to (9) for each group are not reported.

Figure 9. IRFs to a regional employment shock in Groups 1 and 2 of regions.



These results have important implications. If both groups have similar dynamic adjustments in terms of participation and unemployment rates, but people living in "high unemployment" regions are more willing to migrate, this implies that the differences in the relative unemployment rates of these groups should be ascribed to either regional structural differences, to nation-wide shocks producing different impacts, or to both circumstances. Not, in any case, to different responses to one-off regional employment shocks.

This result can be checked, numerically, in Table 7. While in regions with relative high unemployment rates there is a much stronger tendency to leave and there is large spatial mobility, low employment regions are more resilient, and the shock causes larger adjustments in the participation and unemployment rates.

Table 10. IRFs decomposition to a 1% regional employment shock in high and low unemployment regions.

	Group 1 (1996 – 2012)			Group 2 (1996 – 2012)		
	Participation	Unemployment	Migration	Participation	Unemployment	Migration
Year 1	39%	38%	23%	43%	38%	20%
Year 2	17%	37%	46%	23%	45%	32%
Year 3	5%	32%	63%	13%	45%	42%

Note: quarterly data aggregated to annual values and normalised by the employment response in the year.

We find this result consistent with the one we have obtained in Section 5, when splitting the sample period to distinguish the wild-ride from the steep-fall years. We uncovered a stronger tendency to migrate in bad times which, according to these new set of results, may be connected with the fact that people face a relatively bad period (such as the 2008-

2012 one), or with the fact that they live in a region with a relatively high unemployment rate.

8 Conclusions

How does the labour market of the average Spanish region respond to an employment shock? Is this response symmetric across business cycle phases? How do prices adjust in response to such shocks? Do regions react alike in spite of their unemployment rate differences? These are the central questions we have tried to answer.

Our aggregate analysis shows that persistence in the Spanish regional labour market is not substantially different today than in 1976-1994, as documented by Jimeno and Bentolila (1998). Although the Spanish labour market may be, on aggregate, much more flexible than it was, there has been little progress in terms of regional unemployment persistence (recall Figure 2b). The main difference lies in the larger adjustment, today, through changes in participation rates, although with a lower degree of persistence. This gives room to a more important role of interregional migration in the long-run, which may be related to the change in the industry composition, and the availability of a larger amount of educated workers in the Spanish labour market, as shown in Bover and Arellano (2002).

We also find evidence of different responses across business cycle phases. The main mechanism to adjust to a regional shock in expansion is the change in the participation rate, whereas unemployment and migration become the central ones in recession (in the short- and long-run, respectively).

Another finding is that the long-run relative employment level in the aftermath of the shock is higher in 2008-2012 than in 1996-2007. This is an outcome of the growing relevance, in the second period, of the migration adjustment mechanism to regional shocks. Although it is well-known that the Spanish labour market is not characterised by a high degree of regional mobility since the 1970's, our results provide support to the hypothesis that people have become more willing to migrate when confronted to a shock that takes place in a recessive context than when confronted to an equivalent shock taking place in good times.

We also find that price stickiness is still very strong nowadays (1996-2012), a result that should come as no surprise since this was already documented by Jimeno and Bentolila (1998). Our contribution on this regard is that strong real wage rigidities arise both from nominal wages and consumer prices, and are present both in expansion and recession. This seems to indicate that the product and labour markets are still operating with a substantial degree of imperfect competition, in which case policy measures implemented to foster competition and a larger responsiveness of market prices to the changing (regional) economic environment have, to a large extent, failed to achieve their target. Regarding house prices, we have also found a very persistent reaction when there is a regional labour demand shock during the expansionary period.

Disaggregation of the Spanish regions by groups uncovers very similar responses in the short-run, and some divergence as time goes by. In the long-run, high unemployment regions are more reactive in terms of spatial mobility. This reveals a larger propensity to migrate from regions with high than from regions with low unemployment rates. It follows that differences in regional unemployment need to be explained by factors other than regional labour market dynamics. For example, by differences in regional amenities, or different responses to nation-wide shocks.

In any case, we need time to reduce our unemployment rate –participation rate is the main adjustment mechanism in good times–, but we are quick in rising it –unemployment becomes the most important adjustment channel in troublesome times. Given the current difficulties in reducing unemployment, this is an asymmetry that will surely receive attention in future work.

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Appendix

Figure A1. IRFs to a regional employment shock. 1996-2007 and 2008-2012.

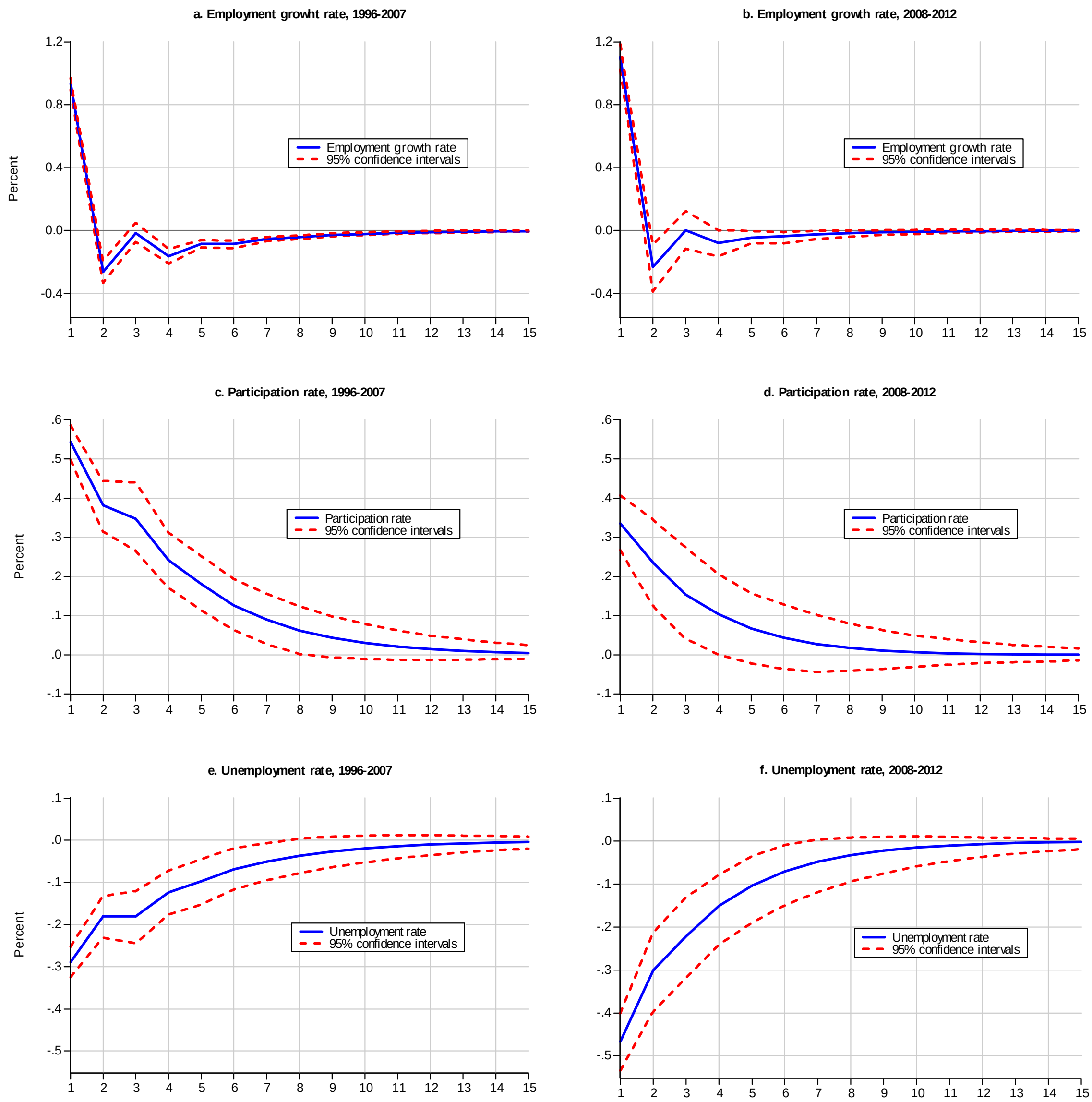

Figure A2. IRFs to a regional employment shock. Total costs, 2000-2007.

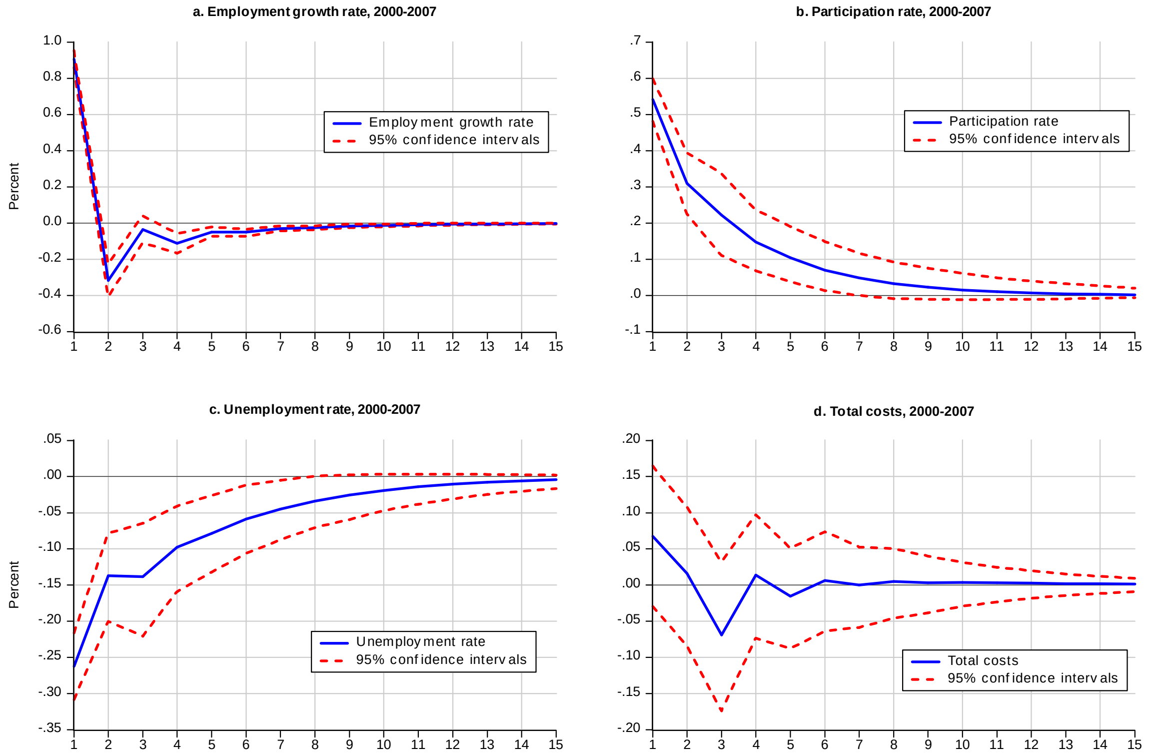

Figure A3. IRFs to a regional employment shock. Total costs, 2008-2012.

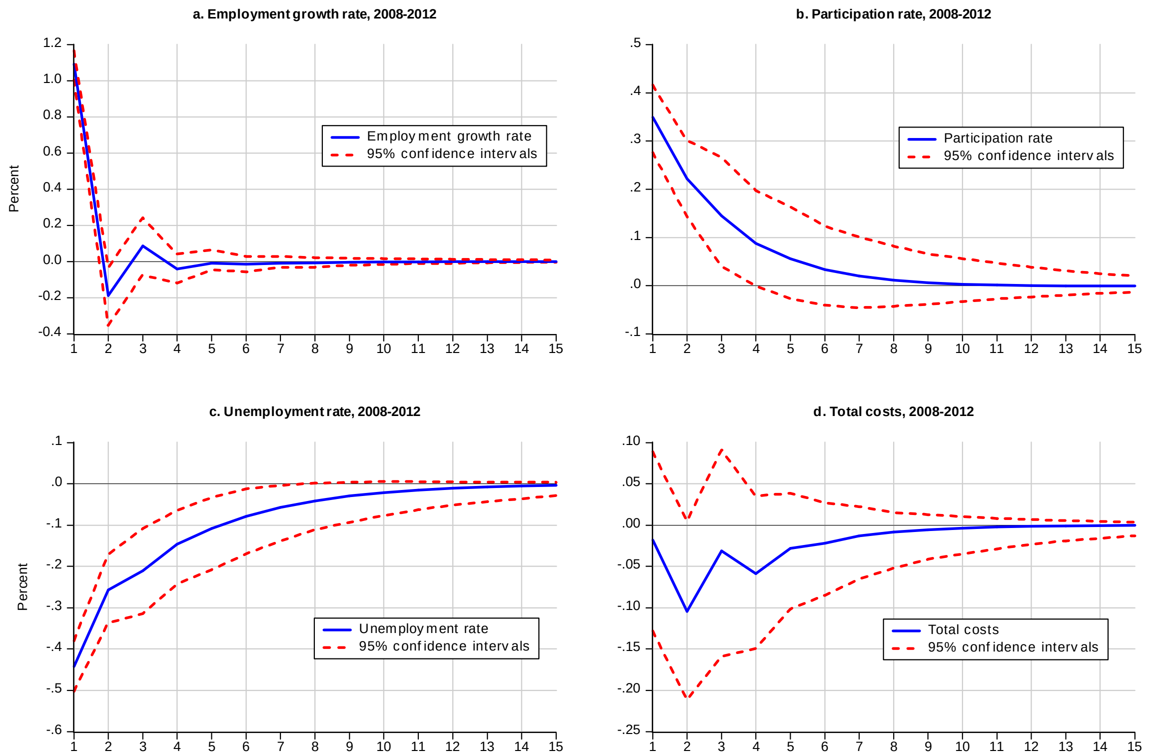

Figure A4. IRFs to a regional employment shock. CPI index, 1996-2007.

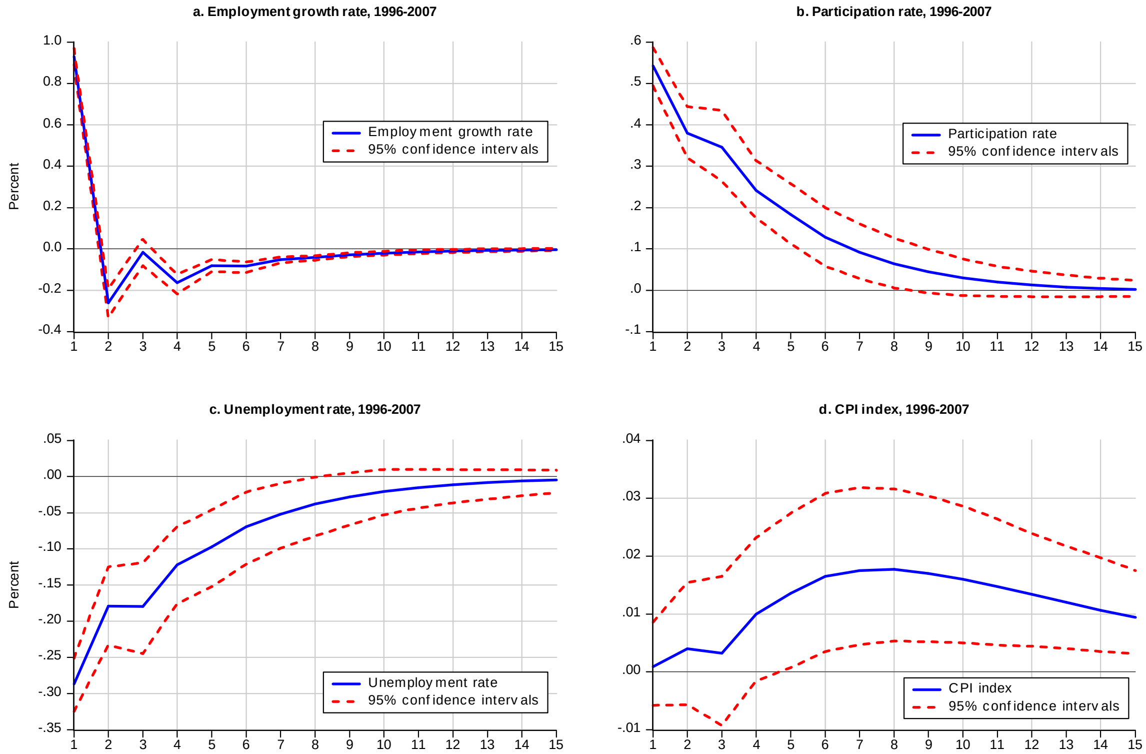

Figure A5. IRFs to a regional employment shock. CPI index, 2008-2012.

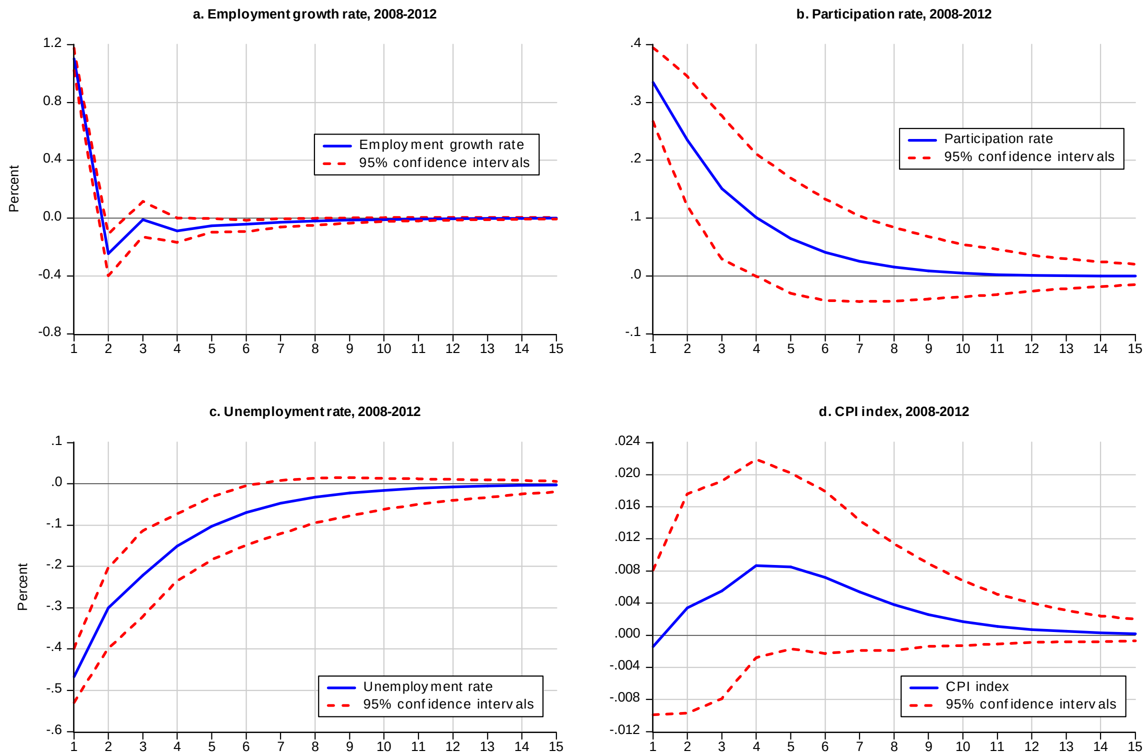

Figure A6. IRFs to a regional employment shock. House prices, 1996-2007.

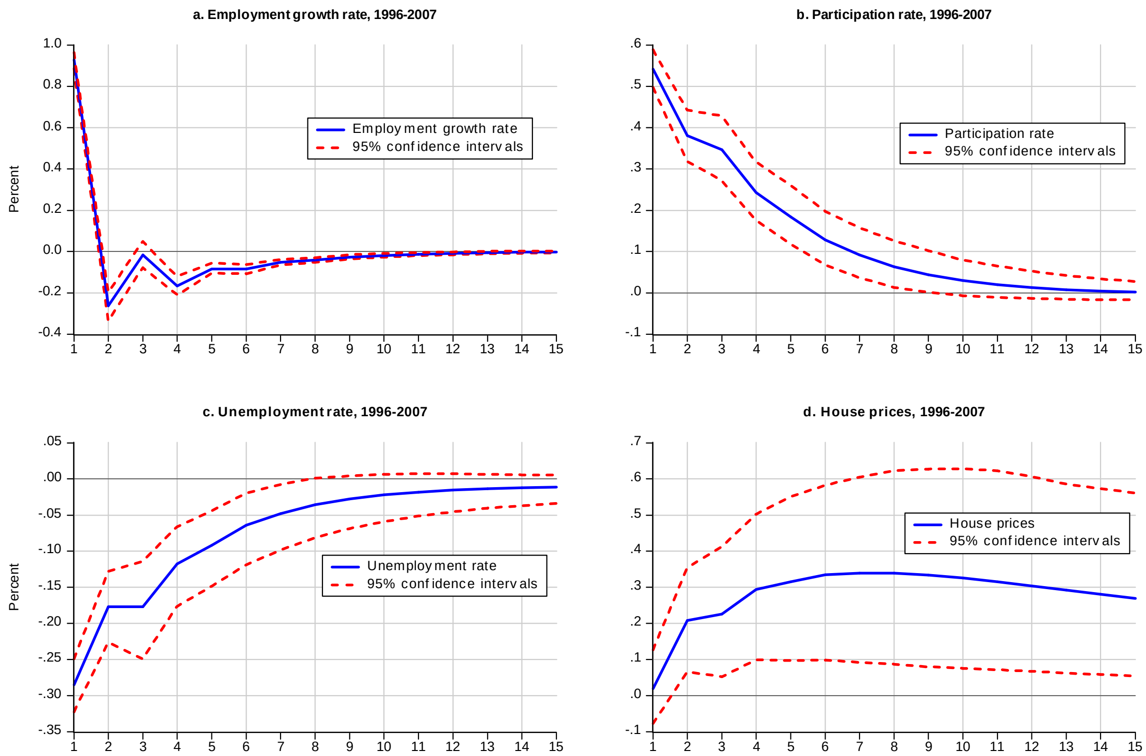

Figure A7. IRFs to a regional employment shock. House prices 2008-2012.

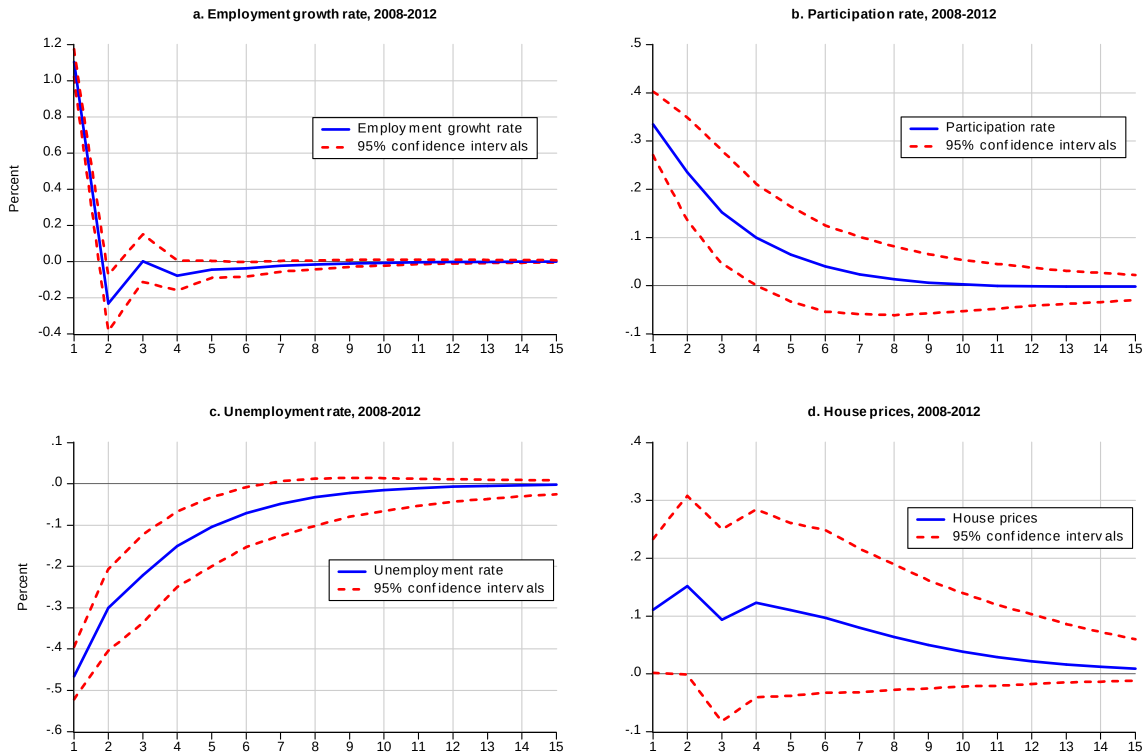

Figure A8. IRFs to a regional employment shock. Cluster analysis, 1996-2012.