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ORIGINAL PAPER

# On the persistence of Spanish unemployment rates

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**Abstract** This paper tests the hysteresis hypothesis for the unemployment rate of the Spanish regions over the period 1976–2004. For that purpose, we employ a large battery of recently developed panel tests which explicitly control for cross-sectional correlation in addition to the panel stationarity test of Carrión-i-Silvestre et al. (Econom J 8:159–175, 2005) which allows for multiple structural breaks and cross-sectional dependence. Overall, our confirmatory analysis with three different proxies for the excess of labour supply renders strong support for the hysteresis hypothesis in regional Spanish unemployment. The results are robust across panel techniques and datasets and accord well with the common belief among scholars that attaches a high degree of persistence to Spanish unemployment due to labour market malfunctioning. We provide a detailed description of the clusters of breaks identified in the analysis, which appear to be closely associated with some macroeconomic shocks and institutional arrangements.

**Keywords** Unemployment rate · Persistence · Panel stationarity test · Structural breaks · Cross-dependence

## 1 Introduction

The difficulties in interpreting the high and persistent unemployment levels recorded in the developed world, and mainly in Europe since the first oil shock of the early seventies have led macroeconomists and labour economists to reconsider the core of

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the natural rate theory.[1] This traditional strand of the literature was pioneered by the work of Friedman (1968) and Phelps (1967, 1968) who entertain the hypothesis that the unemployment rate fluctuates around some natural or equilibrium rate which is associated with a fully equilibrated labour market. Hence, shocks to unemployment are temporary, which implies that unemployment will revert to its equilibrium level.

The labour market developments of the seventies and eighties led to a refinement of the natural rate theory, which gave rise to the structuralist hypothesis. According to these authors (see Phelps 1994; Phelps and Zoega 1998), the rise in unemployment over the seventies and eighties was caused by changes in structural factors leading to the adjustment to a higher equilibrium rate of unemployment. Therefore, most shocks to unemployment appear to be temporary, but occasionally the natural rate permanently changes as a result of infrequent and large shocks. Several major postwar macroeconomic shocks have been identified (Phelps 1994). However, authors have shown some dissatisfaction with the structuralist explanation of the rise in unemployment, since the US unemployment rate—which has also been hit by these shocks—has reverted back to its pre-shock level rather than steadily rising as in Europe. In addition, over the past two decades many European countries have devoted substantial effort to deregulate their labour markets. However, unemployment rates have not significantly fallen as one would expect with the structuralist hypothesis.

These developments gave rise to the hysteresis hypothesis which advocates that differences in institutional arrangements governing the functioning of labour markets across countries lead to marked differences in the way economies adjust to macroeconomic shocks.[2] So labour market rigidities may be responsible for the sluggishness of European labour markets in adjusting to adverse shocks. As opposed to the natural rate and structuralist hypotheses, the hysteresis hypothesis implies that the unemployment rate is path-dependent, i.e. current unemployment levels heavily depend on past levels.[3] Hysteresis' defenders claim that the key to the rise in European unemployment should not be sought on the occurrence of adverse shocks to unemployment, but rather, in the way countries adjust to these shocks.[4] Since unemployment rates can remain at a new higher level indefinitely even if the adverse shock has ended, it is advisable to implement policy measures aimed at increasing labour market flexibility. This literature is also very concerned with restrictive demand policies.

In this context, unit root and stationarity tests have been widely used to investigate the dynamic behaviour of unemployment rates in an attempt to discriminate among these competing theories. Essentially, the hysteresis hypothesis is formulated as a unit root process, and its rejection gives support to the natural rate theory if no breaks are

[1] See Blanchard (2006), among others.

[2] See Blanchard and Summers (1986); Blanchard and Summers (1987).

[3] Two versions of the hysteresis hypothesis exist: (1) the full hysteresis case implies that every temporary shock affects unemployment permanently, thus making the traditional concept of natural rate irrelevant (see Karanassou and Snower 1997); (2) the partial hysteresis or persistence case, which is a special case of the natural rate theory, entails that temporary shocks to unemployment have long-lasting but not permanent effects. Full hysteresis can be modelled as a unit root process while persistence as a near unit root.

[4] See Blanchard and Wolfers (2000).

included in the specification or the structuralist view if mean shifts in unemployment are allowed for.

Chronologically, we find three main groups of studies on the basis of the type of unit root and stationarity tests used to investigate the hysteresis hypothesis. In the first place, there is a group of studies using traditional unit root tests, basically of the Augmented Dickey–Fuller (ADF) type—see, for instance, Blanchard and Summers (1986), Alogoskoufis and Manning (1988), Andrés (1993), Jaeger and Parkinson (1994), Roed (1996) and Everaert (2001) for OECD countries.[5] Mostly, the evidence supports the hypothesis of hysteresis for the EU economies and is mixed for the US. However, those conclusions are based on unit root tests that under the alternative assume a constant, unique, natural rate of unemployment. As noted by Perron (1989) and Zivot and Andrews (1992) among others, unit root tests that fail to control for structural breaks are biased towards the non-rejection of the null of nonstationarity. To overcome this limitation, Mitchell (1993), Bianchi and Zoega (1998), Arestis and Biefang-Frisancho Mariscal (1999, 2000), Papell et al. (2000), Everaert (2001) and Ewing and Wunnava (2001) apply unit root tests that allow for structural breaks in the unemployment rate of OECD countries. Once the specification includes structural breaks, the null hypothesis of hysteresis tends to be rejected in favour of the alternative of stationarity around a changing equilibrium rate for the majority of the countries analysed, and particularly for the US which usually shows a low degree of unemployment persistence. That finding seems to be in accordance with the structuralist theories of unemployment.[6]

A third group of studies are based on panel unit root tests, which try to exploit the cross-sectional variation of the series. The most commonly used panel unit root tests for the case of no breaks are the tests of Levin et al. (2002) and Im et al. (2003); and in the case of breaks the test of Im et al. (2005) that allows for up to two mean shifts. In general, studies applying these techniques to unemployment rates for OECD countries, without allowing for structural breaks (Song and Wu 1998; León-Ledesma 2002; Camarero and Tamarit 2004) and, mainly, when structural breaks are allowed for (Murray and Papell 2000; Strazicich et al. 2001; Camarero et al. 2006), tend to reject, even stronger than in the case of univariate tests allowing for structural breaks in the individual series, the null hypothesis of hysteresis in favour of the alternative of stationarity around a changing equilibrium rate. To sum up, the inclusion of structural breaks, reinforced with the panel dimension leads to the rejection of the hypothesis of hysteresis.

In this paper, we focus on the analysis of the dynamic behaviour of Spanish unemployment at the regional level. The high and persistent level of Spanish unemployment which has sustained over the past three decades can be seen as one of the most striking cases for the hysteresis hypothesis and has drawn a great attention among scholars in the field (see Blanchard and Jimeno 1995). In order to conduct unit root testing with

[5] More specifically for the case of Spain, Alogoskoufis and Manning (1988), Andrés (1993) and Everaert (2001) find evidence favouring the nonstationarity of the unemployment rate, while Roed (1996) provides mixed evidence.

[6] More particularly for the Spanish case, Arestis and Biefang-Frisancho Mariscal (2000), Papell et al. (2000) and Carrión-i-Silvestre et al. (2004) find evidence of hysteresis even after accounting for structural instability, while Arestis and Biefang-Frisancho Mariscal (1999) and Everaert (2001) find support for regime-wise stationarity accompanied with a high degree of persistence.

"state-of-the-art" panel unit root and stationarity tests, we adopt a regional approach which allows us to investigate the time-series properties of the unemployment rate of the 17 Spanish regions over the period 1976–2004 using three different proxies for the excess of labour supply.[7]

The main contributions of the paper are in the following directions. Firstly, we apply a large battery of recently developed panel tests which explicitly control for cross-dependencies and show better power and size properties than other tests previously used in the literature. Secondly, we make use of the recently developed panel stationarity test of Carrión-i-Silvestre et al. (2005, CBL hereafter) which allows for a highly flexible trend function by incorporating an unknown number of structural breaks. This test is thus more general than the one developed by Im et al. (2005) that only allows for two level shifts.[8] Thirdly, by taking stationarity as the null hypothesis rather than nonstationarity, we can be more confident of the presence of a unit root in unemployment as the null will tend to be rejected when there is strong evidence against it. Fourthly, in contrast to previous studies, we carry out a formal analysis of the presence of cross-sectional dependence in unemployment innovations by applying the test for cross-dependence recently developed by Pesaran (2004). In addition, we allow for more general forms of cross-sectional dependence by computing the bootstrap distribution of the panel stationarity test with multiple breaks. Overall, our confirmatory analysis renders strong support for the hysteresis hypothesis in Spanish regional unemployment. This appears robust to the panel technique and the proxy for excess of labour supply used in each case.

The remainder of the paper is structured as follows. Section 2 describes the data and the methodology of the panel unit root and stationarity tests employed in the paper. Section 3 presents the results from the analysis of the time series properties of the unemployment rates of the Spanish regions using three different proxies, focuses on the interpretation of the breaks and puts forward some policy implications. Finally, Sect. 4 concludes.

## 2 Data and econometric methodology

### 2.1 Data description

In our empirical analysis we employ three proxies for the excess of labour supply in the 17 Spanish regions. In the first place, we employ quarterly data on unemployment rates obtained by dividing the number of unemployed by the corresponding labour

[7] At the Spanish regional level, there is only one study related to our approach. Carrión-i-Silvestre et al. (2001) investigate the hysteresis hypothesis for the 50 Spanish provinces using annual data that cover the period 1964–1997. By using the Harris and Tzavalis (1999) test extended to allow for a common shift in the mean, their main conclusions point to the rejection of the full hysteresis hypothesis in favour of regime-wise stationarity associated with a low unemployment regime in the sixties and early seventies and a high unemployment regime from the mid-eighties onwards. Our approach improves on this contribution in several dimensions: data sources, number of breaks and the treatment of cross-sectional dependence.

[8] In previous work, we followed the approach in Im et al. (2005) which includes a linear time trend in the specification. Since this may be problematic for the structuralist approach, we prefer not to report the results.

force figure over the period 1976(3)–2004(4). These data come from the Spanish Labour Force Survey (*Encuesta de Población Activa,* EPA), which are provided by the National Statistical Office (*Instituto Nacional de Estadística*). In the second place, we employ monthly data on unemployment rates based on registered or administrative unemployment figures for the period 1976(7)–2004(12). These data are supplied by the Spanish Labour Office (*Instituto Nacional de Empleo,* INEM). In addition, we use a third proxy for the excess of labour supply also obtained from the INEM for the period 1977(1)–2004(12), which focuses on the total number of persons looking for a job (*demandantes totales de empleo*). This is a broader category than unemployment since it includes on-the-job search, on-the-study search, agrarian workers receiving a special unemployment subsidy, among others. Henceforth we denote this broader definition as "job-seekers rate". All variables have been seasonally adjusted.[9]

It is important to note that we complement the empirical analysis using the EPA source with the INEM data for several reasons. First, it is the only official source that provides monthly data, so the series can be longer than those obtained from the EPA which are supplied on a quarterly basis; this fact can be relevant for the type of analysis (persistence) that we implement in this paper. Second, it is a data source that has been used less frequently than the EPA or other international databases, so it is interesting to pay attention to this new line of results. Third, this data source provides two proxies for the excess of labour supply, which allow us to check for the robustness of our findings.

## 2.2 Test for cross-sectional dependence in panels

Pesaran (2004) presents a simple test of error cross-sectional dependence which is based on the average of pair-wise correlation coefficients of ordinary least squares (OLS) residuals obtained from standard ADF regressions for each individual. Let $\hat{\rho}_{ij}$ be the sample estimate of the pair-wise correlation coefficient of OLS residuals:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{T} e_{it} e_{jt}}{\left(\sum_{t=1}^{T} e_{it}^2\right)^{1/2} \left(\sum_{t=1}^{T} e_{jt}^2\right)^{1/2}} \tag{1}$$

where $e_{it}$ represents the OLS estimated residuals for individual $i$. On the basis of pair-wise correlation coefficients, Pesaran (2004) proposes a test of cross-sectional dependence with good finite-sample properties given by:

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left( \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right) \xrightarrow{d} N(0, 1) \tag{2}$$

[9] We keep up with the extant literature in the field by using seasonally adjusted data rather than unadjusted data. Our main results remain unchanged when we employ seasonally unadjusted data and are available from the authors upon request.

The CD statistic tests the null of cross-sectional independence and is distributed as a two-tailed standard normal distribution. We anticipate that Pesaran's statistic points to the existence of cross-sectional correlation in the three datasets analysed. This motivates the use of the panel unit root and stationarity tests outlined below, which explicitly allow for cross-sectional dependence.

### 2.3 Panel tests under cross-sectional dependence and no breaks

Smith et al. (2004) provide panel versions of some powerful modifications of univariate ADF $t$-statistics such as the Max test of Leybourne (1995) and the weighted symmetric (WS) test of Pantula et al. (1994). Smith et al. (2004) consider a panel specification of the form:

$$\Delta y_{it} = \alpha_i + \gamma_i y_{it-1} + \sum_{j=2}^{p_i} \delta_{ij} \Delta y_{i,t-j-1} + \varepsilon_{it}. \tag{3}$$

where $p_i$ is the required degree of lag augmentation to make the residuals white noise, $\alpha_i$ represents the country-specific fixed effects, and $i = 1, \ldots, N$ and $t = 1, \ldots, T$ stand for the number of panel members and time periods, respectively. To achieve the most parsimonious model compatible with white noise residuals, $p_i$ is determined by the conventional step-down procedure.

Leybourne (1995) proposed to obtain the ADF $t$-statistic from original data $(DF_{fi})$, and from time-reversed data $(z_{it} = y_{i,T+1-t})$ yielding $DF_{ri}$. The Max $t$-statistic for individual $i$ is obtained as $Max_i = Max(DF_{fi,} DF_{ri})$. In a panel framework, the panel Max $t$-statistic takes the form:

$$\Psi_{Max} = \frac{\sqrt{N}(Max_{NT} - E(Max_i))}{\sqrt{Var(Max_i)}} \tag{4}$$

where $Max_{NT} = N^{-1}\sum_{i=1}^{N} Max_i$. Likewise, individual WS tests are computed as in Pantula et al. (1994), and the panel counterpart is given by:

$$\Psi_{WS} = \frac{\sqrt{N}(WS_{NT} - E(WS_i))}{\sqrt{Var(WS_i)}} \tag{5}$$

where $WS_{NT} = N^{-1}\sum_{i=1}^{N} WS_i$. Finally, Smith et al. (2004) present a more powerful variant of the Lagrange Multiplier (LM) statistic on the basis of forward and reverse ADF regressions which yield the univariate $LM_{fi}$ and $LM_{ri}$. Since both statistics take a positive value, the minimum LM statistic is computed as $Min_i = Min(LM_{fi,} LM_{ri})$. The panel version of the test is as follows:

$$\Psi_{Min} = \frac{\sqrt{N}(Min_{NT} - E(Min_i))}{\sqrt{Var(Min_i)}} \tag{6}$$

where $Min_{NT} = N^{-1}\sum_{i=1}^{N} Min_i$. $\Psi_{Max}$ and $\Psi_{WS}$ reject the null hypothesis for large negative values of the statistic and $\Psi_{Min}$ rejects the null for large positive values.[10] Given that these tests assume both cross-sectional independence and asymptotic normality, Smith et al. (2004) develop a modified bootstrap procedure to compute $p$-values that are robust to small-sample bias and to general forms of cross-sectional dependencies across panel members.[11]

Breitung and Das (2005) consider a model like (3) while assuming the existence of weak cross-sectional dependence. For that purpose, they write the model as a seemingly unrelated-type system of equations in matrix form:[12] $\Delta y_t = \phi y_{t-1} + \varepsilon_t$, where $\Delta y_{t,} y_{t-1}$ and $\varepsilon_t$ are $N \times 1$ vectors. The cross-sectional correlation is represented by a non-diagonal covariance matrix $\Omega = E(\varepsilon_t \varepsilon_t')$ for all $t$, with bounded eigenvalues. Breitung and Das (2005) demean the data such that $\tilde{y}_t = y_t - y_0$, where $y_0$ represents the value of the initial observation, and estimate consistently the variance–covariance matrix of the OLS estimator, which is denoted by $\hat{\upsilon}_{\hat{\varphi}}$. They then obtain the robust $t$-statistic free of size distortions due to contemporaneous cross-sectional correlation for $N$ and $T$ tending to infinity:

$$t_{rob} = \frac{\hat{\phi}}{\sqrt{\hat{\upsilon}_{\hat{\varphi}}}} = \frac{\sum_{t=1}^{T} \tilde{y}_{t-1}' \Delta \tilde{y}_t}{\sqrt{\sum_{t=1}^{T} \tilde{y}_{t-1}' \hat{\Omega} \tilde{y}_{t-1}}} \xrightarrow{d} N(0, 1) \tag{7}$$

Harris et al. (2005, HLM) propose a panel stationarity test that is able to handle time-series and cross-sectional dynamics, thereby allowing for heterogeneity in the deterministics across units. They propose a test that addresses cross-sectional dependence through a factor model with an unknown number of factors like:

$$\begin{aligned} y_{it} &= \beta_i' x_{it} + z_{it} \\ z_{it} &= \lambda_i' f_t + e_{it} \end{aligned} \tag{8}$$

where $f_t$ is an $r \times 1$ vector of latent factors which needs to be estimated to determine the rank, $\lambda_i$ is an $r \times 1$ vector of loading parameters and $e_{it}$ is the idiosyncratic term for each $i$. They further assume that $f_t$ and $e_{it}$ are mutually independent of one another. They then compute the $\hat{S}_k^F$ test for the estimated components $\hat{f}_t$ and $\hat{e}_{it}$ jointly, which is robust to cross-sectional correlation and serves as a test for the null hypothesis that the series $z_{it}$ are stationary for all $i$.[13] More specifically, the resulting statistic takes

10 All the three tests take a unit root in all individuals as null hypothesis versus the alternative of stationarity for at least one individual panel member.

11 See Smith et al. (2004, pp. 165–166) for details on the bootstrap procedure that generates bootstrap innovations through resampling using a block size of 30 and 20,000 replications. The maximum lag order of autocorrelation used to compute the statistics is set at eight. The results are robust to the use of a block size equal to 100 and different values for the maximum lag order.

12 For expositional simplicity we abstract from lagged augmented terms.

13 The null hypothesis implies that all cross-sectional units are stationary against the alternative that at least one unit is nonstationary.

the form:

$$\hat{S}_k^F = \frac{\tilde{C}_k}{\hat{\omega}\{\tilde{a}_{k,t}\}} \xrightarrow{d} N(0,1) \tag{9}$$

where $\tilde{z}_{it}$ are standardised residuals, $\hat{\omega}^2\{\tilde{a}_{k,t}\}$ is the long-run variance estimator, $\tilde{C}_k = T^{-1/2}\sum_{t=k+1}^{T} \tilde{a}_{k,t}$ and $\tilde{a}_{k,t} = \sum_{i=1}^{N} \tilde{z}_{it}\tilde{z}_{it-k}$. It can be shown that $\hat{S}_k^F$ follows a standard normal distribution even when it is based on residuals with large $T$ and fixed $N$.

### 2.4 Panel stationarity test with multiple structural breaks and cross-sectional dependence

The panel stationarity test developed by CBL allows for an unknown number of structural breaks under the null hypothesis of joint stationarity and is a generalisation of Hadri (2000)'s test for the case of multiple changes in level and slope. Let the unemployment rate be the set of stochastic processes $\{U_{i,t}\}$ given by:

$$\begin{aligned} U_{i,t} &= \alpha_{i,t} + \varepsilon_{i,t} \\ \alpha_{i,t} &= \sum_{k=1}^{mi} \gamma_{i,k} DU_{i,k,t} + \alpha_{i,t-1} + \upsilon_{i,t} \end{aligned} \tag{10}$$

where $\upsilon_{i,t} \to i.i.d.(0, \sigma_{\upsilon,i}^2)$ and $\alpha_{i,0} = \alpha_i$ represents a constant. $\{\varepsilon_{i,t}\}$ and $\{\upsilon_{i,t}\}$ are assumed mutually independent and, under the null hypothesis of regime-wise stationarity, $\sigma_{\upsilon,i}^2 = 0, \forall i = 1, \ldots, N$. The dummy variables for the changes in level are given by $DU_{i,k,t} = 1$ for $t > T_{b,k}^i$ and 0 otherwise, with $T_{b,k}^i$ denoting the $kth$ break location for the $ith$ individual, for $k = 1, \ldots, m_i, m_i \geq 1$. The specification in (10) is general enough to allow for unit-specific effects and unit-specific mean shifts.[14] CBL compute the panel stationarity test as the average of univariate KPSS tests:[15]

$$\eta(\hat{\lambda}) = N^{-1} \sum_{i=1}^{N} \left( \hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}_{i,t}^2 \right) \tag{11}$$

where $\eta_i(\hat{\lambda}_i) = \hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}_{i,t}^2$ is the univariate KPSS test for individual $i$, and $\hat{S}_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ stands for the partial sum of the estimated OLS residuals from (10).

[14] The test also allows for deterministic trends and shifts in slope, but due to the non-trending behaviour of unemployment—see Fig. 1 in the Appendix—we prefer not to include them in the regression. In addition, on economic grounds, it has more economic appeal to concentrate the analysis on explaining the infrequent changes in the mean of the unemployment rate series, as predicted by the structuralist approach. Notwithstanding, we computed the panel stationarity test allowing for mean and slope shifts, and the results which are available from the authors upon request remain unaffected.

[15] KPSS refers to the stationarity test proposed by Kwiatkowski et al. (1992).

$\hat{\omega}_i^2$ represents a consistent estimate of the long-run variance of $\varepsilon_{i,t}$, which allows for serial correlation and heteroskedasticity in the cross-sectional dimension. For robustness purposes, we allow for both heterogeneity and homogeneity in the estimation of the long-run variance.[16] Following Kurozumi (2002), we estimate the long-run variance non-parametrically using the Bartlett kernel with a bandwidth set according to the following expression:

$$\widehat{l} = \min\left\{1.1447\left(\frac{4\widehat{a}^2 T}{(1+\widehat{a})^2(1-\widehat{a})^2}\right)^{1/3}, 1.1447\left(\frac{4k^2 T}{(1+k)^2(1-k)^2}\right)^{1/3}\right\} \quad (12)$$

where $\widehat{a}$ is the autoregressive parameter estimated with the method proposed by Andrews (1991) and $k = 0.7$ is the preferred value according to Kurozumi's simulations. The test is dependent on the vector $\lambda_i = (\lambda_{i,1}, \ldots, \lambda_{i,mi})' = (T_{b,1}^i/T, \ldots, T_{b,mi}^i/T)'$ for each $i$, which indicates the location of the breaks relative to the whole period $(T)$. This vector is estimated employing the procedure of Bai and Perron (1998) which allows each individual unit to have a different number of breaks positioned heterogeneously across units. This procedure is based upon the global minimisation of the sum of squared residuals (SSR) and chooses as estimate of the breaks location the argument that minimises the sequence of unit-specific $SSR(T_{b,1}^i, \ldots, T_{b,mi}^i)$ obtained from (10):[17]

$$(\hat{T}_{b,1}^i, \ldots, \hat{T}_{b,mi}^i) = \arg\min_{(T_{b,1}^i, \ldots, T_{b,mi}^i)} SSR(T_{b,1}^i, \ldots, T_{b,mi}^i) \quad (13)$$

After estimating the dates for all possible $m_i \leq m^{\max}$ for each individual $i$, we use the sequential procedure proposed by Bai and Perron (1998) which employs pseudo F-statistics to compute and detect sequentially the optimal number of breaks, i.e. the optimal $m_i$. We then calculate the normalised test statistic of CBL which follows a standard normal distribution:[18]

$$Z(\hat{\lambda}) = \frac{\sqrt{N}\left(\eta(\hat{\lambda}) - \bar{\xi}\right)}{\bar{\varsigma}} \xrightarrow{d} N(0, 1) \quad (14)$$

where $\bar{\xi}$ and $\bar{\varsigma}^2$ are computed as averages of individual means and variances of $\eta_i(\lambda_i)$. The limiting distribution of the statistic is derived using sequential asymptotic theory in which $T \to \infty$ is followed by $N \to \infty$. The computation of the $Z(\hat{\lambda})$ statistic requires the individual series to be cross-sectionally independent along with asymptotic normality. Since these assumptions are unlikely to hold in practice, the finite sample distribution of the panel stationarity test with multiple breaks is approximated by the

16 Homogeneity can be assumed by replacing $\hat{\omega}_i^2$ in (11) by $\hat{\omega}^2 = N^{-1}\sum_{i=1}^{N} \hat{\omega}_i^2$.

17 A trimming value of 0.15 is used to eliminate endpoints.

18 The value of the $Z(\hat{\lambda})$ statistic must be compared with the critical values from an upper-tailed standard normal distribution.

**Table 1** Cross-sectional dependence test

| | EPA | INEM | Job-seekers |
|---|---|---|---|
| CD test | 11.305[a] | 17.266[a] | 16.977[a] |
| $p$ value | 0.000 | 0.000 | 0.000 |

The CD-statistic tests for the null of cross-sectional independence and is distributed as a two-tailed standard normal distribution

[a] Implies rejection of the null hypothesis at the 1% significance level

bootstrap distribution. For that purpose, we adopt the bootstrap procedure suggested by Maddala and Wu (1999) which allows for general forms of cross-dependencies across units.[19]

## 3 Empirical results and policy prescriptions

### 3.1 Results of the tests

Traditional panel unit root tests are derived under the assumption of cross-sectional independence of innovations, which is unlikely to hold in practice given the high degree of interdependencies across units. We begin the analysis by applying the CD statistic of Pesaran (2004) to innovations in unemployment rates for the three panels of the 17 Spanish regions. For each unit $i$ we compute OLS residuals from ADF regressions like (3), where the optimal lag-order is determined using the general-to-specific procedure suggested by Ng and Perron (1995) with a maximum lag-order of eight. As reported in Table 1, the null hypothesis that unemployment innovations are cross-sectionally independent is strongly rejected for all of the three variables employed to proxy for the excess of labour supply in the Spanish regions. This finding seems plausible and accords well with the fact that the Spanish regions are highly integrated in economic terms. This demonstrates that inferences deriving from the application of traditional panel unit root tests which are computed under the assumption of error cross-sectional independence are likely to be misleading as they are subject to dramatic size distortions.[20]

We now proceed to report the results from the application of a battery of recently developed panel unit root and stationarity tests which explicitly control for cross-sectional correlation. This includes the more powerful unit root tests developed by Smith et al. (2004), the test of Breitung and Das (2005) and the panel stationarity test of HLM that allows for at least one factor ($\hat{S}_k^F$). As reported in Table 2, we fail to

[19] In a nutshell, this method is based on resampling the estimated residuals from a fitted equation like (10) with a fixed cross-section index aimed at preserving the cross-sectional dependence structure present in the actual data. Then pseudo-observations are generated on the basis of the resampled pseudo-residuals. We then run the panel test on the pseudo-observations in order to derive the empirical distribution of the panel stationarity test with multiple breaks. In the process, we employ 20,000 replications to generate the bootstrap distribution.

[20] See O'Connell (1998), Maddala and Wu (1999), Strauss and Yigit (2003) and Banerjee et al. (2005).

**Table 2** Panel tests without breaks assuming cross-section dependence

| | EPA | | INEM | | Job-seekers | |
|---|---|---|---|---|---|---|
| | Test | $p$ value | Test | $p$ value | Test | $p$ value |
| $\Psi_{Max}$ | −0.892 | 0.618 | −0.892 | 0.618 | 0.098 | 1.000 |
| $\Psi_{Min}$ | 2.158 | 0.319 | 2.158 | 0.319 | 0.327 | 1.000 |
| $\Psi_{WS}$ | −1.092 | 0.551 | −1.092 | 0.551 | −0.280 | 0.999 |
| $t_{rob}$ | −0.653 | 0.257 | −1.094 | 0.137 | −0.477 | 0.317 |
| $\hat{S}_k^F$ | 3.737[a] | 0.000 | 4.473[a] | 0.000 | 4.282[a] | 0.000 |
| Rank (number of factors) | 2 | | 3 | | 4 | |

The $p$ values for the panel versions of the Max $t$ test, Min LM test and WS test are computed employing 20,000 bootstrap replications and defining a block size equal to 30. The maximum lag order is set at 8. The $\hat{S}_k^F$ statistic is a panel stationarity test and the others are panel unit root tests

[a] Implies rejection of the null hypothesis at 1%

reject the null of joint nonstationarity with Smith et al. (2004) and Breitung and Das (2005) tests, while we strongly reject the null of joint stationarity in unemployment with the $\hat{S}_k^F$ statistic. These results hold true irrespective of the proxy used to measure the excess of labour supply.

At this point, it is worth noting that all of the above tests fail to allow for structural breaks in the data generating process of unemployment. This could bias the results towards the non-rejection of the unit root null which may conceal the presence of stationarity with structural change. This may be particularly relevant in our case, as we find strong evidence of a unit root in Spanish regional unemployment when using tests that do not allow for structural breaks. To account for this fact, we extend the analysis by applying the recently developed panel stationarity test of CBL, which allows for multiple level shifts and is able to accommodate general forms of cross-sectional correlation through bootstrap methods. In Tables 3, 4 and 5 we present the results for the three different variables used to proxy for the excess of labour supply in the Spanish regions. In each of the tables, Panel A reports the individual KPSS tests allowing for a maximum of five structural breaks, which serve as a basis to compute the panel KPSS test with multiple breaks. Since comparing statistics obtained with relatively short data samples with asymptotic critical values may lead to a sizable finite-sample bias, we compute finite-sample critical values for the univariate KPSS tests with multiple breaks through Monte Carlo simulations employing 20,000 replications. Panel B reports the panel KPSS test with multiple breaks assuming cross-sectional independence and asymptotic normality and Panel C displays the bootstrap-based critical values of the test.

Beginning with the EPA unemployment rate, Panel A of Table 3 shows that all of the 17 Spanish regions have experienced at least two abrupt changes in their rate of unemployment over the period 1976–2004. More specifically, we find that nine regions (Aragón, Asturias, Baleares, Cantabria, Cataluña, Extremadura, Madrid, País Vasco and Comunidad Valenciana) present four breaks, seven regions (Andalucía, Canarias, Castilla-León, Castilla-La Mancha, Galicia, Murcia and Navarra) experience three

**Table 3** Panel KPSS stationarity test with multiple structural breaks. EPA unemployment rate

Panel A: Region-specific test

| Region | KPSS test | $m_i$ | Mean before the break | $\hat{T}^i_{b,1}$ | $\hat{T}^i_{b,2}$ | $\hat{T}^i_{b,3}$ | $\hat{T}^i_{b,4}$ | Finite sample critical values | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | | 10% | 5% | 1% |
| Andalucía | 0.097[b] | 3 | 15.397 | 27.565 | 32.283 | 22.643 | | 0.072 | 0.085 | 0.115 |
| | | | | *83Q01* | *92Q03* | *99Q01* | | | | |
| Aragón | 0.076[b] | 4 | 5.094 | 14.143 | 11.120 | 16.188 | 9.150 | 0.058 | 0.067 | 0.087 |
| | | | | *80Q04* | *88Q03* | *92Q04* | *97Q04* | | | |
| Asturias | 0.069[b] | 4 | 5.402 | 13.016 | 18.052 | 20.872 | 18.631 | 0.058 | 0.068 | 0.089 |
| | | | | *80Q03* | *84Q04* | *92Q04* | *98Q01* | | | |
| Baleares | 0.072[b] | 4 | 4.791 | 12.802 | 10.612 | 15.560 | 9.398 | 0.06 | 0.071 | 0.094 |
| | | | | *80Q04* | *87Q04* | *92Q02* | *96Q03* | | | |
| Canarias | 0.070 | 3 | 9.739 | 18.003 | 23.852 | 14.003 | | 0.098 | 0.124 | 0.183 |
| | | | | *80Q03* | *84Q04* | *98Q02* | | | | |
| Cantabria | 0.070[b] | 4 | 5.019 | 12.954 | 17.659 | 22.419 | 14.661 | 0.058 | 0.069 | 0.092 |
| | | | | *81Q01* | *85Q02* | *93Q03* | *98Q01* | | | |
| Cataluña | 0.098[a] | 4 | 6.352 | 19.629 | 13.336 | 19.167 | 10.185 | 0.059 | 0.068 | 0.091 |
| | | | | *80Q03* | *88Q04* | *93Q01* | *98Q01* | | | |
| Castilla-León | 0.072 | 3 | 6.331 | 14.504 | 19.091 | 13.159 | | 0.087 | 0.107 | 0.157 |
| | | | | *80Q03* | *92Q04* | *98Q04* | | | | |
| Castilla-La Mancha | 0.105[b] | 3 | 6.212 | 15.734 | 20.008 | 13.650 | | 0.078 | 0.094 | 0.131 |
| | | | | *81Q03* | *92Q02* | *98Q04* | | | | |
| Extremadura | 0.095[a] | 4 | 13.254 | 26.245 | 25.397 | 30.003 | 26.089 | 0.055 | 0.064 | 0.083 |
| | | | | *83Q04* | *88Q03* | *92Q04* | *98Q04* | | | |
| Galicia | 0.115[b] | 3 | 3.879 | 12.028 | 18.232 | 15.579 | | 0.072 | 0.085 | 0.117 |
| | | | | *82Q03* | *92Q01* | *98Q04* | | | | |
| Madrid | 0.080[b] | 4 | 8.088 | 17.091 | 12.664 | 18.967 | 10.769 | 0.058 | 0.067 | 0.088 |
| | | | | *80Q03* | *88Q03* | *92Q04* | *98Q04* | | | |
| Murcia | 0.104[b] | 3 | 7.010 | 17.258 | 23.373 | 13.601 | | 0.081 | 0.098 | 0.141 |
| | | | | *81Q01* | *92Q02* | *97Q04* | | | | |
| Navarra | 0.083[c] | 3 | 6.206 | 15.704 | 12.196 | 7.136 | | 0.078 | 0.094 | 0.133 |
| | | | | *80Q03* | *88Q02* | *98Q04* | | | | |
| País Vasco | 0.065[c] | 4 | 7.003 | 17.913 | 22.378 | 20.488 | 11.682 | 0.064 | 0.077 | 0.108 |
| | | | | *80Q03* | *84Q04* | *89Q01* | *98Q03* | | | |
| La Rioja | 0.049 | 2 | 3.601 | 12.972 | 9.040 | | | 0.14 | 0.178 | 0.273 |
| | | | | *81Q03* | *98Q03* | | | | | |
| Comunidad Valenciana | 0.101[a] | 4 | 5.862 | 18.021 | 15.696 | 22.108 | 12.525 | 0.055 | 0.064 | 0.082 |
| | | | | *81Q01* | *88Q01* | *92Q02* | *98Q01* | | | |

**Table 3** continued

| Panel B: Panel KPSS test with multiple breaks assuming cross-section independence | | | | |
|---|---|---|---|---|
| | Test | $p$ value | | |
| $Z(\hat{\lambda})$ (homogeneous) | 6.708[a] | 0.000 | | |
| $Z(\hat{\lambda})$ (heterogeneous) | 6.990[a] | 0.000 | | |
| Panel C: Bootstrap distribution | | | | |
| | 90% | 95% | 97.5% | 99% |
| $Z(\hat{\lambda})$ (homogeneous) | 2.182 | 2.512 | 2.833 | 3.246 |
| $Z(\hat{\lambda})$ (heterogeneous) | 1.706 | 2.134 | 2.566 | 3.079 |

The $Z(\hat{\lambda})$ (homogeneous) and $Z(\hat{\lambda})$ (heterogeneous) denote the panel KPSS test with multiple breaks developed in CBL for the case of homogeneity and heterogeneity in the estimation of the long-run variance, respectively

[a,b,c] Imply rejection of the null hypothesis at 1, 5 and 10%, respectively. The entries in columns 4–8 are the mean unemployment rates in each regime, with the break dates reported in italics

breaks, and only La Rioja does have two breaks—see Fig. 1 in the Appendix.[21] In addition, we can observe some clustering of the timing of the breaks and a clear pattern in the fluctuations of the mean unemployment rate across different regimes which roughly coincide with the phases of the business cycles experienced over the past 30 years.[22] Among the 59 breaks identified, 21 breaks occurred in the first half of the eighties (period 1980–1985) of which 14 were concentrated in the years 1980–1981; eight breaks occurred during the period 1987–1989; 13 breaks occurred between 1992 and 1993 of which 11 were concentrated in 1992; and 17 breaks took place during the period 1996–1999 of which 13 occurred in 1998.

It is also interesting to observe how the mean rate of unemployment has shifted in response to major macroeconomic shocks occurring over the past three decades. Columns 4 to 8 of Panel A of Table 3 report the mean unemployment rate for each of the regimes identified, which allows us to determine the sign (and magnitude) of the mean shifts in unemployment. Among the breaks identified during the period 1980–1985, all of them are positive, which coincide with the wide-spread rise in unemployment rates during the cyclical downturn that followed the second oil shock. In addition, the contractive aggregate demand policies implemented around that time with the aim of containing the excessive elevation of prices, which began with the *Pacto de la Moncloa*

[21] As noted by an anonymous referee, the procedures employed to identify significant structural breaks may lead to too many breaks. As shown in Fig. 1, some regular upward movements in unemployment rates appear to be captured by an upward shift in the trend function of the variable. Notwithstanding, the presence of multiple mean shifts in the series appears to be a robust feature of the data under analysis.

[22] Despite the fact that Spanish regional unemployment somewhat displays a cyclical pattern of unemployment which is more typical of the US experience, each upward shift in unemployment in response to adverse shocks has tended to become permanent as suggested by the hysteresis hypothesis.

**Table 4** Panel KPSS stationarity test with multiple structural breaks. INEM unemployment rate

Panel A: Region-specific test

| Region | KPSS test | $m_i$ | Mean before the break | $\hat{T}^i_{b,1}$ | $\hat{T}^i_{b,2}$ | $\hat{T}^i_{b,3}$ | $\hat{T}^i_{b,4}$ | $\hat{T}^i_{b,5}$ | Finite sample critical values | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | | | 10% | 5% | 1% |
| Andalucía | 0.131[a] | 4 | 11.190 | 20.233 | 26.754 | 21.153 | 12.488 | | 0.059 | 0.069 | 0.094 |
| | | | | *80M09* | *84M012* | *89M09* | *96M07* | | | | |
| Aragón | *0.083*[a] | 5 | 4.120 | 10.909 | 14.643 | 10.506 | 11.124 | 6.945 | 0.044 | 0.05 | 0.066 |
| | | | | *80M09* | *84M12* | *89M06* | *93M09* | *98M03* | | | |
| Asturias | 0.090[a] | 5 | 4.708 | 13.794 | 18.374 | 16.753 | 17.494 | 12.230 | 0.043 | 0.049 | 0.062 |
| | | | | *80M09* | *84M12* | *89M08* | *93M11* | *98M12* | | | |
| Baleares | 0.065[c] | 4 | 7.042 | 14.762 | 14.046 | 7.910 | 6.287 | | 0.057 | 0.066 | 0.087 |
| | | | | *80M09* | *88M05* | *95M09* | *99M12* | | | | |
| Canarias | 0.080[b] | 4 | 6.483 | 16.925 | 21.105 | 17.505 | 11.679 | | 0.062 | 0.075 | 0.106 |
| | | | | *80M09* | *84M12* | *94M08* | *98M11* | | | | |
| Cantabria | 0.093[b] | 4 | 4.410 | 12.871 | 17.061 | 14.427 | 9.241 | | 0.058 | 0.068 | 0.093 |
| | | | | *81M02* | *85M05* | *89M08* | *97M12* | | | | |
| Cataluña | 0.090[b] | 3 | 7.056 | 17.540 | 11.148 | 6.423 | | | 0.070 | 0.082 | 0.108 |
| | | | | *81M09* | *88M12* | *97M07* | | | | | |
| Castilla-León | 0.229[a] | 3 | 4.261 | 10.508 | 15.005 | 13.11 | 9.41 | | 0.092 | 0.115 | 0.171 |
| | | | | *80M09* | *84M12* | *89M04* | *98M03* | | | | |
| Castilla-La Mancha | 0.075[b] | 4 | 5.443 | 12.613 | 18.130 | 14.875 | 10.333 | | 0.060 | 0.072 | 0.099 |
| | | | | *80M09* | *84M12* | *89M04* | *98M03* | | | | |
| Extremadura | 0.099[a] | 5 | 9.258 | 17.639 | 22.645 | 19.518 | 14.001 | 11.812 | 0.042 | 0.047 | 0.059 |
| | | | | *81M07* | *86M06* | *91M03* | *96M03* | *2000M06* | | | |
| Galicia | 0.089[b] | 4 | 3.846 | 11.223 | 15.592 | 15.269 | 11.251 | | 0.057 | 0.068 | 0.091 |
| | | | | *81M09* | *85M12* | *94M06* | *98M09* | | | | |
| Madrid | 0.102[a] | 4 | 5.762 | 14.386 | 10.752 | 13.137 | 7.695 | | 0.055 | 0.064 | 0.084 |
| | | | | *81M07* | *89M01* | *93M04* | *98M06* | | | | |
| Murcia | 0.106[b] | 3 | 5.996 | 15.051 | 12.371 | 7.297 | | | 0.070 | 0.083 | 0.111 |
| | | | | *81M12* | *89M04* | *98M03* | | | | | |
| Navarra | 0.126[a] | 4 | 5.702 | 17.426 | 12.940 | 10.368 | 6.907 | | 0.054 | 0.062 | 0.081 |
| | | | | *82M04* | *89M07* | *94M10* | *99M01* | | | | |
| País Vasco | 0.109[a] | 4 | 6.056 | 13.868 | 19.870 | 14.383 | 8.168 | | 0.058 | 0.069 | 0.093 |
| | | | | *80M09* | *84M12* | *89M10* | *98M01* | | | | |
| La Rioja | 0.072[c] | 4 | 1.952 | 8.510 | 12.738 | 10.186 | 6.483 | | 0.061 | 0.074 | 0.103 |
| | | | | *80M09* | *84M12* | *89M03* | *98M06* | | | | |
| Comunidad | 0.116[a] | 4 | 5.378 | 15.201 | 20.810 | 16.441 | 7.806 | | 0.060 | 0.070 | 0.096 |
| Valenciana | | | | *80M09* | *84M12* | *89M03* | *97M12* | | | | |

**Table 4** continued

Panel B: Panel KPSS test with multiple breaks assuming cross-section independence

| | Test | $p$ value |
|---|---|---|
| $Z(\hat{\lambda})$ (homogeneous) | 16.602[a] | 0.000 |
| $Z(\hat{\lambda})$ (heterogeneous) | 15.492[a] | 0.000 |

Panel C: Bootstrap distribution (%)

| | 90% | 95% | 97.5% | 99% |
|---|---|---|---|---|
| $Z(\hat{\lambda})$ (homogeneous) | 8.756 | 9.870 | 10.877 | 12.064 |
| $Z(\hat{\lambda})$ (heterogeneous) | 7.040 | 8.044 | 8.990 | 10.223 |

See Table 3

(1977), contributed to the unemployment rise.[23] In contrast to the first cluster of breaks, the period 1987–1989 is characterised by steadily falling unemployment rates as confirmed by all of the breaks showing a negative sign. This prevailing tendency for unemployment rates to fall during the second half of the eighties can be attributed to the cyclical upturn setting in during these years. The third cluster of breaks is associated with the recessionary period of the early nineties (1992–1993) that experienced a sharp rise in unemployment rates caused by monetary policy tightening. This led to a sharp elevation of interest rates as well as to the subsequent crisis of the European Monetary System. In addition, the commitment of Spain to fulfil the Maastricht criteria aimed at achieving nominal convergence towards EU average levels, led to contractive fiscal policies which further prevented unemployment rates from falling.[24] This is confirmed by all of the mean shifts being positive. The fourth cluster is associated with the cyclical upturn of the second half of the 1990s, which coincides with the marked decline in interest rates and the apparent effects of the reforms implemented with the aim of increasing labour market flexibility (1992, 1993, 1994 and 1997). As shown in Fig. 1, all of the regions appear to experience a significant fall in unemployment during these years.

Interestingly, we observe some evidence of business cycle asymmetries in regard to the adjustment of unemployment rates in response to shocks: there appears to be sharp increases in unemployment during cyclical downturns (mainly those of the early 1980s and early 1990s) and smaller declines during cyclical upturns. This is confirmed by the fact that the mean unemployment rate estimated for the final regime in the late 1990s or early 2000s is usually well above that estimated for the first unemployment

[23] In addition, from the late seventies to the early nineties there was a huge expansion in the level and range of welfare entitlements such as unemployment benefits which could have contributed to the fall in the willingness to work.

[24] Blanchard and Wolfers (2000) also note that the occurrence of adverse shifts in labour demand as indicated by the decrease in the labour share since the early eighties—which may be caused by technological bias away from labour or by a decrease in the wage relative to marginal labour productivity—may have contributed to the high unemployment levels recorded in the 1980s and 1990s.

**Table 5** Panel KPSS stationarity test with multiple structural breaks. Job-seekers rate

Panel A: Region-specific test

| Region | KPSS test | $m_i$ | Mean before the break | $\hat{T}^i_{b,1}$ | $\hat{T}^i_{b,2}$ | $\hat{T}^i_{b,3}$ | $\hat{T}^i_{b,4}$ | Finite sample critical values | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | | 10% | 5% | 1% |
| Andalucía | 0.094[b] | 4 | 13.187 | 30.808 | 44.641 | 34.620 | 27.423 | 0.066 | 0.081 | 0.118 |
| | | | | *81M10* | *85M12* | *96M03* | *2000M05* | | | |
| Aragón | 0.093[a] | 4 | 4.973 | 12.259 | 17.398 | 19.525 | 12.630 | 0.055 | 0.065 | 0.085 |
| | | | | *81M02* | *85M04* | *91M05* | *97M11* | | | |
| Asturias | 0.094[b] | 4 | 5.846 | 15.928 | 21.967 | 28.030 | 19.475 | 0.063 | 0.076 | 0.107 |
| | | | | *81M02* | *85M04* | *89M06* | *98M12* | | | |
| Baleares | 0.064[c] | 4 | 8.319 | 14.717 | 16.722 | 22.166 | 13.982 | 0.057 | 0.068 | 0.092 |
| | | | | *81M02* | *85M09* | *90M03* | *96M04* | | | |
| Canarias | 0.113[a] | 4 | 7.515 | 17.568 | 24.691 | 26.212 | 21.830 | 0.057 | 0.067 | 0.091 |
| | | | | *81M03* | *85M05* | *90M02* | *97M06* | | | |
| Cantabria | 0.135[a] | 4 | 6.524 | 15.852 | 23.232 | 20.233 | 13.842 | 0.057 | 0.068 | 0.093 |
| | | | | *81M06* | *86M09* | *95M02* | *99M04* | | | |
| Cataluña | 0.067[c] | 4 | 7.092 | 16.711 | 21.010 | 18.013 | 10.931 | 0.060 | 0.071 | 0.097 |
| | | | | *81M02* | *85M04* | *89M06* | *97M09* | | | |
| Castilla-León | 0.097[a] | 4 | 5.226 | 11.663 | 19.449 | 22.321 | 15.812 | 0.057 | 0.067 | 0.091 |
| | | | | *81M02* | *85M04* | *90M06* | *98M05* | | | |
| Castilla-La Mancha | 0.100[a] | 4 | 6.382 | 13.829 | 20.802 | 23.947 | 16.704 | 0.055 | 0.064 | 0.083 |
| | | | | *81M02* | *85M04* | *92M01* | *98M03* | | | |
| Extremadura | 0.114[b] | 4 | 10.155 | 26.053 | 39.154 | 33.226 | 28.050 | 0.068 | 0.083 | 0.124 |
| | | | | *81M10* | *85M12* | *96M07* | *2000M09* | | | |
| Galicia | 0.176[a] | 3 | 4.837 | 13.962 | 22.568 | 17.785 | | 0.079 | 0.097 | 0.143 |
| | | | | *82M04* | *87M07* | *98M09* | | | | |
| Madrid | 0.089[b] | 4 | 5.839 | 13.488 | 20.072 | 21.485 | 13.950 | 0.057 | 0.066 | 0.089 |
| | | | | *81M02* | *85M04* | *93M03* | *98M06* | | | |
| Murcia | 0.113[a] | 4 | 6.281 | 15.212 | 20.292 | 23.554 | 13.255 | 0.055 | 0.064 | 0.085 |
| | | | | *81M08* | *85M10* | *91M03* | *97M06* | | | |
| Navarra | 0.101[a] | 4 | 6.386 | 18.699 | 23.029 | 18.249 | 13.480 | 0.057 | 0.067 | 0.092 |
| | | | | *82M04* | *86M06* | *94M10* | *98M12* | | | |
| País Vasco | 0.098[b] | 4 | 8.246 | 16.178 | 26.044 | 21.527 | 13.235 | 0.062 | 0.075 | 0.105 |
| | | | | *81M02* | *85M05* | *94M11* | *99M01* | | | |
| La Rioja | 0.155[b] | 3 | 2.481 | 9.804 | 17.132 | 11.697 | | 0.112 | 0.143 | 0.221 |
| | | | | *81M02* | *85M04* | *98M06* | | | | |
| Comunidad Valenciana | 0.114[a] | 4 | 6.542 | 17.238 | 24.325 | 28.519 | 15.019 | 0.056 | 0.065 | 0.084 |
| | | | | *81M02* | *85M04* | *90M12* | *97M12* | | | |

**Table 5** continued

Panel B: Panel KPSS test with multiple breaks assuming cross-section independence

| | Test | $p$ value |
|---|---|---|
| $Z(\hat{\lambda})$ (homogeneous) | 14.725[a] | 0.000 |
| $Z(\hat{\lambda})$ (heterogeneous) | 14.667[a] | 0.000 |

Panel C: Bootstrap distribution (%)

| | 90% | 95% | 97.5% | 99% |
|---|---|---|---|---|
| $Z(\hat{\lambda})$ (homogeneous) | 4.342 | 5.134 | 5.889 | 6.790 |
| $Z(\hat{\lambda})$ (heterogeneous) | 3.317 | 4.131 | 4.921 | 5.837 |

See Table 3

regime (before the occurrence of the first mean shift) which is generally associated with the second half of the 1970s.

Turning to the analysis of the non-stationarity properties of the EPA unemployment rate, the univariate KPSS tests with multiple breaks reject the null of regime-wise stationarity for all the regions at conventional levels with the exception of Canarias, Castilla-León and La Rioja.[25] With the panel KPSS test with multiple breaks which is shown in Panel B of Table 3, we are able to reject the null of joint regime-wise stationarity in favour of hysteresis at the 1% level. In Panel C of Table 3 we report the bootstrap distribution of the panel KPSS test with multiple breaks so that we can check that the rejection of the null is not caused by the size distortions experienced by panel tests that fail to control for cross-sectional dependence. Despite the fact the bootstrap critical values are considerably higher than those associated with an upper-tailed standard normal distribution, the panel KPSS test is still able to reject the null at the 1% level. Therefore, there appears to be strong evidence supportive of a unit root in the unemployment rate of the Spanish regions, which is consistent with the hysteresis hypothesis.

As far as the INEM unemployment rate is concerned, Table 4 shows that there is also strong evidence of structural change in unemployment since all of the regions experience at least three changes in their unemployment mean. As with the EPA unemployment rate, the INEM unemployment rate shows somewhat similar clustering of breaks and a number of unemployment regimes which roughly coincide with the phases of the business cycles.[26] Furthermore, we are able to reject the null of regime-wise stationarity at conventional significance levels for all the regions. Also, the panel

[25] The null of regime-wise stationarity is rejected for Cataluña, Extremadura and Comunidad Valenciana at the 1% level, for Andalucía, Aragón, Asturias, Baleares, Cantabria, Castilla-La Mancha, Galicia, Madrid and Murcia at the 5% level, and for Navarra and País Vasco at the 10% level.

[26] Due to space limitations, we only show the figure that corresponds to the EPA unemployment rate. For similar reasons, we do not stop to discuss the clustering of breaks and the dynamic behaviour of the INEM unemployment and job seekers rates, which follow similar patterns to the EPA unemployment rate. We refer the readers to the information contained in Tables 4 and 5.

KPSS test with multiple breaks appears to reject the null of regime-wise stationarity at the 1% level irrespective of the assumption of cross-sectional dependence.[27]

The evidence favouring hysteresis in Spanish regional unemployment is also strong for the job-seekers rate. As shown in Table 5, we again reject the null of regime-wise stationarity with the univariate KPSS test at conventional significance levels for all the regions. Along the same lines, the panel KPSS test with multiple breaks is also able to reject the null of regime-wise stationarity at the 1% level irrespective of the assumption of cross-sectional dependence.

Taken as a whole, our thorough investigation of the time-series properties of Spanish regional unemployment provides strong and robust evidence supportive of the hysteresis hypothesis. In our opinion, this evidence should be taken very seriously as we try to properly control for cross-sectional correlation, multiple mean shifts and finite-sample bias. Furthermore, by taking stationarity as the null, we can be more confident of the presence of a unit root in unemployment, as the null tends to be rejected when there is strong evidence against it.

### 3.2 Policy prescriptions

Overall, the high degree of persistence leading to hysteretic behaviour in Spanish regional unemployment may be the result of a number of institutional arrangements governing the functioning of the economy that have caused sluggishness in the adjustment process in response to macroeconomic shocks.[28] There are several contributions in this line for the Spanish economy. Blanchard and Wolfers (2000) emphasise the explanation of unemployment in the OECD countries by means of a combination of shocks and institutions. Bentolila and Jimeno (2003) explain the behaviour of Spanish unemployment in recent decades using this approach, concluding that Spain experienced similar shocks to other OECD countries, except for stronger labour demand shifts. Therefore, their explanation relies mostly on Spain having a set of exceptionally unemployment-prone labour institutions, in particular employment protection, unemployment benefits and collective bargaining.

Using a structural vector autoregression approach, Dolado and Jimeno (1997) point out that the combination of a plausible mixture of different types of shocks and extreme persistence in their propagation mechanisms, mainly due to rigidities in both good and labour markets, can satisfactorily explain the dismal performance of the Spanish labour market. Furthermore, Jimeno and Bentolila (1998) find a high degree of persistence in Spanish regional unemployment due to low wage flexibility. This appears to be mainly caused by the low responsiveness of migration and labour participation to

[27] Note that the panel KPSS test takes a much greater value than the critical values from an upper-tailed normal distribution and than the bootstrap critical values which correct for cross-sectional dependence as well as for finite-sample bias.

[28] Nickell (1997) considers a labour index which takes into consideration the following five dimensions: working time, fixed-term contracts, employment protection, minimum wages and employees' representation rights. Spain and Italy appear to be the countries with the most restrictive rules and regulations and in turn the least flexible labour markets.

regional wages, which is reinforced by the low elasticity of these two channels to regional unemployment.

To sum up, from an economic policy perspective, our results call for policy measures aimed at improving labour market flexibility conditions which speed up the adjustment process in response to adverse shocks, thereby preventing upward shifts in unemployment from becoming permanent. Along these lines, we have to conclude that the main labour market reforms endorsed in 1984, 1994 and 1997 were not optimally designed and implemented (Segura 2001). As a result, some of the aforementioned sources of hysteresis still remain and should be corrected in future reforms (Usabiaga 2006).

## 4 Conclusions

Unit root testing has now become common practice in applied macroeconomics. The fact that conventional unit root and stationarity tests normally lack power has led to the development of panel unit root and stationarity tests which are more powerful than their univariate counterparts. Since the seminal work of Perron (1989), it has become widely recognised that failure to control for structural breaks favours the nonrejection of the null of nonstationarity. More recently, it has also become well known the fact that panel tests that fail to allow for cross-sectional correlation are subject to severe size distortions. In this paper, we have investigated the hysteresis hypothesis for the Spanish regional unemployment over the past three decades taking into account the aforementioned considerations.

We have first applied the CD statistic of Pesaran (2004) to unemployment innovations, rendering strong evidence of cross-sectional dependence in the error structure of our panel of regions. Second, we have accounted for this fact by employing a battery of panel unit root and stationarity tests that explicitly allow for cross-sectional dependence across regions. Our findings are strongly supportive of hysteresis in Spanish regional unemployment. These results appear reinforced with the evidence obtained from the panel KPSS stationarity test with multiple breaks, which controls for general forms of cross-sectional dependence and for finite-sample bias through bootstrap methods. In addition, the fact that we use tests that take as null both nonstationarity and regime-wise stationarity implies that our analysis can be thought of as confirmatory, thus making us more confident of the presence of hysteresis in Spanish regional unemployment over the past 30 years.

An important policy implication of our results is that stabilisation policy may have permanent (or at least long-lasting) effects on Spanish unemployment. In this context, the level of aggregate demand and the corresponding policies deserve great attention. The restrictive demand policies aiming at the achievement of the Maastricht criteria and the disinflation objectives of central banks may have imposed a very costly burden on the Spanish economy. This has contributed to the prevailing tendency of Spanish regional unemployment rates to rise from the already high level reached after the first oil shock. In addition, this high degree of persistence in unemployment further gives an indication that labour market reforms implemented in the Spanish economy in recent decades were not optimally designed to combat the underlying sources of hysteresis.

## Appendix

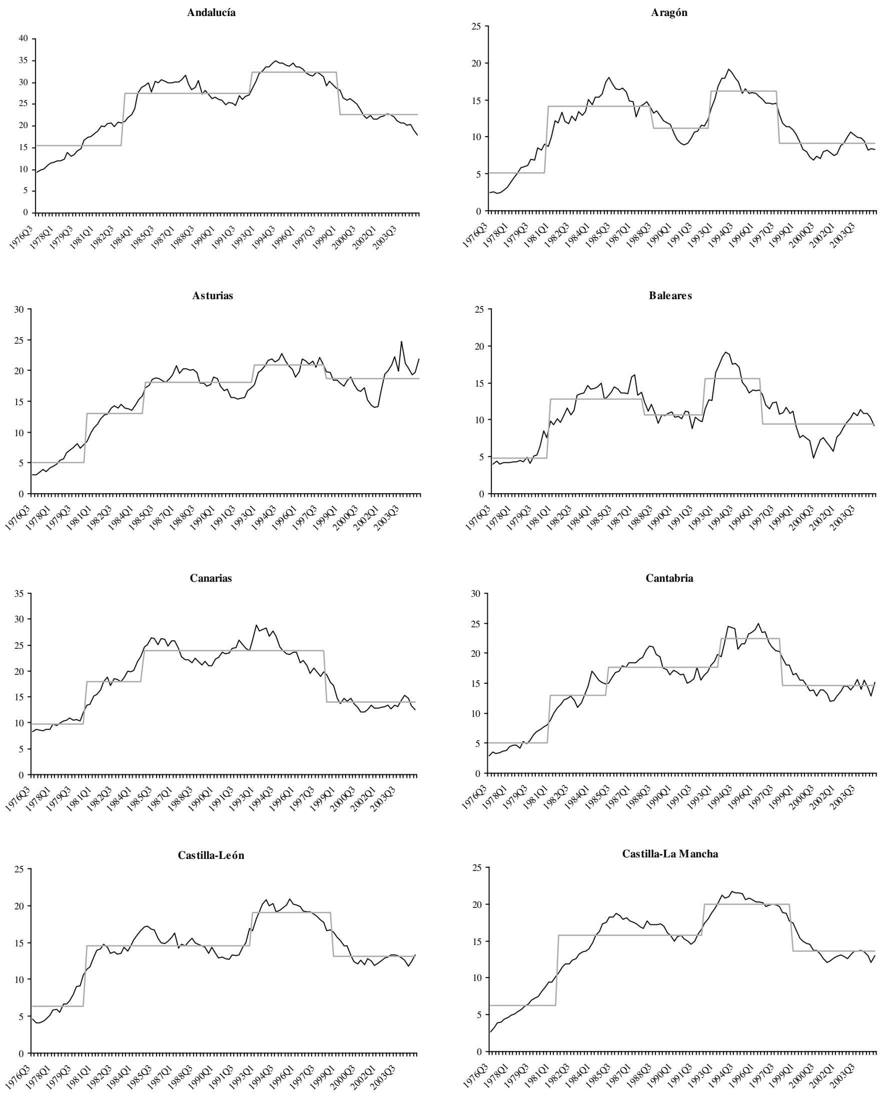


**Fig. 1** Actual and mean unemployment rate (EPA). Spanish regions. 1976Q1-2004Q4

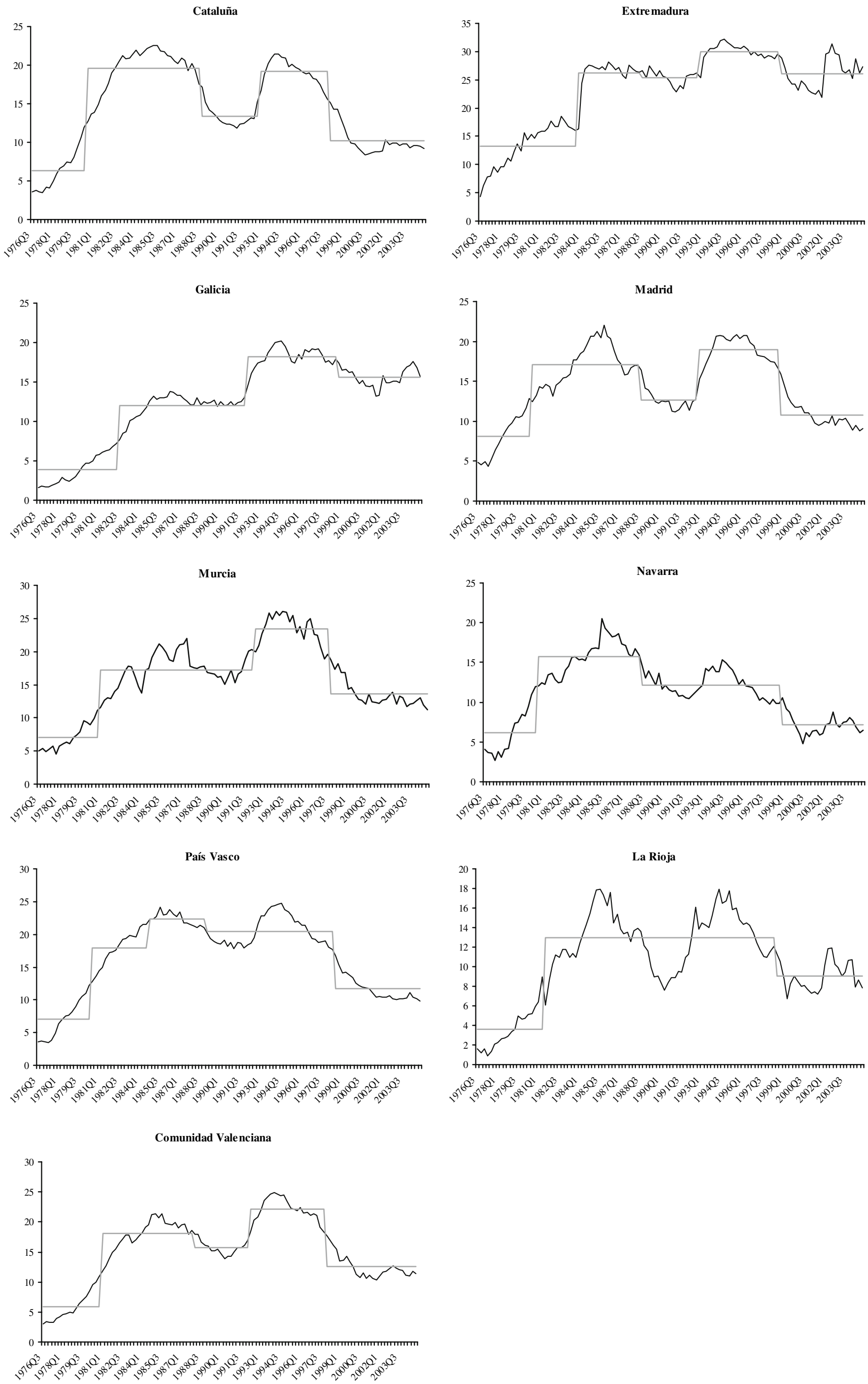


**Fig. 1** continued

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