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PANIC in Spanish Unemployment

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Abstract: This paper investigates the stochastic properties of Spanish regional unemployment rates using quarterly data spanning the period 1976(1)-2013(2). Towards this end, we employ the PANIC procedures of Bai and Ng (2004) and Bai and Ng (2010), which allow for the decomposition of the observed unemployment rate series into a common component and an idiosyncratic component. This enables us to identify the exact source behind the nonstationary behaviour in Spanish regional unemployment. Our analysis provides clear-cut evidence for the hysteresis hypothesis, which appears to be caused by a common stochastic trend driving all the regional unemployment series. These results are largely backed up by those obtained from the median-unbiased estimation of the persistence parameter which equals one for the common factor and to a lesser extent from the half-life estimate obtained from impulse-response functions that equals 12.5 years for the case of shocks hitting the common trend. This persistence analysis also provides evidence of stationarity for most of the idiosyncratic series (two regions being the exception), as found in the PANIC analysis.

Keywords: Unemployment rate, Persistence, PANIC, Common Factor, Median-Unbiased Estimation, Half-Lives.

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1. Introduction

This paper investigates the stochastic properties of the unemployment rate of the Spanish regions over the period 1976-2013. This is done by using the Panel Analysis of Nonstationarity in Idiosyncratic and Common components (PANIC) framework developed by Bai and Ng (2004) and further extended by Bai and Ng (2010). Unlike most panel tests employed in the literature, the PANIC approach allows for strong forms of cross-sectional correlation such as cross-cointegration. Previous studies have emphasised the fact that failure to allow for cross-dependence when it is present in the data can cause severe size distortions (see O'Connell, 1998; Maddala and Wu, 1999; Banerjee et al., 2005). In addition, PANIC methods will enable us to (1) investigate whether there is pervasive cross-sectional dependence driven by one or several common factors and (2) determine whether the source of nonstationarity is the common factors and/or the idiosyncratic components.

The high and persistent level of Spanish unemployment which has sustained over the past three and a half decades is considered to be one of the most striking cases of hysteresis and has drawn much attention among scholars in the field. The explanations about the poor performance of the Spanish unemployment have relied on the notion of "Eurosclerosis" (unemployment-prone institutions), the adverse interaction between shocks and institutions, the long-term unemployment predicament, the real wage rigidity, the limited geographical mobility, and mainly the insider-outsider model (dual labour market).[1]

Determining which component(s) of the observed unemployment rate series are driving the high persistence in the series is therefore necessary if one wants to formulate economic policy prescriptions in order to tackle the unemployment problem over the medium term.

Overall, the empirical analysis renders overwhelming evidence of nonstationarity in the regional unemployment rate series, which is found to be driven by a common stochastic factor. In the case of stationary idiosyncratic series, the evidence supports the existence of pairwise cointegration across regional unemployment rate series. In practical terms, the results indicate that the hysteretic behaviour in Spanish unemployment observed across all regions is caused by a common force that appears to exhibit nonstationarity.

As a complement to the PANIC analysis we also provide estimates of the persistence parameter and the half-life of a shock to both the common factor and the idiosyncratic series. This is because unit root tests may provide incomplete information regarding the degree of persistence of a series due to their exclusive focus on the too restrictive distinction between the I(0) and I(1) hypotheses, whereas the half-lives analysis can provide an accurate and exact measure of persistence by

[1] Some representative contributions in this area are Blanchard and Summers (1986, 1987), Lindbeck and Snower (1988), Layard et al. (1991), Blanchard and Jimeno (1995), Nickell (1997), Jimeno and Bentolila (1998), Blanchard and Wolfers (2000), Nickell et al. (2005), Bentolila and Jimeno (2006), Ball (2009) and Bentolila et al. (2012a). This list combines Spain-focused studies with more international essays.

capturing the number of years for a unit impulse (hitting either the common factor or the idiosyncratic components) to dissipate by one half.[2] Towards this end, we deploy the median-unbiased approach of Andrews and Chen (1994) that computes half-lives directly from impulse-response functions. To the best of our knowledge, this is the first attempt in the literature at establishing the degree of persistence of both the common factor and idiosyncratic components obtained from the decomposition of observed unemployment rate series following the PANIC methodology.

The analysis of persistence calculating median-unbiased estimates of the persistence parameter and half-lives for both components provides evidence lending support to those obtained with the PANIC procedures. First, region-specific idiosyncratic components appear to revert to their mean, as given by finite half-lives in all the idiosyncratic series with the exception of two regions. This implies that the effects from idiosyncratic shocks hitting region-specific unemployment rates will vanish as time passes. Second, there is clear-cut evidence of a median-unbiased point estimate of the persistence parameter for the common factor that takes a value of unity, which is consistent with the unit root hypothesis. Besides, a half-life point estimate of 12.5 years for the case of shocks hitting the common factor implies that it takes a relatively long period (for a variable like unemployment) for a unit impulse on the common component of regional unemployment rates to dissipate by half.

The remainder of the paper is structured as follows. Section 2 briefly reviews the main theoretical hypotheses behind the behaviour of the unemployment rate and presents the main previous studies on the issue. Section 3 describes the PANIC methodology employed in the paper. Section 4 presents the results from the decomposition of the observed unemployment rate series of the Spanish regions into a common and an idiosyncratic component and their respective time series properties obtained via PANIC. Section 5 is devoted to providing median-unbiased estimates of the persistence parameter and half-lives for both the common and idiosyncratic components. Section 6 concludes.

2. Literature Review

There are three main theoretical hypotheses regarding the behaviour exhibited by the unemployment rate. The traditional natural rate theory pioneered by the work of Friedman (1968) and Phelps (1967, 1968) advocates that the unemployment rate fluctuates around some natural or equilibrium rate implying a fully equilibrated labour market. Shocks to unemployment are thus temporary and the unemployment rate is assumed to revert to its equilibrium level.

The labour market developments of the 1970s and 1980s gave rise to the structuralist hypothesis, as a refinement of the natural rate theory. Proponents of this hypothesis such as Phelps (1994) and

[2] Since the idiosyncratic series reflect deviations of the observed unemployment rate series from the common trend, the measurement of the half-life of a shock to the idiosyncratic series is of particular economic relevance, since it can be informative of how fast individual regions adjust to economic shocks.

Phelps and Zoega (1998) support that the natural rate of unemployment has tended to shift upward as a result of changes in structural factors of the economy that include the slowdown in productivity growth (Hoon and Phelps, 1997), the steep rises in oil prices in the seventies (Carruth et al., 1998) and the changes in world interest rates (Phelps, 1994). Under this hypothesis, most shocks to unemployment appear to be temporary, but occasionally the natural rate permanently changes as a result of infrequent and large shocks. However, since the US unemployment rate reverted back to its pre-shock level rather than steadily rising as in Europe, such prominent authors as Blanchard and Summers (1986, 1987) turned their attention to the hysteresis hypothesis which advocates that cross-country differences in institutional arrangements governing the functioning of labour markets lead to marked differences in the way economies adjust to macroeconomic shocks. In some European labour markets the prevalence of labour market rigidities may be responsible for the sluggishness in the adjustments to adverse shocks. As opposed to the natural rate hypothesis, the hysteresis hypothesis implies that the unemployment rate is path-dependent so that current unemployment levels heavily depend on past levels.

These theoretical hypotheses have been tested in the empirical literature assessing the time series properties of the unemployment rate. In short, the hysteresis hypothesis is formulated as a unit root process, and its rejection gives support to the natural rate theory if no breaks are 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. First, there is a group of studies using traditional univariate 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. The evidence generally supports the hypothesis of hysteresis for the EU economies and appears mixed for the US.

Second, such studies as Mitchell (1993), Bianchi and Zoega (1998), Arestis and Biefang-Frisancho Mariscal (1999, 2000), Papell et al. (2000) and Everaert (2001) apply univariate unit root tests that allow for structural breaks in the unemployment rate of OECD countries. 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.

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 tests of Im et al. (2005) and Carrión-i-Silvestre et al. (2005) that allow for up to two and five mean shifts, respectively. In this group we find Song and Wu (1998), León-Ledesma (2002) and Camarero and Tamarit (2004), for the case of no breaks, and Murray and Papell (2000), Strazicich et al. (2001), Camarero et al. (2006) and Romero-Ávila and Usabiaga (2008, 2009) for the case of structural breaks in the specification. The evidence tends to strongly favor stationarity in unemployment in the US, while there is mixed evidence for European unemployment. At the Spanish regional level, there is one study that employs panel methods to test for a unit root in Spanish regional unemployment. More specifically, Romero-Ávila and Usabiaga (2008) employed the panel stationarity test of Carrión-i-Silvestre et al. (2005) to test for a unit root in the unemployment rate of the 17 Spanish regions over the period 1976-2004, finding strong evidence of nonstationarity in the unemployment rate across the 17 Spanish regions.

However, it is remarkable that none of the aforementioned studies have allowed for strong forms of cross-sectional dependence. Therefore, this paper takes a step forward in this issue by employing the PANIC procedures proposed by Bai and Ng (2004, 2010). Unlike previous studies, we will be able to decompose the Spanish unemployment rates into idiosyncratic and common components as well as to determine their separate stochastic properties.

3. Econometric Methodology

The extant literature in the field has developed several panel unit root tests allowing for weak forms of cross-correlation like the non-linear instrumental variables panel unit root test of Chang (2002), the five bootstrap panel unit root tests of Smith et al. (2004) and the Breitung and Das (2005) test. However, the problem with these tests is that they only allow for contemporaneous short-run cross-correlation, but not for stronger forms such as cross-sectional cointegration.

Breitung and Pesaran (2008) point out that the use of dynamic linear factor models allow to make some parametric assumptions on the nature of the cross-sectional dependence, thereby allowing for much stronger forms of cross-dependence than the above tests. The panel procedures modelling cross-sectional dependence through a factor structure includes Pesaran (2007), Moon and Perron (2004) and Bai and Ng (2004, 2010), but only the latter is general enough to allow for cointegration across units, which implies that the observed series can contain common stochastic trends. A further advantage is that, unlike Pesaran (2007) and Moon and Perron (2004) tests, the PANIC approach is flexible enough as to allow for a different order of integration in the common factor(s) and idiosyncratic components.

The observed data on regional unemployment rates are modelled as the sum of a deterministic part, a common component and an idiosyncratic error term:

$$U_{it} = D_{it} + \lambda'_i F_t + e_{it} \tag{1}$$

where λ_i is an $r \times 1$ vector of factor loadings, F_t is an $r \times 1$ vector of common factors and e_{it} is the idiosyncratic component. D_{it} can contain a constant and a linear trend. Since λ_i and F_t can only be estimated consistently when $e_{it} \sim I(0)$, a model in first-differences like $\Delta U_{it} = \lambda'_i f_t + z_{it}$ is estimated, where $z_{it} = \Delta e_{it}$ and $f_t = \Delta F_t$. We then use principal components to estimate the common factors $\hat{f}_t$, the corresponding factor loadings $\hat{\lambda}_i$ and the residuals $\hat{z}_{it} = y_{it} - \hat{\lambda}'_i \hat{f}_t$, so that we preserve the order of integration of F_t and e_{it}. U_{it} is normalised for each cross-section unit to have a unit variance as in Bai and Ng (2002). The common factors and the residuals are then re-cumulated as follows: $\hat{F}_t = \sum_{s=2}^{t} \hat{f}_s$ and $\hat{e}_{it} = \sum_{s=2}^{t} \hat{z}_{is}$, which are used to test for a unit root in the common and idiosyncratic components, respectively.

Before conducting the tests for a unit root in the common and idiosyncratic components, the number of common factors must be determined. The preferred criterion is BIC_3 because Bai and Ng (2002, pp. 205-207) showed that the BIC_3 criterion exhibits very good properties even under a sufficiently general scenario in which the idiosyncratic errors are serially correlated and cross-correlated (see also Bai and Ng, 2010, p. 1097-1098).[3] The BIC_3 information criterion takes the form:

$$BIC_3(k) = \hat{\sigma}_e^2(k) + k\hat{\sigma}_e^2(k_{\max})\left(\frac{(N+T-k)\ln(NT)}{NT}\right) \tag{2}$$

where k is the number of factors included in the model, $\hat{\sigma}_e^2(k)$ is the variance of the estimated idiosyncratic components, and $\hat{\sigma}_e^2(k_{\max})$ is the variance of the idiosyncratic components estimated with the maximum number of factors (k_{max}=5). The second argument in the loss function represents the penalty for overfitting, which is thought to correct for the fact that models with a larger number of factors can fit at least as good as models with fewer common factors, but efficiency is reduced as more factor loading parameters are estimated (Bai and Ng, 2002). We choose the optimal number of common factors $\hat{k}$ as $\arg\min_{0 \le k \le 5} BIC_3(k)$. For the sake of completeness, we also present the IC_1, IC_2 and IC_3 panel information criteria proposed by Bai and Ng (2002), though in the case of discrepancies across criteria, the optimal number of common factors will be set on the basis of the BIC_3 criterion.

3.1. Analysis of the Idiosyncratic Component

To test for a unit root in the idiosyncratic components, Bai and Ng (2004) estimate standard ADF specifications of the following form:

$$\Delta \hat{e}_{it} = \delta_{i,0}\hat{e}_{i,t-1} + \sum_{j=1}^{p_i} \delta_{i,j}\Delta \hat{e}_{i,t-j} + u_{it} \tag{3}$$

[3] For further support for choosing this method in short panels, see Moon and Perron (2007).

They then employ the ADF t-statistic for testing $\delta_{i,0} = 0$, which is denoted by $ADF_{\hat{e}}^{c}(i)$ or $ADF_{\hat{e}}^{\tau}(i)$ for the cases of only constant and constant and linear trend in specification (1), respectively.[4] To raise statistical power, Bai and Ng (2004) proposed to employ pooled statistics based on the Fisher-type inverse chi-square tests of Maddala and Wu (1999) and Choi (2001), which can be safely applied to the idiosyncratic components when they are found to be cross-sectionally independent.[5] Letting $\pi_{\hat{e}}^{c}(i)$ be the p-value associated with $ADF_{\hat{e}}^{c}(i)$, we have:[6]

$$P_{\hat{e}}^{c} = -2\sum_{i=1}^{N} \log\ \pi_{\hat{e}}^{c}(i) \xrightarrow{d} \chi^2_{(2N)} \text{ for } N \text{ fixed, } T \to \infty \tag{4}$$

$$Z_{\hat{e}}^{c} = \frac{-\sum_{i=1}^{N} \log\ \pi_{\hat{e}}^{c}(i) - N}{\sqrt{N}} \xrightarrow{d} N(0,1) \text{ for } N, T \to \infty \tag{5}$$

In addition, we employ the two Moon and Perron (2004) type pooled tests using the PANIC residuals to estimate a bias-corrected pooled PANIC autoregressive estimator and a panel version of the Sargan-Bhargava (1983) statistic that employs the sample moments of the residuals without the need to estimate the pooled autoregressive coefficients. More specifically, the bias-corrected pooled PANIC autoregressive estimator can be constructed by estimating a panel regression in the cumulated idiosyncratic errors estimated via PANIC, i.e. $\hat{e}_{it}$. For the specification with no trend, the bias-corrected pooled PANIC autoregressive estimator ρ computed on the basis of the following pooled autoregressive specification $\hat{e}_{it} = \rho\hat{e}_{it-1} + \varepsilon_{it}$ equals:

$$\hat{\rho}^{+} = \frac{trace(\hat{e}_{-1}^{'}\hat{e}) - NT\hat{\lambda}_{\varepsilon}}{trace(\hat{e}_{-1}^{'}\hat{e}_{-1})} \tag{6}$$

and the PANIC pooled Moon-Perron type statistics are:

$$P_{a}^{c} = \frac{\sqrt{N}T(\hat{\rho}^{+} - 1)}{\sqrt{(2\hat{\phi}_{\varepsilon}^{4} / \hat{\omega}_{\varepsilon}^{4})}} \xrightarrow{d} N(0,1) \text{ for } N, T \to \infty \text{ with } N/T \to 0 \tag{7}$$

$$P_{b}^{c} = \sqrt{N}T(\hat{\rho}^{+} - 1)\sqrt{\frac{1}{NT^2} trace(\hat{e}_{-1}^{'}\hat{e}_{-1})\frac{\hat{\omega}_{\varepsilon}^{2}}{\hat{\phi}_{\varepsilon}^{4}}} \xrightarrow{d} N(0,1) \text{ for } N, T \to \infty \text{ with } N/T \to 0 \tag{8}$$

where $\hat{\omega}_{\varepsilon}^{2}$ and $\hat{\lambda}_{\varepsilon}$ are cross-sectional averages of the consistent estimates of the long-run and one-sided variance of ε_{it}, respectively, and $\hat{\phi}_{\varepsilon}^{4}$ is the cross-sectional average of $\hat{\omega}_{\varepsilon,i}^{4}$.

[4] The asymptotic distribution of $ADF_{\hat{e}}^{c}(i)$ coincides with the Dickey-Fuller (DF) distribution for the case of no constant, while that of $ADF_{\hat{e}}^{\tau}(i)$ is proportional to the reciprocal of a Brownian bridge.

[5] Note that under a factor structure, it is not appropriate to pool individual unit root tests for the observed series, since the limiting distribution of the test would contain terms that are common across cross-sectional units.

[6] The same holds for the case of a trend, where $\pi_{\hat{e}}^{\tau}(i)$ is the p-value associated with $ADF_{\hat{e}}^{\tau}(i)$. The pooled statistics for the trend specification are denoted as $P_{\hat{e}}^{\tau}$ and $Z_{\hat{e}}^{\tau}$.

For the specification with a trend, we have:

$$\hat{\rho}^{+} = \frac{trace(\hat{e}_{-1}^{'}\hat{e})}{trace(\hat{e}_{-1}^{'}\hat{e}_{-1})} + \frac{3}{T}\frac{\hat{\sigma}_{\varepsilon}^{2}}{\hat{\omega}_{\varepsilon}^{2}} \tag{9}$$

$$P_{a}^{\tau} = \frac{\sqrt{N}T(\hat{\rho}^{+}-1)}{\sqrt{(36/5)\hat{\phi}_{\varepsilon}^{4}\hat{\sigma}_{\varepsilon}^{4}/\hat{\omega}_{\varepsilon}^{8}}} \xrightarrow{d} N(0,1) \text{ for } N,T \to \infty \text{ with } N/T \to 0 \tag{10}$$

$$P_{b}^{\tau} = \sqrt{N}T(\hat{\rho}^{+}-1)\sqrt{\frac{1}{NT^{2}}trace(\hat{e}_{-1}^{'}\hat{e}_{-1})\frac{5}{6}\frac{\hat{\omega}_{\varepsilon}^{6}}{\hat{\sigma}_{\varepsilon}^{4}\hat{\phi}_{\varepsilon}^{4}}} \xrightarrow{d} N(0,1) \text{ for } N,T \to \infty \text{ with } N/T \to 0 \tag{11}$$

where $\hat{\sigma}_{\varepsilon}^{2}$ is the cross-sectional average of the consistent estimate of the short-run variance of ε_{it}.

In addition, the PANIC pooled Sargan-Bhargava (1983) statistic does not require the estimation of ρ. It takes the following form for the specification without trends:

$$PMSB^{c} = \frac{\sqrt{N}\left(trace\left(\frac{1}{NT^{2}}\hat{e}^{'}\hat{e}\right) - \hat{\omega}_{\varepsilon}^{2}/2\right)}{\sqrt{\hat{\phi}_{\varepsilon}^{4}/3}} \xrightarrow{d} N(0,1) \text{ for } N,T \to \infty \text{ with } N/T \to 0 \tag{12}$$

where $\hat{\omega}_{\varepsilon}^{2}/2$ estimates the asymptotic mean of $(1/NT^{2})trace(\hat{e}^{'}\hat{e})$ and the denominator estimates its standard deviation. For the specification with trends, we have:

$$PMSB^{\tau} = \frac{\sqrt{N}\left(trace\left(\frac{1}{NT^{2}}\hat{e}^{'}\hat{e}\right) - \hat{\omega}_{\varepsilon}^{2}/6\right)}{\sqrt{\hat{\phi}_{\varepsilon}^{4}/45}} \xrightarrow{d} N(0,1) \text{ for } N,T \to \infty \text{ with } N/T \to 0 \tag{13}$$

where the variables $\hat{\omega}_{\varepsilon}^{2}$ and $\hat{\phi}_{\varepsilon}^{4}$ are estimated from residuals obtained from $\hat{\varepsilon} = \hat{e} - \hat{\rho}\hat{e}_{-1}$, where $\hat{\rho}$ represents the pooled least squares estimator based on $\hat{e}$.

The advantage of these six pooled statistics is that they do not require least squares linear detrending that could lead to a fall in statistical power. Notwithstanding, the statistics for the constant only specification are expected to perform better relative to those associated with the constant and trend specification.

3.2. Analysis of the Common Component

To test for nonstationarity in the common factors, Bai and Ng (2004) propose using an ADF statistic for the case of a single common factor (*k=1*) by estimating an ADF specification for $\hat{F}_{t}$ with the same deterministic components as in model (1):

$$\Delta\hat{F}_{t} = D_{t} + \gamma_{i,0}\hat{F}_{t-1} + \sum_{j=1}^{p}\gamma_{i,j}\Delta\hat{F}_{t-j} + v_{it} \tag{14}$$

The corresponding ADF t-statistics are denoted by $ADF_{\hat{F}}^{c}$ and $ADF_{\hat{F}}^{\tau}$ and follow the limiting distribution of the Dickey-Fuller (1979) test for the specifications with only a constant, and a constant and a trend, respectively. For the case of multiple common factors (*k>1*), the number of common stochastic trends in the common factors are determined using the modified rank tests given by the filter test MQ_f and the corrected test MQ_c.

4. Empirical Results from the PANIC Procedures

The measure of unemployment used is the *Encuesta de Población Activa* (EPA) unemployment rate, which is available on a quarterly basis for the period 1976(3)-2013(2) and is provided by the National Statistical Office (*Instituto Nacional de Estadística*). The unemployment rates are obtained by dividing the number of unemployed by the corresponding labour force figure.

We now proceed to report the results obtained from the application of the PANIC techniques. Before that, we conduct a formal analysis of the presence of cross-sectional dependence in regional unemployment rate innovations by applying the tests for cross-dependence developed by Breusch and Pagan (1980) and Pesaran (2004). Pesaran's test is given by $CD = \sqrt{2T/(N(N-1))}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \xrightarrow{d} N(0,1)$, where $\hat{\rho}_{ij}$ represents the average of pair-wise correlation coefficients of OLS residuals obtained from standard ADF regressions for each individual with the order of the autoregressive model being selected using the *t-sig* criterion in Ng and Perron (1995), with the maximum number of lags set at $p = 4(T/100)^{1/4}$. Breusch and Pagan (1980) employ the Lagrange Multiplier (LM) statistic $CD_{lm} = T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}^2 \xrightarrow{d} \chi^2_{N(N-1)/2}$ to also test the null hypothesis of cross-sectionally independent errors.

The LM test of Breusch and Pagan (1980) equals 67.99 and 69.44 for the specifications without and with trends, and the respective values of the CD test of Pesaran (2004) are 32.60 and 32.97. The null hypothesis of cross-sectionally independent errors is thus strongly rejected at the 1% significance level in all cases. Hence, the use of the PANIC methodology is justified on the grounds that it allows for cross-sectional dependence, thereby avoiding large size distortions in the tests.

Prior to testing for a unit root in the idiosyncratic and common components, the factors need to be estimated through principal components. We then select the optimal number of common factors present in the panel. As noted above, this is done with the BIC_3 procedure developed in Bai and Ng (2002), which appears to perform better than alternative information criteria. Setting a maximum number of factors to five, the BIC_3 criterion selects one common factor, as shown in Table 1. This implies that if we found the common factor to be nonstationary and the idiosyncratic components to be stationary, the evidence would support the presence of pair-wise cointegration among the 17 regional unemployment rates due to the existence of a common stochastic trend driving them all.

[Insert Table 1 about here]

Table 2 reports the results of the univariate ADF statistics for both the observed unemployment rate series (column 3) and the idiosyncratic components (column 4), the ADF statistic for the common factor in addition to the pooled PANIC statistics (those proposed in Bai and Ng (2004) and those in Bai and Ng (2010)). For the sake of completeness, we report the results for the specification with and without trends, though admittedly the specification with no deterministic trends is expected to be more meaningful for the analysis of unemployment rates. Thus we base our main conclusions on the specification without trends, though the results are fairly robust to the inclusion of deterministic trends.

We begin with the direct testing for a unit root in the observed data (column 3 of Table 2) and the results clearly point to the nonstationarity in the observed unemployment rates, as we fail to reject the unit root null hypothesis for any of the 17 individual series. Therefore, the next step is to establish whether the nonstationary behaviour of Spanish regional unemployment is caused by the common factor and/or the idiosyncratic component. Column 4 presents the univariate ADF statistic, the pooled statistics based on the Fisher-type inverse chi-square tests of Maddala and Wu (1999) and Choi (2001), and the pooled PANIC tests of Moon and Perron (2004) and Sargan and Bhargava (1983) type proposed by Bai and Ng (2010), for testing the unit root null hypothesis for the idiosyncratic component. The evidence from the application of the univariate ADF statistic indicates that the unit root hypothesis can be rejected for eight regions at least at the 10% significance level: Andalusia, Asturias, Catalonia, Extremadura, Rioja and Valencian Community at the 5% level and Canary Islands and Madrid at the 10% level. It is also worth noting that the five pooled statistics from Bai and Ng (2004, 2010) are able to strongly reject the joint unit root null hypothesis at the 1% level of significance. This, in turn, implies that the idiosyncratic component can be considered jointly stationary.

Regarding the analysis of the time series properties of the common factor, the ADF statistic for the common factor is presented in the bottom part of the table. Remarkably, the ADF test for a unit root in the common factor fails to reject the null even at the 10% level. This result is further supported by the application of the IPC_1, IPC_2 and IPC_3 information criteria of Bai (2004) to establish the number of *non-stationary* common factors in the panel (setting the maximum number of factors to five), which clearly favour the existence of only one common stochastic factor. The above decomposition of the original series into the idiosyncratic and common components thus indicates that the source of nonstationarity is primarily a common stochastic trend driving the nonstationarity in the observed unemployment rate series. This finding together with the presence of jointly stationary idiosyncratic components provides strong evidence of pairwise cointegration across Spanish regional unemployment rates due to the existence of a strong common component linking all regional unemployment series together.[7]

Regarding the specification with a trend (see columns 5 to 7), the same results essentially follow, though in this case the evidence of joint stationarity in the idiosyncratic component is tempered as the joint unit root null can only be rejected at the 1% with the pooled statistic based on the Fisher-type inverse chi-square tests of Maddala and Wu (1999) and Choi (2001), but not with the pooled PANIC tests of Moon and Perron (2004) and Sargan and Bhargava (1983) type proposed by Bai and Ng (2010). Note, though, that the latter tests appear to perform better in terms of statistical power in the specification without trends, even if they do not require least squares detrending for their construction (see Bai and Ng, 2010).

[Insert Table 2 about here]

Finally, columns 8 and 9 present the ratio of the standard deviation of the idiosyncratic component to the standard deviation of the observed data (both expressed in first-differences) in addition to the standard deviation of the common to the idiosyncratic component. These statistics should be informative about the relative importance of the common and idiosyncratic components. A first ratio close to one coupled with a relatively small second ratio would indicate that region-specific variations prevail in that region over the common component. In relative terms, this appears to be the case for Navarre and to a lower extent Balearic Islands, which are mainly driven by idiosyncratic variations. In contrast, unemployment rates in the other regions are (to a differing degree) more affected by the common component.

Taken as a whole, the PANIC analysis of the stochastic properties of our panel of Spanish regional unemployment rates appears to render clear-cut evidence for the existence of hysteresis in Spanish unemployment caused by the prevalence of a nonstationary common factor. This in turn implies that the 17 regional unemployment rates appear closely linked together due to the presence of a common stochastic trend driving them all.

5. Persistence of Shocks to the Common and Idiosyncratic Components of Spanish Regional Unemployment

As a complement to the previous PANIC analysis, we next compute median-unbiased estimates of the persistence parameter and the half-life of a shock for both the common and idiosyncratic components that form the observed regional unemployment rate series. These two persistence measures are accompanied by their respective 90% confidence intervals, which serve as a measure

[7] In an unpublished appendix available from the authors upon request we present a figure depicting the evolution of the observed regional unemployment rates and the common stochastic trend. It shows that the common trend can track very closely the observed unemployment rates for all the regions of Spain.

of their accuracy.[8] The half-life of a shock –that measures the number of years for a unit impulse to dissipate by one half– is computed directly from the impulse-response function following Cheung and Lai (2000).

The median-unbiased method, originally proposed by Andrews (1993) for AR(1) processes and later extended by Andrews and Chen (1994) for autoregressive processes of order greater than one, can account for the downward bias associated with least squares estimates of the persistence parameter. This bias is caused by the skewness to the left in the distribution of the estimators of the persistence parameter in autoregressions, which leads the median to exceed the mean. This makes the median a better measure of central tendency than the mean.[9] Andrews (1993) and Andrews and Chen (1994) correct the least squares estimator for the median bias as follows. If we let $\hat{\alpha}$ be an estimator of the true value of the persistence parameter α, with a median function $[m(\alpha)]$ which is unique $\forall \alpha \in (-1,1]$, the median-unbiased estimator of α can be defined as follows:

$$\hat{\alpha}_{MU} = \left\{ \begin{array}{c} 1 \quad if \quad \hat{\alpha} > m(1), \\ m^{-1}(\hat{\alpha}) \quad if \quad m(-1) < \hat{\alpha} < m(1), \\ -1 \quad if \quad \hat{\alpha} \le m(-1) \end{array} \right\} \tag{15}$$

where $m(-1) = \lim_{\alpha \to -1} m(\alpha)$, and $m^{-1} : (m(-1),\ m(1)] \to (-1,1]$ is the inverse function of $m(.)$ satisfying $m^{-1}(m(\alpha)) = \alpha$ for $\alpha \in (-1,\ 1]$. So if there is a function for which each true value of α renders the median value of $\hat{\alpha}$ implied by the 0.50 quantile, the inverse of this function can be used to obtain median-unbiased estimates of α. The value of α which gives the least squares estimator with a median value of $\hat{\alpha}$ can thus be used as the median-unbiased estimator of the persistence parameter.[10] The 90% confidence intervals are obtained on the basis of the 0.05 and 0.95 quantiles of $\hat{\alpha}$ and are used as a measure of the accuracy of the persistence parameter estimate.

Median-unbiased estimates of half-lives and associated confidence intervals are also provided.[11] For AR(1) processes, half-lives are obtained using the formula $HL = \ln(0.5)/\ln(\alpha)$. For higher-order autoregressive processes, the estimate of α is no longer exact –as the series no longer decays monotonically–, and must be derived approximately from the impulse-response function.

[8] Note that the previous analysis focused exclusively on establishing whether the sum of the autoregressive coefficients –i.e. the persistence parameter– is equal or less than unity, rather than determine an exact measure of persistence.

[9] The estimation of a more accurate measure of the persistence of shocks to a series through median-unbiased half-lives along with confidence intervals is essential in addressing the low power of univariate unit root tests. They can inform about whether non-rejection of the unit root null is caused by low power or is due to the existence of a high degree of uncertainty about the true value of the persistence parameter.

[10] Thus, there is a 50% probability that the confidence interval from -1 to $\hat{\alpha}_{MU}$ contains the true value of α and another 50% probability that the confidence interval from $\hat{\alpha}_{MU}$ to 1 contains the true α.

[11] The median-unbiased property that characterise the point estimates of the persistence parameter carries over to any scalar measure of persistence calculated from them, as is the case for the half-life point estimates.

Columns 3 and 4 of Table 3 present the median-unbiased estimates of the persistence parameter ($\hat{\alpha}_{MU}$) and half-lives along with their respective 90% confidence intervals for the ADF specifications without a linear trend, whereas columns 6 and 7 do so for the specification with trends. Since the results remain unaltered when a linear trend is included, we focus on the evidence obtained from a specification without a trend.[12]

Let us start commenting on the median-unbiased estimates of the persistence parameter associated with the idiosyncratic series. We find four regions with a persistence parameter below 0.96 (Asturias, Extremadura, Murcia and Rioja), seven regions with a persistence parameter greater than 0.96 and lower than unity (Andalusia, Aragon, Canary Islands, Cantabria, Catalonia, Madrid and Valencian Community), and six regions with an estimate of the persistence parameter of one (Balearic Islands, Basque Country, Castilla-Leon, Castilla-La Mancha, Galicia and Navarre), which is consistent with the unit root hypothesis. This result lends support to the existence of stationary idiosyncratic series in all but six regions in which $\hat{\alpha}_{MU} = 1$. Besides, the lower bound of the 90% confidence interval appears below 0.96 in all regions with the exception of the Basque Country (0.985) and Castilla-La Mancha (1.00). The upper bound of the confidence interval is always one with the exceptions of Madrid (0.909), Rioja (0.965) and Asturias (0.989).

As far as the median-unbiased estimates of the half-life are concerned, there is a considerable degree of variation in the degree of persistence associated with the idiosyncratic series, though it is worth pointing out that the half-life point estimate is always a finite number, except in the case of the Basque Country and Castilla-La Mancha that exhibit infinite half-lives (in congruence with the presence of a unit root in these two series). This implies that any region-specific shock hitting the unemployment rates of these two regions will exert a permanent effect on them. In stark contrast stand the other regions, particularly those with the lowest persistence, with a half-life equal or below 2.5 years (Andalusia, Aragon, Asturias, Balearic Islands, Cantabria, Murcia and Rioja), for which the effect of half of a shock will vanish in 2.5 years or less. Seven other regions exhibit half-lives above 2.5 and below 5 years (Canary Islands, Castilla-Leon, Catalonia, Extremadura, Madrid, Navarre and Valencian Community) and only Galicia presents a finite half-life above this (7 years). To measure the persistence of shocks to idiosyncratic series for the whole panel, we compute the median half-life which equals 3.25 years, implying a speed of convergence of 19% per year. The median lower bound of the 90% confidence interval for half-lives equals 1.75 years, which indicates that shocks decay at a rate of 33% per year.[13] The upper bound of the confidence interval is infinite in all but six regions (Andalusia, Asturias, Catalonia, Extremadura, Rioja and Valencian

[12] If after 40 years (480 months) shocks have not vanished by half, the half-life is assumed to be infinite.
[13] The mean half-life equals 3.05 years, which implies a speed of convergence of 20% per year; the mean lower bound equals 2.25 years, with an associated speed of convergence of 27% per year.

Community). The width of the confidence interval associated with the half-life point estimates reflects a high degree of uncertainty in their estimation, since they would be consistent with a wide range of degrees of persistence. Summing up, the relatively low value of half-lives for the 15 region-specific idiosyncratic series exhibiting finite half-lives, coupled with a median value of half-lives of 3.25 years for these 15 series, largely stand in favour of the joint stationarity of the idiosyncratic component found in the PANIC analysis.

Regarding the analysis of the common factor we find evidence of a median-unbiased estimate of the persistence parameter of one, which is congruent with the unit root hypothesis (as found in the PANIC analysis). However, when we measure the degree of persistence associated with the common factor through the half-life, we find an estimate of 12.25 years which appears below a value of infinity (as would be predicted in the unit root case). Still, our half-life estimates indicate that the rate at which common shocks hitting regional unemployment rates decay equals 6% per year. This implies a relatively high degree of persistence in the common component as it takes more than a decade for the effect of a shock hitting regional unemployment to vanish by half. This time period is above a normal business cycle, thus reflecting the difficulty for Spanish regional unemployment rates to reach pre-shock levels even after the recovery of the economy has set in.

[Insert Table 3 about here]

Taken as a whole, the median-unbiased estimation of the persistence parameter and the half-life of shocks to the idiosyncratic series appear to render evidence largely consistent with that obtained from the application of the PANIC approach, which supports the joint stationarity of the idiosyncratic component. The only exceptions were the Basque Country and Castilla-La Mancha that exhibited infinite half-lives. In addition, the evidence gathered in this section about the persistence of the common component appears largely in line with that obtained in the PANIC analysis, since the median-unbiased estimate of the persistence parameter is unity (lending support to the unit root hypothesis) and the half-life of a shock to the common factor is relatively high (for a variable like unemployment) as given by 12.25 years. These results suggest that all the policy measures implemented with the aim of removing any impediment to the adjustment mechanisms in response to an increase in unemployment should be welcome. Thus, greater labour market flexibility is called for as a way to reduce the sluggishness in adjusting to adverse common shocks.[14]

Recent work by IMF economists, in accordance also with OECD and the European Commission, comes to reinforce the view that product and labour market reforms are urgently needed –especially

[14] This is particularly important in the case of the Basque Country and Castilla-La Mancha that exhibit the highest persistence in the idiosyncratic component according to half-life estimates. As such, it should be more difficult for these two regions to prevent unemployment rises from becoming permanent. By the same token, appropriate stabilisation policy management at the different levels should be more relevant in the regions with the highest persistence.

in countries with malfunctioning institutions– if large rises in unemployment are to be averted in the aftermath of a sizable negative shock, like the Great Recession.

In particular, Jaumotte (2011) points to two institutions as the main culprits of rendering the Spanish labour market so dysfunctional: the intermediate-level collective bargaining and the high protection of permanent workers. This author then centres on how supply-side reforms could combine to substantially cut unemployment and the share of temporary workers. By employing standard regression methods, she finds that either strongly centralising or decentralising the bargaining system would give rise to a significant decrease in the unemployment rate (10 and 7 percentage points respectively). Reductions in unemployment benefits (or in the length of time of perception) and in the tax wedge undertaken together would also cut the unemployment rate, although in a more modest manner. In addition, the introduction of some flexibility-enhancing measures in the product markets would help reduce the unemployment rate somewhat. When the focus is shifted to the high share of temporary workers, the measure that would lead this share to decrease is a reduction in the employment protection on permanent contracts. Again, lowering the tax wedge and the replacement rates of the unemployment benefits would also help. Finally, it is worth mentioning that all these measures are likely to be highly complementary, which means that carrying them out in one stroke would reduce unemployment as well as the weight of temporary employment.

Blanchard et al. (2013) address the issue of whether the set of labour market policies that the IMF has recommended (or "imposed" in some cases) to advanced economies during the Great Recession has proven right. These policies can be split into micro and macro flexibility-enhancing measures, and both types of flexibility are truly critical to achieving higher productivity and lower unemployment. Micro flexibility refers to shift protection from jobs towards workers so as to ease the desirable reallocation of labour. This flexibility amounts to designing labour institutions in a way that the correct level of flexibility for firms and security for workers can be achieved. For example, countries where employment protection is still high should bring firing costs down into alignment with OECD average. At the same time, a generous unemployment insurance system, as long as it decreases with duration, coupled with a full-fledged set of active labour market policies (ALMPs) that ensures that workers will be put back to work once they lose their jobs, should be provided. The macro flexibility concerns the ability to adjust to negative macroeconomic shocks without letting unemployment increase much. The crucial element here is the collective bargaining structure. It is well-established in the literature that intermediate-level systems (sector or industry level) are the most unemployment-prone ones. According to this, moving towards the extremes proves convenient. For instance, in a centralised system, national agreements could be fostered in the face of adverse shocks whereby wages can grow in line with labour productivity.

5. Conclusions

This paper has applied the PANIC framework developed by Bai and Ng (2004) and further extended by Bai and Ng (2010) to the unemployment rate of the Spanish regions spanning the period 1976-2013. This methodology has not only enabled us to allow for strong forms of cross-sectional correlation among the individual unemployment rate series, but also to establish the source of hysteresis in Spanish regional unemployment, i.e. whether it stems from the common component and/or the region-specific idiosyncratic series. The PANIC analysis has provided evidence for the joint stationarity of the idiosyncratic component as well as for the existence of a common stochastic factor driving the observed unemployment rate series. The findings from the median-unbiased estimation of the persistence parameter and the half-life of a shock associated with both components largely accord with those obtained from the PANIC analysis.

It is worth highlighting that our results carry some important policy implications. In short, according to our results, to combat the source of hysteresis in Spanish regional unemployment, it is necessary to implement policy measures aimed at reducing the sluggishness of the Spanish labour market in adjusting to adverse shocks hitting the common component of regional unemployment rates. This is because there exist important rigidities in the labour and goods and services markets that prevent regional unemployment rates from returning to pre-shock levels, thus making unemployment rises largely permanent.

In the spirit of the aforementioned line of advice from international organisations, have recent labour market reforms been conducted in Spain (2010, 2011 and 2012)[15] aiming, among other things, at making wages more responsive to increases in unemployment. Mainly, these reforms have facilitated the objective dismissal on economic grounds; fostered company-based bargaining whereby firms, under certain conditions, can opt out from the wages and other labour conditions bargained at a superior level; and reduced the gap in the severance pay between open-ended and fixed-term contracts.[16] In short, the reforms have tried to reallocate bargaining power from workers and unions to firms, although the way these changes are brought into fruition remains to be seen, especially the interpretation of these new rules by labour courts. Furthermore, the contentious debate over the proposal of introducing a single open-ended labour contract, with increasing –with tenure– severance pay, for all new hiring, remains prominent today. On the other hand, over the past four years wages and unit labour costs have experienced a downward adjustment in Spain relative to other Eurozone members and some steps have recently been taken towards helping attenuate the

[15] See Bentolila et al. (2012b, c).

[16] These reforms have also removed a wide range of administrative procedures that previously hindered entrepreneurial activity, adopted partially the German model of working hours, taken some preliminary steps to adopt the Austrian capitalisation fund scheme, and limited the inertia of labour agreements. among other minor measures. However, a greater emphasis should be placed on ALMPs. Likewise, the reforms in goods and services markets implemented up to now have not gone far enough, e.g. those with regard to competition enforcement.

indexation of the economy. We shall see whether all these new factors significantly affect the observed patterns in the Spanish labour market, inter alia reducing the high unemployment persistence.

TABLES

Table 1: Information Criteria

Number of factors (k)	$IC_1(k)$	$IC_2(k)$	$IC_3(k)$	$BIC_3(k)$
0	-0.0191	-0.0191	-0.0191	0.9811
1	-0.3382	-0.3310	-0.3503	0.7508*
2	-0.3413*	-0.3269*	-0.3655*	0.8042
3	-0.3097	-0.2881	-0.346	0.8869
4	-0.2967	-0.2680	-0.3451	0.9701
5	-0.3012	-0.2653	-0.3617	1.0561

Note: * represents the lowest value of the information criteria. See the text for the equations associated with the information criteria.

Table 2: Panel Analysis of Nonstationarity in Idiosyncratic and Common Components

	No Trend Specification			Trend Specification				
Region	k	$ADF_y^c(i)$	$ADF_{\hat{e}}^c(i)$	k	$ADF_y^\tau(i)$	$ADF_{\hat{e}}^\tau(i)$	$\frac{\sigma(\Delta\hat{e}_{it})}{\sigma(\Delta y_{it})}$	$\frac{\sigma(\lambda'_i F_t)}{\sigma(\hat{e}_{it})}$
Andalusia	3	-2.382	-1.964**	2	-2.381	-1.964	0.468	3.832
Aragon	5	-1.929	-1.617	5	-1.870	-1.251	0.639	3.112
Asturias	3	-1.946	-2.191**	3	-1.943	-2.176	0.562	2.993
Balearic Islands	8	-1.768	-1.204	8	-1.897	-1.851	0.745	1.836
Basque Country	6	-2.049	-0.341	6	-2.647	-3.979***	0.590	1.024
Canary Islands	8	-1.643	-1.672*	8	-1.662	-1.733	0.634	1.908
Cantabria	8	-2.062	-1.518	2	-2.059	-1.631	0.673	2.226
Castilla-Leon	3	-2.428	-1.328	3	-2.408	-1.068	0.384	3.193
Castilla-La Mancha	4	-1.767	0.205	4	-1.882	-0.983	0.459	2.506
Catalonia	5	-2.461	-2.211**	4	-2.443	-2.520*	0.551	1.882
Extremadura	7	-1.798	-2.347**	8	-1.819	-2.339*	0.501	2.692
Galicia	3	-2.160	-1.513	3	-2.405	-1.421	0.642	1.308
Madrid	5	-2.477	-1.700*	2	-2.494	-2.344*	0.574	2.335
Murcia	7	-2.124	-1.404	5	-2.216	-1.829	0.723	2.728
Navarre	5	-1.785	-1.184	5	-1.716	-1.152	0.839	1.183
Rioja	0	-2.135	-2.066**	0	-2.163	-2.233	0.717	3.293
Valencian Community	3	-2.141	-1.984**	3	-2.194	-2.326*	0.473	3.144
Critical Values								
1%		-3.430	-2.580		-3.960	-3.167		
5%		-2.860	-1.950		-3.410	-2.577		
10%		-2.570	-1.620		-3.120	-2.314		
Bai and Ng (2004) Pooled Statistics								
		$P_{\hat{e}}^c$	80.155***		$P_{\hat{e}}^\tau$	63.191***		
		$Z_{\hat{e}}^c$	5.597***		$Z_{\hat{e}}^\tau$	3.540***		
Bai and Ng (2010) Pooled Statistics								
		P_a^c	-9.290***		P_a^τ	-1.305		
		P_b^c	-4.527***		P_b^τ	-1.145		
		$PMSB^c$	-2.435***		$PMSB^\tau$	-0.952		

Common Factor Analysis		**Critical Values**					**Critical Values**		
Statistic		**1%**	**5%**	**10%**	**Statistic**		**1%**	**5%**	**10%**
$ADF_{\hat{F}}^{c}$	-2.296	-3.430	-2.860	-2.570	$ADF_{\hat{F}}^{\tau}$	-2.301	-3.960	-3.410	-3.120

Note: The augmented autoregressions employed in the ADF analysis set a maximum lag-order to $p = 4(T/100)^{1/4}$. The information criterion BIC_3 has chosen an optimal rank equal to 1 for the tests of Bai and Ng (2004) and Bai and Ng (2010). $P_{\hat{e}}$ is distributed as χ^2_{34}, with 1%, 5% and 10% critical values of 56.061, 48.602 and 44.903, respectively. $Z_{\hat{e}}$, P_a, P_b and $PMSB$ are distributed as N(0, 1) with 1%, 5% and 10% critical values of -2.326, -1.645 and -1.282. ***, ** and * imply rejection of the null hypothesis at 1%, 5% and 10%, respectively.

Table 3: Half-lives of Shocks to the Common and Idiosyncratic Components of Spanish Regional Unemployment

	k	$\hat{\alpha}_{MU}$	HL_{IRF}	k	$\hat{\alpha}_{MU}$	HL_{IRF}
Idiosyncratic Series	**No Trend Specification**			**Trend Specification**		
Andalusia	3	0.961	2.00	3	1.000	2.00
		[0.883, 1.000]	[1.00, 13.75]		[0.883, 1.000]	[1.00, 13.75]
Aragon	5	0.968	1.00	5	1.000	1.00
		[0.864, 1.000]	[1.00, ∞]		[0.871, 1.000]	[1.00, ∞]
Asturias	3	0.916	2.25	3	0.919	2.25
		[0.850, 0.989]	[1.50, 5.00]		[0.845, 1.000]	[1.50, 5.00]
Balearic Islands	8	1.000	2.00	8	1.000	2.00
		[0.920, 1.000]	[1.75, ∞]		[0.914, 1.000]	[1.75, ∞]
Basque Country	6	1.000	∞	6	1.000	∞
		[0.985, 1.000]	[8.50, ∞]		[1,000, 1.000]	[8.50, ∞]
Canary Islands	8	0.974	4.25	8	1.000	4.25
		[0.915, 1.000]	[2.50, ∞]		[0.919, 1.000]	[2.50, ∞]
Cantabria	8	0.983	2.50	8	0.946	2.50
		[0.904, 1.000]	[1.00, ∞]		[0.873, 1.000]	[1.50, 8.50]
Castilla-Leon	3	1.000	3.50	3	1.000	3.50
		[0.926, 1.000]	[1.75, ∞]		[0.920, 1.000]	[1.75, ∞]
Castilla-La Mancha	4	1.000	∞	4	1.000	∞
		[1.000, 1.000]	[3.25, ∞]		[1,000, 1.000]	[3.25, ∞]
Catalonia	5	0.975	4.50	5	1.000	4.50
		[0.929, 1.000]	[2.75, 24.75]		[0.926, 1.000]	[2.75, 25.75]
Extremadura	7	0.938	3.25	7	0.946	3.25
		[0.865, 1.000]	[2.50, 7.75]		[0.874, 1.000]	[2.50, 10.25]
Galicia	3	1.000	7.00	3	1.000	7.00
		[0.947, 1.000]	[3.50, ∞]		[0.953, 1.000]	[3.50, ∞]
Madrid	5	0.979	4.00	5	1.000	4.00
		[0.909, 0.909]	[2.25, ∞]		[0.916, 1.000]	[2.00, ∞]
Murcia	7	0.956	1.00	7	1.000	1.00
		[0.865, 1.000]	[1.00, ∞]		[0.890, 1.000]	[0.75, ∞]
Navarre	5	1.000	4.25	5	1.000	4.25
		[0.930, 1.000]	[1.75, ∞]		[0.930, 1.000]	[1.75, ∞]
Rioja	0	0.849	1.00	0	0.846	1.00
		[0.740, 0.965]	[0.50, 1.75]		[0.754, 1.000]	[0.50, 1.75]
Valencian Community	3	0.974	3.25	3	1.000	3.25
		[0.899, 1.000]	[1.75, 41.25]		[0.910, 1.000]	[1.75, 41.25]

Common Component						
Common Factor	3	1.000	12.50	2	1.000	12.50
		[0.975, 1.000]	[6.25, ∞]		[0.978, 1.000]	[6.25, ∞]

Note: The degree of augmentation of the ADF regression is selected with a general-to-specific criterion setting a maximum lag-order equal to $p = 4(T/100)^{1/4}$ (for instance, k=4 implies an AR(5) process). The entries in columns 3 and 6 are the approximately median-unbiased point estimates of the persistence parameter computed as in Andrews and Chen (1994). We use 2,000 iterations in our numerical simulations in order to generate quantile functions of $\hat{\alpha}$. The entries in columns 4 and 7 are the median-unbiased half-life point estimates. Figures in brackets represent the 90% confidence intervals associated with the persistence parameter and the half-life of a shock. Half-lives are measured in years and are computed directly from the impulse-response function.

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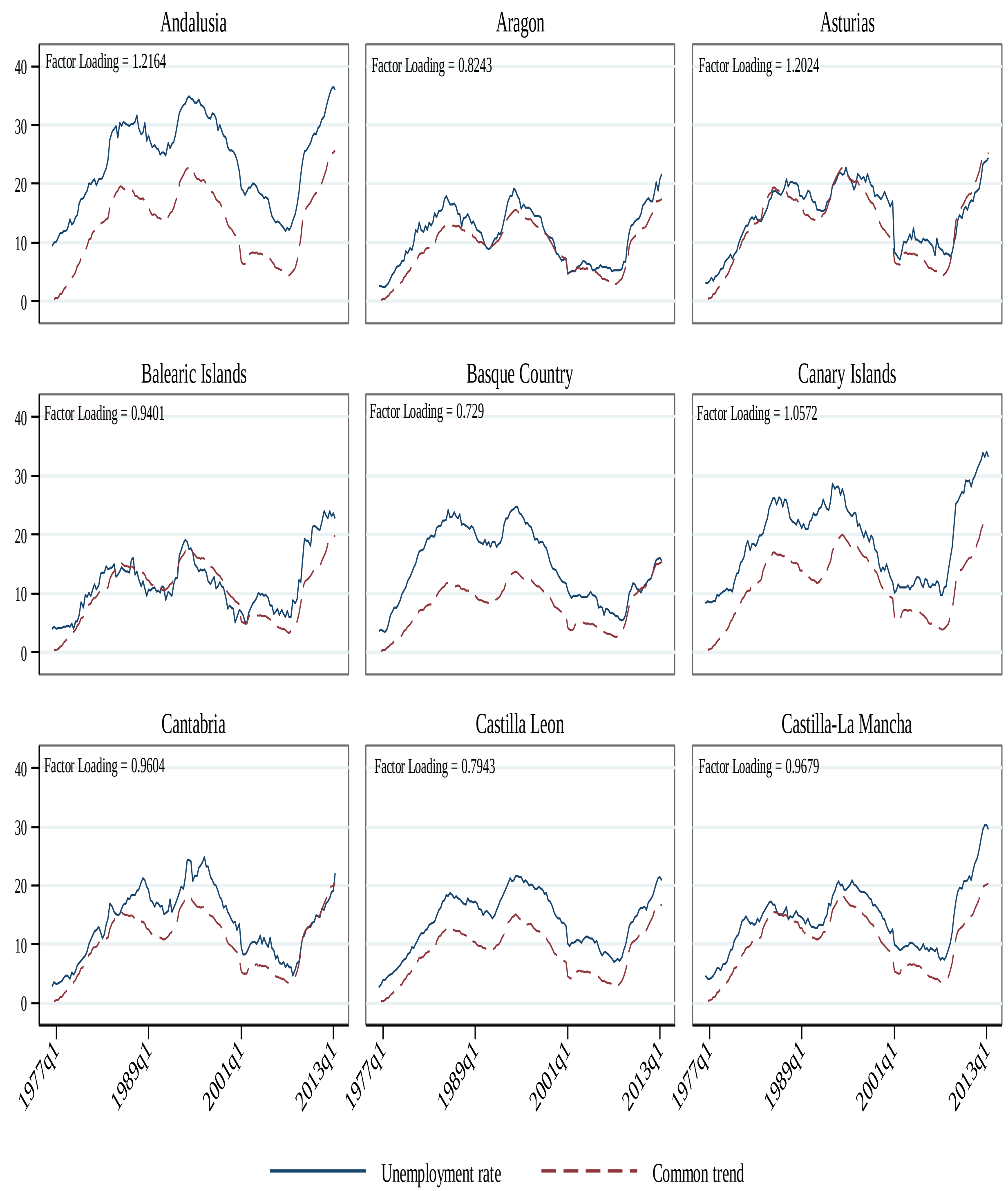
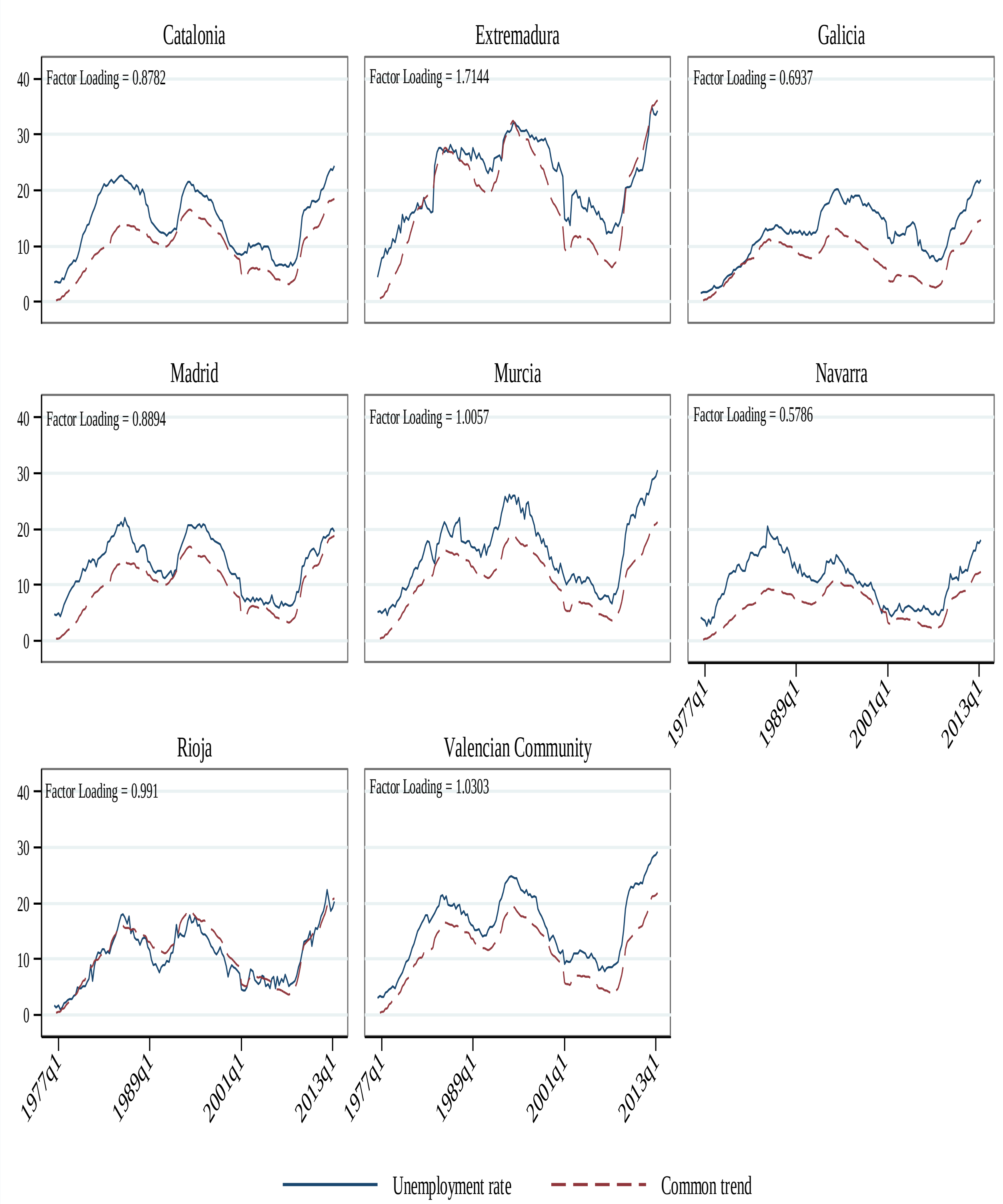
Appendix

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Extremadura
Factor Loading = 1.7144
Galicia
Factor Loading = 0.6937
Madrid
Factor Loading = 0.8894
Murcia
Factor Loading = 1.0057
Navarra
Factor Loading = 0.5786
Rioja
Factor Loading = 0.991
Valencian Community
Factor Loading = 1.0303
0
10
20
30
40
1977q1
1989q1
2001q1
2013q1
Unemployment rate
Common trend