

# Unemployment Hysteresis and Cycle Asymmetry: A case study

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## Abstract

This paper proposes an exhaustive step-by-step methodology guide to study in detail the behaviour of unemployment time-series, applied to the Portuguese case. In the first part of the chapter, we assess if the series follows a unit root process as to confirm the hysteresis hypothesis. In the second part, we develop a baseline nonlinear model to test for the asymmetric behaviour of unemployment across cycle phases. Our results lend support for hysteresis and show that the Portuguese unemployment dynamics is better described by a nonlinear with three types of transition variables: (a) annual change of cyclical unemployment (b) annual change of unemployment; and (c) annual GDP growth rate. We also analyse the impact of Labour Market Institutions (LMI) on its asymmetric behaviour, concluding that LMI can affect not only the regimes but also the equilibrium unemployment rate. Thus, strong enough short-run increases in unemployment, as those observed during the recent fiscal consolidation effort, have non-negligible impacts on raising the Portuguese natural rate of unemployment.

**Keywords:** Unemployment; Hysteresis; Nonlinear models; Cycle-asymmetric adjustment; Portugal.

**JEL Codes:** E24; C22; C52; C32.

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# 1 Introduction

The concept of hysteresis in unemployment was first introduced in time series analysis by Blanchard and Summers (1986, 1987). The idea behind this term is simple but powerful: transitory shocks may have permanent effects on the unemployment rate. This theory was then rapidly embraced by several researchers as there were, at least, two main reasons for its popularity. First, the “hysteresis hypothesis” helped to explain the differences between the United States (US) and the European equilibrium unemployment rates (Layard et al., 1991). Second, it challenged the “non-accelerating inflation rate of unemployment (NAIRU)” hypothesis which suggested that shocks only have transition effects and, therefore, unemployment should be considered as a stationary, mean reverting, process (Phelps, 1967, 1972 and Friedman, 1968). The “first wave” of research focused on testing, empirically, these two hypotheses, and led to the following well-known results: (a) the hysteresis hypothesis seems to be confirmed for the European case, whereas the natural rate hypothesis is supported for the US (Blanchard and Summers, 1986; Røed, 1996, 2002); (b) it is possible to identify an asymmetric-cycle pattern in the dynamics of the unemployment rates - apparently, unemployment rapidly peaks in recessions, but it smoothly decreases in expansions (Granger and Teräsvirta, 1993). However, the “hysteresis hypothesis” was soon challenged by the “structuralist hypothesis” (Phelps, 1994; Perron, 1989). This theory argues, instead, that the unemployment rate might be subject to occasional but persistent structural shocks, affecting the long-term unemployment rate. Hence, in a “second wave” of research, several econometric techniques were developed to improve the accuracy of tests (Perron, 1989; Zivot and Andrews, 1992 and Lee and Strazicich, 2003), with the “structuralist hypothesis” arising as a substitute for the “hysteresis hypothesis” in modeling European unemployment (e.g., Franchi and Ordóñez, 2008; Lin et al., 2008; Ayala et al., 2012). Table 1 summarises the main results from the recent empirical literature on testing the three competing theories.<sup>1</sup>

Interestingly, the asymmetric behaviour of unemployment only regained attention after Skalin and Teräsvirta (2002)’s empirical application. They argue that unemployment dynamics might be better modeled within a nonlinear framework, which allows not only for different regimes of equilibrium unemployment but

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<sup>1</sup>Some authors argue that these models fail to capture the “genuine” definition of hysteresis. Indeed, these models fail to capture the selective shocks property, but fully account for what we want to assess: the permanent effects of transitory shocks (remanence property) and non-linearity dependence on past shocks (nonlinear property) - e.g., Göcke (2002).

also for different adjustment processes of unemployment over the cycle. Therefore, a “third wave” of research emerged, aiming at validating cycle-related nonlinearities on the dynamics of unemployment, and which may endogenously drive different equilibrium rates (see Table 2).

<b>Authors</b>	<b>Empirical Strategy</b>	<b>Period</b>	<b>Countries</b>	<b>Results</b>
<b>Panel A: Evidence for NAIRU</b>				
Gustavsson and Österholm (2006)	Unit root tests in the nonlinear STAR framework	1948:M1-2005:M1	5 OECD countries	NAIRU hypothesis is supported (except for Australia).
Lee (2010)	Unit root tests in nonlinear heterogeneous panels	1960-2008	29 OECD countries	Natural rate hypothesis found in 23 out of 29 countries.
<b>Panel B: Evidence for Hysteresis</b>				
Logeay and Tober (2006)	Kalman-filter techniques	1970:Q1-2002:Q4	12 Euro Area members	Hysteresis hypothesis is confirmed.
Lee and C.H.Lin (2010)	Panel LM test with breaks	1976-2004	9 East Asian countries	Hysteresis hypothesis found in 7 out of 9 countries.
Cuestas et al. (2011)	Unit root tests allowing for non-linearities, structural breaks, and fractional integration	1998:M1-2007:M12	8 Central and Eastern European countries; 15 EU	Hysteresis hypothesis is supported in most of the countries.
Cheng et al. (2014)	Flexible Fourier unit root test	1960-2011	5 countries (PIIGS)	Hysteresis hypothesis is confirmed for 3 countries.
<b>Panel C: Evidence for Structuralist</b>				
Lee and Chang (2008)	LM unit root tests with endogenous structural breaks	1855-2004	14 OECD countries	In favour of the structuralist hypothesis.
Franchi and Ordóñez (2008)	Unit root test against smooth transition stationarity	1956-2005	5 OECD countries	In favour of the structuralist hypothesis.
Ayala et al. (2012)	Unit root tests allowing for structural breaks, and fractional integration	1980-2009	18 Latin American countries	Structuralist hypothesis is supported for 16 out of 18 Latin American countries.

Table 1: A summary of the recent literature on unemployment hysteresis  
*Notes:* This table is an update of the Table 1 presented in Franchi and Ordóñez (2008), p. 314. For a survey, see Røed (1997).

Theoretically, there are several mechanisms supporting cycle asymmetry. Firstly,

the insider-outsider model suggests that, in expansions, unemployment might not decrease as much as it had increased in recessions due to the ability of the insiders to push wages up, making therefore unprofitable for firms to hire more workers (Blanchard and Summers, 1986; Lindbeck and Snower, 1989). Secondly, asymmetric adjustments may also occur when firing costs are smaller than hiring costs (Bentolila and Bertola, 1990; Hamermesh and Pfann, 1996). Additionally, other labour market institutions (LMI), as unemployment protection and wage bargaining structure, may induce unemployment persistency after a shock (Layard et al., 1991) – labour hoarding.

<b>Authors</b>	<b>Empirical Strat- egy</b>	<b>Period</b>	<b>Countries</b>	<b>Results</b>
Skalin and Teräsvirta (2002)	STAR models	1960:Q1-1997:Q3	11 OECD countries	Linearity is rejected for all except for US, Japan, Norway and Canada.
van Dijk et al. (2002)	FI-STAR model	1968:M1-1999:M12	US	Nonlinearities with persistence behaviour.
Akram (2005)	STAR models	1972:Q2-2004:Q1	Norway	Nonlinear dynamics of the unemployment rate and asymmetrically response to shocks.
Cancelo (2007)	STAR models	1970:Q1-2004:Q4	6 OECD countries	Nonlinear unemployment behaviour captured by the GDP growth rate.
Lin et al. (2008)	Threshold unit root test	1970:M1-2005:M5	16 OECD countries	Nonlinear hysteresis hypothesis: Denmark and Portugal. NAIRU hypothesis in the low regime: Australia, Finland, France, Germany, Japan, USA. Natural rate hypothesis in the high regime: Canada, Ireland and Netherlands.
Franchi and Ordóñez (2011)	STAR models	1972:Q4-2008:Q3	Spain	Nonlinear dynamics of the unemployment rate – asymmetric responses to shock.
Bardsen et al. (2012)	STAR models	1979:Q4-2010:Q2	Australian	Nonlinear dynamics of the unemployment rate – asymmetric responses to shocks.
Cevik and Dibooglu (2013)	MS-ADF test	1948:Q1-2011:Q3	US	Regime-dependent nonstationarity in unemployment: mean-reverting in expansion regime; hysteresis in recession regime.

Table 2: A summary of the recent literature on unemployment nonlinearities across the cycle

*Notes:* For a survey on smooth transition models, see, among others van Dijk et al. (2002).

Taking into account the mixed results within the literature, in this paper we propose a novel step-by-step methodology to study in detail the behaviour of unemployment time series. We divide our analysis into two parts. First, we introduce a set of guidelines to assess if the series follows a unit root process as to confirm the hysteresis hypothesis. In the second part, we develop a baseline nonlinear model to test for the asymmetric behaviour of unemployment across cycle phases. We, then, apply our methodology to the Portuguese case, aiming to: (i) provide, to the best of our knowledge, the first in-depth analysis of the Portuguese unemployment dynamics, testing for the three competing hypotheses; (ii) assess, in particular, if the recent rise in the unemployment rate from, roughly, 8% up to 18% is persistent and results from recession-specific dynamics of unemployment. Indeed, this is an outcome from the 2008 recession combined with the fiscal consolidation strategy to which, alongside with several reforms, Portugal committed under the multilateral financial assistance economic adjustment program (2011-2014). Furthermore, (iii) we try to assess how LMI shape the dynamics of the Portuguese unemployment rate. This Chapter proceeds as follows. Section 2 provides the methodology and the empirical assessment of the “hysteresis hypothesis”. In section 3 we develop a nonlinear model for the Portuguese unemployment dynamics. Section 4 introduces LMI and analyses their impact on the behaviour of the unemployment rate. Finally, Section 5 concludes.

## 2 Hysteresis and unit root tests

This section aims to assess the stationarity of the Portuguese unemployment rate. We first provide a step-by-step methodological guide to a linear modeling of the unemployment rate and, in Section 2.2., we present the estimation results.

### 2.1 Methodology

The study of the stationarity of the series is typically associated with testing if the unemployment rate exhibits a unit root process. As a starting-point, several regular unit root tests are usually applied. The Dickey and Fuller (1979) unit root test uses the following model specification:

$$\Delta u_t = \alpha + \gamma t + \rho u_{t-1} + \sum_{j=1}^K \beta_j \Delta u_{t-j} + \varepsilon_t \quad (1)$$

where  $u_t$  is the unemployment rate,  $\alpha$  is a constant term,  $t$  captures trend,  $K$  is the lag-augmentation for the correction of the residual auto-correction (León-Ledesma and McAdam, 2004), and  $\varepsilon_t \sim iidN(0, \sigma^2)$ . The null hypothesis of a unit root ( $H_0 : \rho = 0$ ) against the alternative of a stationary process ( $H_1 : \rho < 0$ ) can be tested using the conventional  $t$ -ratio for  $\rho$  and the critical values from MacKinnon (1991). The Phillips and Perron (1988), and the Ng and Perron (2001) unit root tests are some other tests applied for the null of a unit root, while Kwiatkowski et al. (1992) LM test is run under the null hypothesis of stationarity.

However, since Perron's (1989) seminal work, it is widely accepted that these standard unit tests can lead to misleading results in the presence of structural breaks. "Perron (1989) argued that if there is a structural break, the power to reject a unit root decreases when the stationarity alternative is true and the structural break is ignored" (Cheng et al., 2014, p.143). Thus, it is desirable to control for the existence of structural breaks in the time series. The test by Bai and Perron (2003) is one of the most widely used to control for structural breaks. If the results support the existence of one or more structural breaks, the next step is to update the previous empirical tests by performing unit root tests with endogenous search for structural breaks (e.g., Perron, 1997; Zivot and Andrews, 1992). For instance, Perron's (1997) test is based on the following regression:

$$\Delta u_t = \alpha + \theta D(< T_b)_t + \gamma t + \delta D(> T_b)_t + \rho u_{t-1} + \sum_{j=1}^K \beta_j \Delta u_{t-j} + \varepsilon_t \quad (2)$$

where  $D(< T_b)_t = 1 (t < T_b)$  and  $D(> T_b)_t = 1 (t = T_b + 1)$  with  $T_b$  being the time at which the change in the trend function occurs.

In both Perron (1997) and Zivot and Andrews (1992) tests, the competitive hypotheses are given by:

$H_0$ : Series has a unit root ( $\rho = 1$ ).

$H_1$ : Series is stationary with one structural break.

The unit root test is performed using the  $t$ -statistic for the null hypothesis. More recently, other unit root tests with endogenous structural breaks have also been proposed in the literature under the argument that the previous tests do not allow for the possibility of structural break under the null hypothesis. In other words, rejecting the null hypothesis does not provide any information regarding the existence of structural break under a non-stationary process. Thus, Lee and

Strazicich (2003) extended the tests by introducing a break under the null hypothesis. On the other hand, Lee and Strazicich (2004) unit root test includes two structural breaks under the null hypothesis. This test can be described as follows:

$$\Delta u_t = \delta' \Delta Z_t + \Phi \tilde{S}_{t-1} + \varsigma_t \quad (3)$$

where  $\tilde{S}_{t-1} = u_t - \tilde{\Psi}_x - Z_t \tilde{S}$ ,  $t = 2, \dots, T$ ;  $\tilde{S}$  are coefficients in the regression of  $\Delta u_t$ ,  $\tilde{\Psi}_x$  is given by  $u_1 - Z_1 \tilde{S}$ , and  $Z_t$  is a vector of exogenous variables. Consider the following data-generating process (DGP):

$$u_t = \delta' Z_t + e_t,$$

$$e_t = \beta e_{t-1} + \varepsilon_t$$

where  $\varepsilon_t \sim iddN(0, \sigma^2)$ . Relying on the crash and break models, proposed by Perron (1989), Lee and Strazicich (2003) introduce the possibility of two shifts in level and in both level and trend, respectively. Thus, for example, the crash model can be described by  $Z_t = [1, t, D_1 t, D_2 t]'$ , where  $D_{jt} = 1$  for  $t \geq T_{Bj} + 1$ ,  $j = 1, 2$ , and 0 otherwise;  $T_{Bj}$  corresponds to the date of the break point. The null hypothesis is described by  $H_0 : \alpha_0 + d_1 B_{1t} + d_2 B_{2t} + u_{t-1} + v_{1t}$  and the alternative hypothesis,  $H_1 : \alpha_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + u_{t-1} + v_{2t}$ , where  $v_{1t}$  and  $v_{2t}$  are stationary error terms;  $B_{jt} = 1$  for  $t = T_{Bj} + 1$ ,  $j = 1, 2$ , and 0 otherwise; and  $d = (d_1, d_2)'$ . Therefore, the unit root hypothesis is  $H_0 : \Phi = 0$ , and the tests statistic are given by  $\tilde{\rho} = T\tilde{\phi}$  and  $\tilde{\tau}$ , the latter corresponding to the t-statistic testing the null hypothesis  $\Phi = 0$ .<sup>2</sup>

From the results obtained from applying the modified unit root tests, two alternative results are feasible. If the performed unit root tests with structural breaks provide evidence for the stationarity of the time series, we might conclude in favour of the structuralist hypothesis. However, if the obtained results support that the series is non-stationary, it is possible to conclude for the hysteresis hypothesis. Independently of the result, one should keep in mind that, in order to apply the standard autoregressive models (AR), the time series must be stationary. In other words, if the empirical evidence favours the hysteresis hypothesis rather than the structuralist, to estimate an AR model one needs to first take into account the non-stationarity property of the series. A common method is to

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<sup>2</sup>For an in-depth analysis of the break model and the unit root test, see Lee and Strazicich (2003).

difference the time series (Hamilton, 1994).<sup>3</sup>

## 2.2 Empirical results

The data used in this empirical research is the quarterly Portuguese unemployment rate, seasonally adjusted, from 1983:Q1 to 2013:Q4.<sup>4</sup> Figure 1 plots the time series. Briefly, it seems that the unemployment rate has been rising uninterruptedly since the beginning of 2001, with a steady and sharp increase since 2008 until 2012.



Mean	Median	Maximum	Minimum	Std. Dev.	Skewness	Kurtosis	Jarque-Bera
7.820	7.200	17.500	3.900	3.116	1.262	4.358	42.464 [0.00]

Figure 1: Quarterly Portuguese unemployment rate, seasonally-adjusted (1983:Q1 – 2013:Q4).

Following the proposed methodology, we applied the relevant unit root tests. Table 3 presents the main results. Panel A reports the standard unit root tests with and without a time trend. Almost all tests fail to reject the hypothesis of stationarity of the series in levels. The only two exceptions are the KPSS and MSB tests with an intercept. Nevertheless, the evidence for the first difference suggests an opposite result, with all tests rejecting the non-stationarity hypothesis. Apparently, hysteresis in unemployment seems to be confirmed for the Portuguese case.

<sup>3</sup>Notice that, if the applied tests regarding the presence of structural breaks do not support the existence of, at least, one structural break, two outcomes are also possible. If the unit root tests favour the stationarity of the series, we should conclude for the NAIRU hypothesis. Otherwise, the hysteresis hypothesis prevails.

<sup>4</sup>Data were gathered from the OECD database at <http://stats.oecd.org/> (accessed on March 2014).



Panel B reports the Zivot and Andrews (1992) and the Perron (1997) unit root tests as a starting point, whereas Panel C presents the results for the tests proposed by Lee and Strazicich (2004, 2003). From the first two tests, we fail to provide evidence to reject the null hypothesis at meaningful significance levels, concluding for the nonstationarity of the series. Moreover, including the possibility of one or two structural breaks does not seem to be enough to make the series stationary as well. Hence, none of the applied tests appear to be able to reject the null hypothesis at 1% significance level.

<b>Panel A: Standard Unit root tests</b>										
with intercept					with intercept and trend					
	Level		First Difference		Level		First Difference			
ADF	-1.195353 (2)		-3.927171***(1)		-2.142080 (2)		-6.226079*** (1)			
PP	-0.248721		-6.025601***		-0.748309		-6.276827***			
KPSS	0.620107**		0.383519 *		0.272075***		0.056098			
MZa	-4.06425 (2)		-15.0218*** (1)		-6.47652 (2)		-18.3134** (1)			
MZt	-1.07799 (2)		-2.52755** (1)		-1.63803 (2)		-2.95154** (1)			
MSB	0.26524* (2)		0.16826*** (1)		0.25292 (2)		0.16117** (1)			
MPT	6.42010 (2)		2.42480** (1)		14.1237 (2)		5.43312** (1)			
<b>Panel B: Unit root tests with endogenous structural breaks</b>										
Zivot and Andrews (1992)					Perron (1997)					
Model	LM stat		$T_B$		LM stat		$T_B$			
Both	-3.6483		1999Q3		-3.6483		1999Q2			
<b>Panel C: Unit root tests with endogenous structural breaks in both <math>H_0</math> and <math>H_1</math></b>										
Model	Lee and Strazicich (2004)				Lee and Strazicich (2003)					
	$\hat{k}$	LM stat	$T_B$	$\lambda$	$\hat{k}$	LM stat	$T_{1B}$	$T_{2B}$	$\lambda_1$	$\lambda_2$
Crash	2	-2.6990	1986Q3	0.121	2	-2.8695	1986Q3	2002Q4	0.121	0.645
Break	2	-3.5430	1998Q3	0.508	2	-4.0553	1991Q4	1999Q2	0.290	0.532

Notes: The results from Bai and Perron (2003) test reported in Appendix A. Panel C:  $T_B$  and  $T_{iB}$  denote the year of structural break for Lee and Strazicich (2004) and Lee and Strazicich (2003), respectively. The critical values for Lee and Strazicich (2004) model and Lee and Strazicich (2003) model are presented in Lee and Strazicich (2004) - Table 1 and Lee and Strazicich (2003) - Table 2, respectively. The lag length has been obtained by following a general-to-specific approach from a maximum of 30 lags. \*, \*\*, and \*\*\* denote test statistic at 10, 5 and 1 percent level, respectively.

Table 3: Applied unit root tests for Portuguese quarterly unemployment rate (1983:Q1-2013:Q4)

Table 4 presents the estimated models based on the results of the LS (2004) and LS (2003) and Figure 2 provides a graphical representation of the estimated models. The structural breaks in the constant and trend are statistically significant in almost all models. Moreover, the years where the structural breaks occur

are in line with the economic and political changes in the sequence of the European Economic Community membership in 1986 and with the beginning of the Euro Zone. Thus, combining the results from Table 3, we might conclude that the Portuguese unemployment rate seems to be a non-stationary process with one or two structural breaks, supporting the hysteresis hypothesis rather than the structuralist or the NAIRU hypothesis. These findings are supported in the literature by Chang et al. (2005) and Lin et al. (2008), but are in contrast with those in Lee (2010) and Cheng et al. (2014).

LS (2004)						
Model	$\alpha$	$\gamma$	$d_1$	$\delta_1$		
Crash	8.24***	-0.08***	-6.01***			
Break	7.87***	-0.04***	-2.50***	0.24***		
LS (2003)						
Model	$\alpha$	$\gamma$	$d_1$	$d_2$	$\delta_1$	$\delta_2$
Crash	8.45***	-0.05***	-5.01***	2.39		
Break	9.75***	-0.16***	-1.40	-3.68***	0.20***	0.17***

Notes: \*, \*\* and \*\*\* indicate test statistic significant at 10%, 5% and 1% percent levels, respectively. The

table presents the results for the following regression models:

Crash model Lee and Strazicich (2003):  $u_t = \alpha + \gamma t + d_1 D_{1t} + d_2 D_{2t} + v_t$ .

Break model Lee and Strazicich (2003):  $u_t = \alpha + \gamma t + d_1 D_{1t} + d_2 D_{2t} + \delta_1 DT_{1t} + \delta_2 DT_{2t} + v_t$

Crash model Lee and Strazicich (2004):  $u_t = \alpha + \gamma t + d_1 D_{1t} + v_t$

Break model Lee and Strazicich (2004):  $u_t = \alpha + \gamma t + d_1 D_{1t} + \delta_1 DT_{1t} + v_t$

Table 4: LS (2004) and LS (2003) unit root tests for the Portuguese quarterly unemployment rate (1983:Q1-2013:Q4)

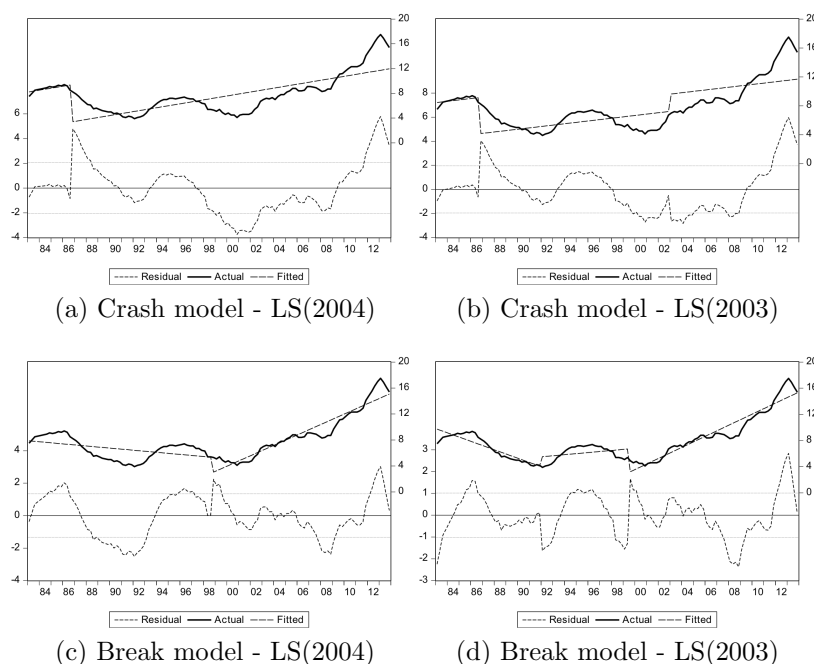


Figure 2: Graphical representation of the LS (2004) and LS (2003) models for the Portuguese quarterly unemployment rate (1983:Q1- 2013:Q4)

Finally, this hysteresis hypothesis also implies that, in order to estimate the standard autoregressive models, we first need to take into account the non-stationarity property. It is worth noting that, according to Table 3, the null hypothesis of a unit root is rejected by all the standard tests applied to first differences, which means that the quarterly change of unemployment rate is stationary. Nevertheless, following van Dijk et al. (2002) and Deschamps (2008), we decided to consider the annual change of unemployment rate rather than its first differences. In theoretical terms, it seems better to consider variations between the same quarters in order to avoid possible problems of seasonal shocks.

Table 5 presents an estimated AR model with five lags. The diagnostic test statistics of the model indicate that there are no significant valuations of the standard assumptions about residuals, with the only exception being the heteroscedasticity test with no white cross terms. The regression specification test (RESET) does not indicate significant functional form misspecification. A word of caution is needed, however. Since this test is constructed to have power against general forms of functional misspecification, it might have low power against specific non-linear forms (Akram, 2005).

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$\Delta_4 U_t =$	-0.165	+ 1.198 $\Delta_4 U_{t-1}$	+ 0.055 $\Delta_4 U_{t-2}$	+ 0.312 $\Delta_4 U_{t-3}$	- 0.472 $\Delta_4 U_{t-4}$	-
	(0.0502)	(0.1023)	(0.1244)	(0.125)	(0.1443)	
		0.401 $\Delta_4 U_{t-5}$	+ 0.218 $D_1$	+ 0.218 $D_2$	+ $\varepsilon_t$	
		(0.0738)	(0.0659)	(0.0293)		

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*Long-run properties:*  $\sum \alpha_i = 0.868954$ ; Diagnostic tests: Log-Likelihood value: -45.72; Standard error of residuals:  $\hat{\sigma} = 0.37$ ; Autocorrelation 1-4:  $X^2(4) = 5.12[0.27]$ ; ARCH 4:  $X^2(4) = 1.80[0.77]$ ; Normality:  $X^2(2) = 4.28[0.12]$ ; Heteroscedasticity  $F_{xixj}$ :  $X^2(32) = 41.54[0.12]$ ; Heteroscedasticity  $F_{x_i^2}$ :  $X^2(7) = 21.00[0.00]$ ; RESET test:  $F(1, 106) = 0.35[0.56]$

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Notes: The lag order was chosen based on the Akaike's information criterion Granger and Teräsvirta (1993). Since there is evidence for the existence of one or two structural break, we introduce one or two dummies in the AR model. To choose the best one, we fit four different AR model – two with one structural break, corresponding to the results from LS (2004), and two with two structural breaks from LS (2003). Comparing the standard criterions among them – AIC, SC, Log Likelihood – we concluded that the model with two structural breaks on 1989:Q2 and 2011:Q1 is preferable. The standard errors are in parentheses below the estimates and *p-values* are shown in square brackets. Autocorrelation 1-4 tests for residuals up to 4 lags; ARCH 4 tests for autoregressive conditional heteroskedasticity up to order 4 (Engle, 1982). The normality test is the Jarque and Bera (1980).  $F_{xixj}$  and  $F_{x_i^2}$  are tests for residuals heteroskedasticity due to omission of cross products of regressors and/or square regressors (White, 1980). RESET is the standard regression specification test (Ramsey, 1969).

Table 5: An AR (5) model for the annual changes in Portuguese quarterly unemployment rates (1983:Q1-2013:Q4)

### 3 A nonlinear benchmark model

The previous applied unit root tests implicitly assumed that the time series is well described by a linear behaviour. In other words, they might lack significance if nonlinearities are present. On the one hand, as we stated in the introduction, there are several theoretical reasons for an asymmetric behaviour of unemployment across the cycle. On the other hand, this asymmetric behaviour can also be empirically observed, through sharper increases (smoother reductions) in the unemployment rate during recessions (expansions). This allows for the possibility of multiple equilibria, endogenously driven by cycle (transitory) conditions. Thus, the next step is to test the AR(q) model against a possible nonlinear model, such as the smooth-transition autoregressive (STAR) model.

#### 3.1 Methodology

One of the most well-known nonlinear methods was proposed by Granger and Teräsvirta (1993) and Teräsvirta (1994), which can be described as follows. In a univariate framework, a STAR model of unemployment can be formulated as follows:

$$\Delta u_t = \alpha + \beta u_{t-1} + \sum_{i=1}^q \phi_i \Delta u_{t-i} + \left( \tilde{\alpha} + \tilde{\beta} u_{t-1} + \sum_{i=1}^q \tilde{\phi}_i \Delta u_{t-i} \right) F(\gamma, \Delta u_{t-d} - c) + \varepsilon_t \quad (4)$$

where  $\alpha, \beta, \phi_i, \gamma$  and  $c$  are parameters to be estimated and  $\varepsilon_t \sim iidN(0, \sigma^2)$ . The transition function  $F(\gamma, \Delta u_{t-d} - c)$  is continuous, non-decreasing, takes values in the 0–1 range, and works as a proxy for the cycle phase of the economy. Franchi and Ordóñez, 2011, p.72, states that “[t]he STAR model can be interpreted as a regime switching model that allows for two regimes, associated with the extremes values  $F(\gamma, \Delta u_{t-d} - c) = 0$  and  $F(\gamma, \Delta u_{t-d} - c) = 1$ , each corresponding to a specific state of the economy” (). The transition between regimes occurs when the transition variable ( $\Delta u_{t-d}$ ) deviates from a constant threshold (steady-state), value  $c$ , and its speed is governed by the parameter  $\gamma$ .

Following Teräsvirta, 1994, p.210, the specification procedure can be viewed as a sequence consisting of three steps: (i) specify a linear autoregressive model; (ii) test linearity for different values of  $d$ , the delay parameter and, if it is rejected, select the appropriate transition function; (iii) choose between the logistic smooth transition autoregressive model (LSTAR) and the exponential smooth transition autoregressive model (ESTAR), by testing a sequence of nested hypothesis.

To the LSTAR model corresponds the logistic function,

$$F(\gamma, \Delta u_{t-d} - c) = (1 + \exp[-\gamma \{\Delta u_{t-d} - c\}])^{-1}, \gamma > 0 \quad (5)$$

and the ESTAR model corresponds the exponential function,

$$F(\gamma, \Delta u_{t-d} - c) = 1 - \exp\{-\gamma (\Delta u_{t-d} - c)^2\}, \gamma > 0 \quad (6)$$

Notice that the LSTAR model reduces to a self-exciting threshold autoregressive (SETAR) model when  $\gamma \rightarrow \infty$  and, therefore, the logistic function approaches 1. Conversely, when  $\gamma = 0$ , the LSTAR model is reduced to an AR model. The different response to positive and negative deviations of  $\Delta u_{t-d}$  from  $c$  makes the LSTAR model convenient for modeling unemployment when asymmetric behaviour arises (Franchi and Ordóñez, 2011).

As pointed out by Teräsvirta (1994), testing linearity against STAR is rather complicated – under the null hypothesis, the parameters defining the STAR model are not identified. Hence, Teräsvirta (1994) proposes a sequence of tests to evaluate the null of an  $AR(q)$  model against the alternative STAR model. These tests are based on estimating the following auxiliary regression for a chosen set

of values of the delay parameter,  $d$ :

$$\Delta u_t = \beta_0 + \sum_{i=1}^q \beta_{1i} \Delta u_{t-i} + \sum_{i=1}^q \beta_{2i} \Delta u_{t-i} \Delta u_{t-d} + \sum_{i=1}^q \beta_{3i} \Delta u_{t-i} \Delta u_{t-d}^2 + \sum_{i=1}^q \beta_{4i} \Delta u_{t-i} \Delta u_{t-d}^3 + v_t \quad (7)$$

where  $v_t$  is the error term. This auxiliary regression was adapted taking into consideration our proposed STAR model in (4). Testing an  $AR(q)$  model against a STAR model is equivalent to:

$$H_0 : \beta_{2i} = \beta_{3i} = \beta_{4i} = 0, i = 1, 2, \dots, q. \quad (8)$$

In order to identify the appropriate lag (value of  $d$ ) to be used in the transition variable, this test should be conducted for different values of  $d$  in the range  $1 \leq d \leq q$ . If linearity is rejected, the next step is to test for LSTAR against ESTAR model. The following sequence of tests on the auxiliary regression was proposed by Granger and Teräsvirta (1993), and Teräsvirta (1994):

$$H_{04}: \beta_{4i} = 0, i = 1, 2, \dots, q.$$

$$H_{03}: \beta_{3i} = 0 | \beta_{4i} = 0, i = 1, 2, \dots, q.$$

$$H_{02}: \beta_{2i} = 0 | \beta_{3i} = \beta_{4i} = 0, i = 1, 2, \dots, q.$$

An LSTAR model should be selected if  $H_{04}$  or  $H_{02}$  is rejected for at least one value of  $i$  and an ESTAR model if  $H_{03}$  is rejected for at least one  $i$ .

## 3.2 Empirical results

In our case, since we have already estimated an  $AR(5)$  model in Section 2.2, above, we can now test linearity against STAR models, following the auxiliary regression described by (7). We apply the method using three transition variables: (a)  $\Delta u_{t-d}$ , the variation of unemployment itself, lagged  $d$  periods; (b)  $\Delta \log GDP_{t-d}$ , the logarithmic variation of GDP, denoting, approximately, the GDP growth rate; and (c)  $\Delta u_{t-d}^c$ , the variation of the cyclical unemployment, lagged  $d$  periods.

	$s_t = \Delta u_{t-d}$		$s_t = \Delta u_{t-d}^c$		$s_t = \Delta \log GDP_{t-d}$	
$D$	Testing Linearity	$p$ -value	Testing Linearity	$p$ -value	Testing Linearity	$p$ -value
1	$X_{24}^2 = 170.25$	[0.000]	$X_{24}^2 = 349.23$	[0.000]	$X_{24}^2 = 202.86$	[0.000]
2	$X_{24}^2 = 179.79$	[0.000]	$X_{24}^2 = 162.53$	[0.000]	$X_{24}^2 = 283.99$	[0.000]
3	$X_{24}^2 = 294.13$	[0.000]	$X_{24}^2 = 732.01$	[0.000]	$X_{24}^2 = 820.95$	[0.000]
4	$X_{24}^2 = 294.13$	[0.000]	$X_{24}^2 = 400.05$	[0.000]	$X_{24}^2 = 232.43$	[0.000]
5	$X_{24}^2 = 405.62$	[0.000]	$X_{24}^2 = 388.75$	[0.000]	$X_{24}^2 = 136.83$	[0.000]

(a) Panel A: LM test for STAR nonlinearities

	$H_{04}$	$p$ -value	$H_{03}$	$p$ -value	$H_{02}$	$p$ -value
<b>Panel B1</b>						
$\Delta u_{t-1}$	$X_8^2 = 25.84$	[0.001]	$X_8^2 = 13.58$	[0.093]	$X_8^2 = 5.62$	[0.690]
$\Delta u_{t-2}$	$X_8^2 = 11.81$	[0.160]	$X_8^2 = 56.86$	[0.000]	$X_8^2 = 13.24$	[0.104]
$\Delta u_{t-3}$	$X_8^2 = 17.63$	[0.024]	$X_8^2 = 12.10$	[0.147]	$X_8^2 = 24.90$	[0.002]
$\Delta u_{t-4}$	$X_8^2 = 4.46$	[0.814]	$X_8^2 = 22.76$	[0.004]	$X_8^2 = 35.96$	[0.000]
$\Delta u_{t-5}$	$X_8^2 = 14.61$	[0.067]	$X_8^2 = 54.52$	[0.000]	$X_8^2 = 24.12$	[0.002]
<b>Panel B2</b>						
$\Delta u_{t-1}^c$	$X_8^2 = 6.00$	[0.648]	$X_8^2 = 17.30$	[0.027]	$X_8^2 = 20.60$	[0.008]
$\Delta u_{t-2}^c$	$X_8^2 = 21.24$	[0.007]	$X_8^2 = 24.47$	[0.002]	$X_8^2 = 24.73$	[0.002]
$\Delta u_{t-3}^c$	$X_8^2 = 50.84$	[0.000]	$X_8^2 = 20.92$	[0.007]	$X_8^2 = 50.45$	[0.000]
$\Delta u_{t-4}^c$	$X_8^2 = 8.71$	[0.367]	$X_8^2 = 20.01$	[0.010]	$X_8^2 = 52.19$	[0.000]
$\Delta u_{t-5}^c$	$X_8^2 = 44.76$	[0.000]	$X_8^2 = 44.42$	[0.000]	$X_8^2 = 36.58$	[0.000]
<b>Panel B3</b>						
$\Delta \log GDP_{t-1}$	$X_8^2 = 25.25$	[0.001]	$X_8^2 = 30.00$	[0.000]	$X_8^2 = 42.68$	[0.000]
$\Delta \log GDP_{t-2}$	$X_8^2 = 9.43$	[0.307]	$X_8^2 = 71.64$	[0.000]	$X_8^2 = 33.57$	[0.000]
$\Delta \log GDP_{t-3}$	$X_8^2 = 52.65$	[0.000]	$X_8^2 = 21.93$	[0.005]	$X_8^2 = 39.22$	[0.000]
$\Delta \log GDP_{t-4}$	$X_8^2 = 14.35$	[0.0731]	$X_8^2 = 12.23$	[0.1413]	$X_8^2 = 23.02$	[0.003]
$\Delta \log GDP_{t-5}$	$X_8^2 = 18.55$	[0.0175]	$X_8^2 = 6.43$	[0.5990]	$X_8^2 = 44.13$	[0.000]

(b) Panel B: STAR model selection

Table 6: LM test for STAR nonlinearities for the annual changes in Portuguese quarterly unemployment rates (1983:Q1-2013:Q4)

Table 6: Panel A presents the test statistics for the null hypothesis. Following Granger and Teräsvirta (1993) approach, we conclude that an AR(5) model is rejected at a 1% significance level for any of the proposed transition variables, for  $d = 1$  to  $d = 5$ . We conjecture that a nonlinear specification seems more appropriate to characterise the Portuguese unemployment process.

Table 6: Panel B shows the results for the second part of the procedure, regarding the choice between LSTAR and ESTAR model specification. The results are straightforward: with only a few exceptions, we can estimate a LSTAR or an ESTAR model with any of the suggested transition variables. Therefore, the next step is to estimate the nonlinear model. Following Teräsvirta (1994); van Dijk et al. (2002); Akram (2005); Camarero et al. (2006); Lin et al. (2008);

Deschamps (2008); Franchi and Ordóñez (2011) and Bardsen et al. (2012), we estimate the nonlinear model with a logistic transition function rather than with an exponential transition function because it seems to better capture the dynamics of unemployment. In order to avoid misspecification of the model, we follow a “general-to-specific” approach: we start by consider all the five possible lags and analyse its relevance to the model by checking the  $p$ -values.

$$\Delta_4 U_t = \begin{matrix} 0.226 & - & 0.256U_{t-1} & + & 0.274\Delta_4 U_{t-1} & - & 0.208\Delta_4 U_{t-3} & - \\ (0.148) & & (0.234) & & (0.926) & & (0.117) & \end{matrix}$$

$$0.442\Delta_4 U_{t-4} + 0.396\Delta U_{t-5} + F(G) \left[ \begin{matrix} 0.231U_{t-1} & + & 0.898\Delta_4 U_{t-1} \\ (0.237) & & (0.914) \end{matrix} \right] + \varepsilon_t$$

$$\text{where } F(G) = \left( 1 + \exp \left[ \frac{-5.894 \left( \begin{matrix} \Delta_4 U_{t-2} - (-1.075) \\ (6.405) & (0.394) \end{matrix} \right)}{\hat{\sigma}(\Delta_4 U_{t-2})} \right] \right)^{-1}$$


---

*Long-run properties:*  $F(G) = 0 : \sum \phi_i = 0.02$ ;  $F(G) = 1 : \sum \phi_i = 0.918$ ;  $\hat{u} = 9.04$ . *Diagnostic tests:*  
*AIC:* -2.09; *SBIC:* -1.84; *Standard error of residuals,*  $\hat{\sigma} = 0.33$ ; *Samples Standard deviation of*  
 $\Delta_4 U_{t-2} : \hat{\sigma}(\Delta_4 U_{t-2}) = 0.975$ ; *Autocorrelation 1-4:*  $F(4, 89) = 0.130[0.27]$ ; *ARCH 4:*  
 $\chi^2(4) = 1.56[0.82]$ ; *Normality:*  $\chi^2(2) = 6.19[0.05]$ ; *Heteroscedasticity*  $F_{xixj} : \chi^2(49) = 33.72[0.38]$ ;  
*RESET test:*  $F(1, 92) = 0.42[0.52]$ ;

---

Table 7: Parsimonious LSTAR model with annual change of unemployment as transition variable

Table 7 proposes a parsimonious LSTAR model with annual change of unemployment as transition variable and its properties. Indeed, two main results arise. First, there is cycle asymmetry, since the sum in absolute value of the autoregressive coefficients is lower(higher) when changes in unemployment are below(above) the threshold ( $\Delta_4 U_{t-2} = -1.075$ ). This suggests that unemployment rises faster than it decreases, which is in line with the theoretical framework and with the available data. Second, there is unemployment persistence in one of the regimes ( $F(G) = 1 : \sum \phi = 0.918$ ). Finally, the model also predicts an equilibrium unemployment rate of 9%. The diagnostic tests do not indicate any misspecification problems regarding autocorrelation, heteroscedasticity nor normality.

Figure 3 presents the transition function and plots the residuals from the linear model used as a departing point for linearity testing together with the residuals from the nonlinear model. Since the value of  $\tilde{\gamma}$  is large, the transition between one regime to another is rather fast. On the other hand, it is interesting to note that from the beginning of the century the transition function is systematically close to 1, implying, therefore, that unemployment rate has been rising and moving



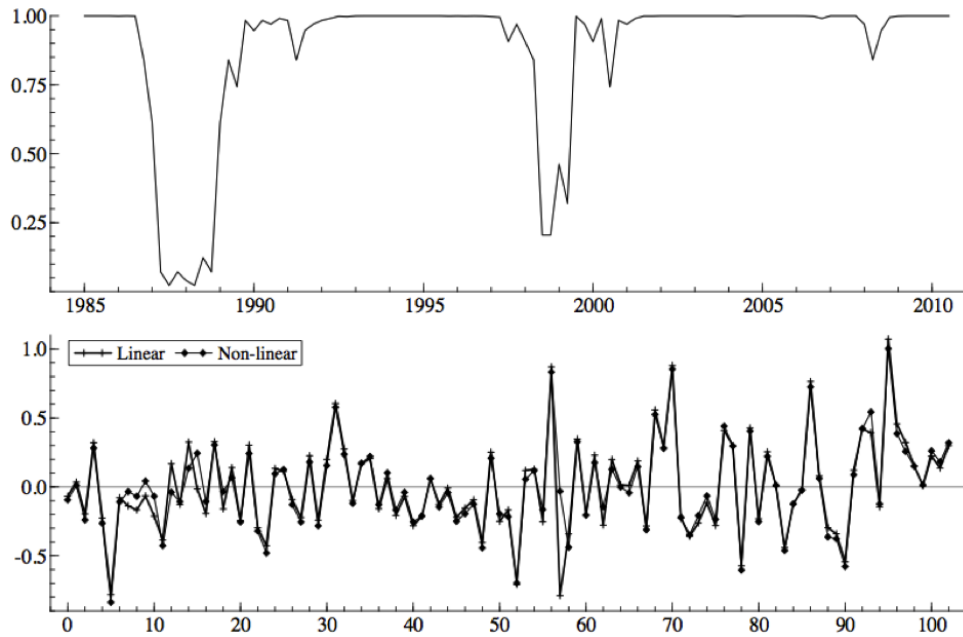


Figure 3: Parsimonious LSTAR model with annual change of unemployment as transition variable

towards a higher value ever since. Finally, plotting the residuals indicates that the major contribution of the nonlinear model is where the unemployment rate is decreasing, which supports the asymmetry property.

Table 8 proposes a parsimonious LSTAR model with annual GDP growth rate as transition variable. Once again, the cycle asymmetry property arises, since the sum in absolute value of the autoregressive coefficients is lower (higher) when changes in unemployment are below (above) the threshold ( $\Delta_4 \log GDP_{t-4} = 4.462$ ). This suggests that unemployment rises faster than it decreases, which is in line with the previous model. Unemployment persistence is also present in one of the regimes ( $F(G) = 0 : \sum \phi = 0.973$ ). Nevertheless, comparing the two models, notice that the latter has a higher persistence in both regimes, which might be explained by the labour hoarding phenomenon: a situation when, for example, firms tend to employ more workers than they need in recessions in order to guarantee that their human capital will be available in expansions. According to Fiorito and Kollintzas (1994, p.258), “labour hoarding (...) is a situation where firms find relatively more costly to adjust employment rather than hours per worker, so that they have an incentive to smooth employment over the business cycle and utilize labor more intensively in expansions and less intensively in contractions”. Moreover, the authors shown that this phenomenon has a particularly impact in Europe and Japan. Finally, since the transition variable is the annual GDP growth rate and not the annual change of unemployment, it

is possible to identify two unemployment regimes. Indeed, the estimated model indicates a low regime with an equilibrium value of 4.49% and a high regime of 7.11%. The former regime corresponds to  $F(G) = 1$  and the latter to  $F(G) = 0$ . The diagnostic tests do not indicate any misspecification problems.

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$$\Delta_4 U_t = \begin{matrix} 0.128 & - & 0.018U_{t-1} & + & 1.202\Delta_4 U_{t-1} & - & 0.218\Delta_4 U_{t-3} & - & 0.442\Delta U_{t-4} & + \\ (0.157) & & (0.023) & & (0.067) & & (0.118) & & (0.133) & \end{matrix}$$

$$0.431\Delta_4 U_{t-5} + F(G) \left[ \begin{matrix} 1.979 & - & 0.451U_{t-1} & - & 0.549\Delta_4 U_{t-1} \\ (0.859) & & (0.196) & & (0.262) \end{matrix} \right] + \varepsilon_t$$

$$\text{where } F(G) = \left( 1 + \exp \left[ \frac{-18.74 \left( \frac{\Delta_4 \log GDP_{t-4} - 4.462}{\hat{\sigma}(\Delta_4 \log GDP_{t-4})} \right) + 0.182}{(12.16)} \right] \right)^{-1}$$


---

*Long-run properties:*  $F(G) = 0 : \sum \phi_i = 0.973; \hat{u}_1 = 7.11; F(G) = 1 : \sum \phi_i = 0.424; \hat{u}_2 = 4.49.$   
*Diagnostic tests:*  $AIC: -2.10; SBIC: -1.81; Standard error of residuals, \hat{\sigma} = 0.33; Samples$   
*Standard deviation of  $\Delta_4 \log GDP_{t-4} : \hat{\sigma}(\Delta_4 \log GDP_{t-4}) = 2.616; Autocorrelation 1-4:$   
 $F(4, 88) = 1.22[0.31]; ARCH 4: \chi^2(4) = 0.87[0.93]; Normality: \chi^2(2) = 7.35[0.03];$   
*Heteroscedasticity  $F_{xixj} : \chi^2(39) = 37.62[0.53]; RESET test: F(1, 91) = 0.09[0.77];$**

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Table 8: Parsimonious LSTAR model with annual GDP growth rate as transition variable

Figure 4 presents the transition function and plots the residuals from the linear model used as a basis for linearity testing together with the residuals from the non-linear model. Since the value of  $\tilde{\gamma}$  is also large, the transition between one regime to another is rather fast. On the other hand, it is interesting to note that from 2001 onwards the transition function is systematically close to 0, implying therefore that unemployment rate has been “stuck” in its high regime. These results are in line with the previous model, which corroborates the consistency of the methodology and the models. Finally, plotting the residuals indicates that there are gains in using the nonlinear model when unemployment is increasing and decreasing.

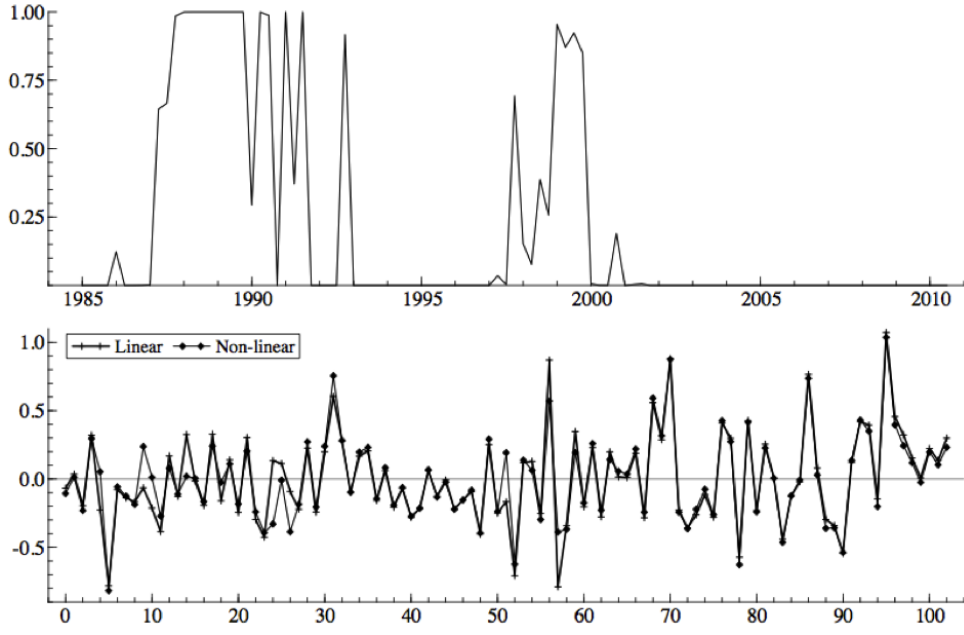


Figure 4: Parsimonious LSTAR model with annual GDP growth rate as transition variable

$$\Delta_4 U_t = \frac{1.037}{(0.255)} - \frac{0.046 U_{t-1}}{(0.021)} + \frac{1.569 \Delta_4 U_{t-1}}{(0.154)} - \frac{0.549 \Delta U_{t-4}}{(0.098)} + \frac{0.3483 \Delta_4 U_{t-5}}{(0.095)} + F(G) \begin{bmatrix} -0.8708 & -0.315 \Delta_4 U_{t-1} \\ (0.241) & (0.168) \end{bmatrix} + \varepsilon_t$$

$$\text{where } F(G) = \left( 1 + \exp \left[ \frac{-3.871 \left( \frac{\Delta_4 U_{t-1} - (-0.5121)}{\hat{\sigma}(\Delta_4 U_{t-1})} \right)}{(3.157) \quad 0.112} \right] \right)^{-1}$$

*Long-run properties:*  $F(G) = 0 : \sum \phi_i = 1.368$ .  $F(G) = 1 : \sum \phi_i = 1.053$ ;  $\hat{u} = 3.61$ . *Diagnostic tests:*  $AIC: -2.24$ ;  $SBIC: -2.01$ ; *Standard error of residuals,*  $\hat{\sigma} = 0.313$ ; *Samples Standard deviation of*  $\Delta_4 U_{t-1} : \hat{\sigma}(\Delta_4 U_{t-1}) = 0.6505$ ; *Autocorrelation 1-4:*  $F(4, 90) = 1.81[0.13]$ ; *ARCH 4:*  $\chi^2(4) = 5.09[0.28]$ ; *Normality:*  $\chi^2(2) = 7.35[0.47]$ ; *Heteroscedasticity*  $F_{x_i x_j} : \chi^2(39) = 26.01[0.41]$ ; *RESET test:*  $F(1, 93) = 1.00[0.32]$ ;

Table 9: Parsimonious LSTAR model with annual change of cyclical unemployment as transition variable

Finally, Table 9 presents a parsimonious LSTAR model with annual change of cyclical unemployment as transition variable. In this case, however, the interpretation is not straightforward. First, both regimes seem to be nonstationary, since the sum of the regressive coefficients is above one. Moreover, although the cycle asymmetry is present in the model, it is not in line with the data and with the theoretical framework. Indeed, this model predicts that unemployment

decreases faster than it increases. Notice that  $F(G) = 0 : \sum \phi_i = 1.368$  and  $F(G) = 1 : \sum \phi = 1.053$ . Nevertheless, the diagnostic tests do not indicate any misspecification problems. Taking into account these results, we argue that one of the possible reasons to explain this behaviour relates with the fact that over the past years the cyclical unemployment fails to account for a substantial part of the unemployment dynamics itself, as Figure 5 illustrates.

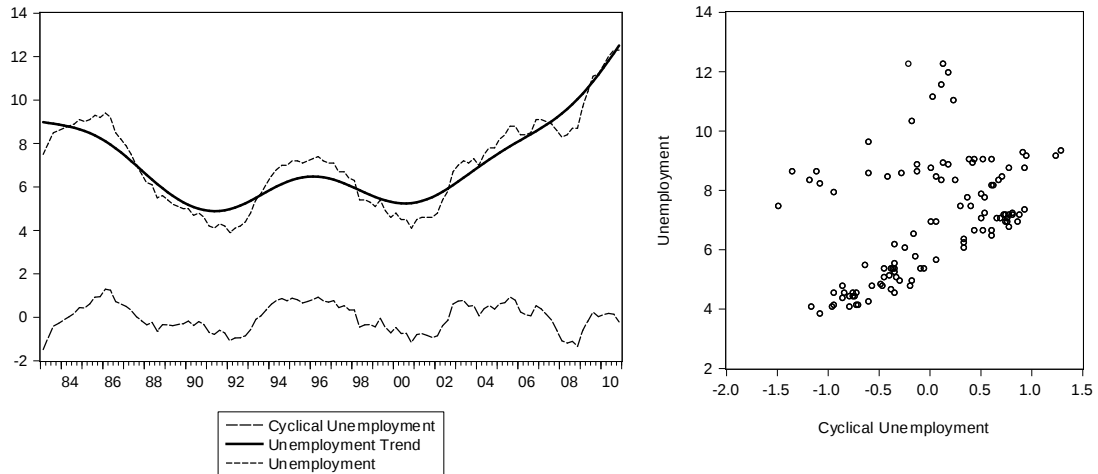


Figure 5: Unemployment dynamics

Broadly speaking, most of the rising in unemployment from the beginning of the century relates with its trend rather than with its cyclical component. If this pattern is confirmed, it would be more difficult than expected to revert this process, in the sense that some of the current unemployed workers might never re-enter into the labour market.

## 4 The impact of LMI - a tentative assessment

The institutional framework of the labour market seems to crucially influence the impact of “shocks” in an economy, as well as its adjustment towards the equilibrium. Theoretically speaking, Flaig and Rottmann (2013, p.637) states “the location and shape of both the price setting and wage setting functions depend on many institutional settings. (...) Consequently, changes in labour market institutions lead to a shift in one or both functions and to a change in the equilibrium values of the real wages and the unemployment rate”. Moreover, Nickell et al. (2005, p. 3) argues that “shocks drive unemployment but the scale of the unemployment consequences of any particular shocks depend on the institutional

framework of the economy”. Furthermore, a different set of institutions seems also to play an important role in explaining differences in the unemployment level among countries, specifically within the European Union. Following Belot and van Ours (2004, p.621), “the search for relationships between unemployment and labor market institutions is motivated by the fact that across countries there are substantial differences in the level and evolution of unemployment”.

<b>Av_Nrr<sup>a</sup></b>	Average net replacement rates for single earners
<b>Coord<sup>b</sup></b>	Coordination of wage setting
<b>EPT<sup>c</sup></b>	Strictness of employment protection - temporary employment.
<b>Level<sup>b</sup></b>	The predominant level(s) at which wage bargaining takes place
<b>UbDur<sup>a</sup></b>	Unemployment benefit duration
<b>UD<sup>b</sup></b>	Union density rate, net union membership as a proportion of wage and salary earners in employment

Table 10: Description of the LMI

Sources:

a) Center for Economic Studies (CES) - CESifo on-line database - table “Unemployment Benefit Schemes”, available at <https://www.cesifo-group.de/ifoHome.html>.

b) Amsterdam Institute for Advanced labour studies (AIAS) on-line database - table “ICTWSS Database 4, April 2013”, available at <http://www.uva-aias.net/208>.

c) OECD on-line database - table “Employment Protection”, available at <http://stats.oecd.org>.

	<b>Av. Nrr</b>	<b>Coord</b>	<b>EPT</b>	<b>Level</b>	<b>Ubdur</b>	<b>UD</b>
<b>1983-85</b>	14.42	2.00	3.38	3.00	0.35	46.36
<b>1986-89</b>	29.94	3.00	3.38	3.50	0.49	36.71
<b>1990-94</b>	34.61	3.00	3.38	3.80	0.53	26.85
<b>1995-99</b>	37.81	3.00	3.04	3.80	0.58	24.38
<b>2000-04</b>	42.91	2.20	2.76	3.00	0.60	21.46
<b>2005-10</b>	51.79	2.33	2.25	3.00	0.63	20.48
<b>Mean</b>	37.52	2.61	2.93	3.36	0.55	27.58
<b>Median</b>	35.39	3.00	2.81	3.00	0.54	25.29
<b>Max</b>	55.19	3.00	3.38	5.00	0.68	47.24
<b>Min</b>	7.18	2.00	1.94	3.00	0.33	19.34
<b>Std. Dev</b>	11.43	0.50	0.49	0.78	0.09	8.62

Table 11: Summary of the LMI

Therefore, in this section we aim to understand how LMI might affect (i) the transition dynamics from one regime to another; (ii) the unemployment dynamics within each regime; and (iii) the actual equilibrium unemployment level suggested by the previous models. Table 5.10 provides a brief description of selected LMI variables and Table 5.11 presents a statistical summary.

Taking into account the lack of theoretical support from the model present in Table 9, we build our analysis of the LMI upon the first two models: a) the parsimonious LSTAR model with annual change of unemployment as transition variable; and b) the parsimonious LSTAR model with annual GDP growth rate as transition variable. We follow a general-to-specific approach. We tested and analysed all the considered LMI in Table 10. Nevertheless, since most LMI variables are high correlated (see Appendix B), in the end only three models remained, each one with only one LMI, either EPT or Level.<sup>5</sup>

Table 5.12 proposes a parsimonious LSTAR model with annual change of unemployment as transition variable and with *EPT* as explanatory variable.

$$\Delta_4 U_t = 1.386 - 0.277U_{t-1} + 0.280\Delta U_{t-1} - 0.201\Delta_4 U_{t-3} - 0.445\Delta_4 U_{t-4} + 0.424\Delta U_{t-5} - 0.293EPT + F(G) \begin{bmatrix} 0.209U_{t-1} + 0.873\Delta_4 U_{t-1} \\ (0.125) \quad (0.502) \end{bmatrix} + \varepsilon_t$$

where  $F(G) = \left( 1 + \exp \left[ \frac{-4.985 \left( \frac{\Delta_4 U_{t-2} - (-0.977)}{\hat{\sigma}(\Delta_4 U_{t-2})} \right)}{(2.561) \quad (0.301)} \right] \right)^{-1}$

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*Long-run properties:*  $F(G) = 0 : \sum \phi_i = 0.06$ ;  $F(G) = 1 : \sum \phi_i = 0.931$ . *Diagnostic tests:* *AIC:* -2.19; *SBIC:* -1.92; *Standard error of residuals,*  $\hat{\sigma} = 0.317$ ; *Samples Standard deviation of*  $\Delta_4 U_{t-2} : \hat{\sigma}(\Delta_4 U_{t-2}) = 0.975$ ; *Autocorrelation 1-4:*  $F(4, 88) = 0.96[0.44]$ ; *ARCH 4:*  $\chi^2(4) = 1.75[0.78]$ ; *Normality:*  $\chi^2(2) = 1.03[0.60]$ ; *Heteroscedasticity*  $F_{xixj} : \chi^2(41) = 54.14[0.08]$ ; *RESET test:*  $F(1, 91) = 0.16[0.69]$ ;

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Table 12: Parsimonious LSTAR model with annual change of unemployment as transition variable

The cycle asymmetry property arises again, since the sum in absolute value of the autoregressive coefficients is lower (higher) when changes in unemployment are below (above) the threshold. Nevertheless, when compared with the baseline model (Table 7), it seems that both regimes are now more persistent. Moreover, the transition between regimes seems now slower ( $\hat{\gamma} = \frac{4.985}{0.975}$ ), which is in line with the theoretical framework in which LMI influence the adjustment to shocks. Interestingly, the sign of *EPT* is negative - this means that higher employment protection regarding temporary employment contributes to a lower equilibrium unemployment level. The equilibrium unemployment rate suggested by the model can be calculated as follows:

<sup>5</sup>Notice that, due to the lack of available information regarding LMI, the new estimated models gather data from 1983:Q1 to 2010:Q4.

$$\hat{u} = \frac{-1.386 + 0.293EPT}{-0.277 + 0.209}$$

Notice that the level of  $EPT$  directly contributes to the equilibrium unemployment rate. In what relates to the Portuguese case, the minimum and maximum values for  $EPT$  are 1.938 and 3.375, implying therefore that:

$$\begin{cases} \hat{u}_{Max} = 12.03 \\ \hat{u}_{Min} = 5.84 \end{cases}$$

It is curious to note that the equilibrium value suggested by the baseline model is between these two. Notice also that, since we are using annual change of unemployment as transition variable, it is not possible to argue for the existence of two equilibrium unemployment rates as in the model with annual GDP growth rate as transition variable. Indeed, what this model presents is the possibility of changes in the equilibrium unemployment rate, depending on the value of  $EPT$ . Finally, the diagnostic tests do not indicate any misspecification problems.

Table 13 introduces a parsimonious LSTAR model with annual change of unemployment as transition variable and with  $Level$  as explanatory variable.

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$$\Delta_4 U_t = \begin{matrix} 0.520 & - & 0.286U_{t-1} & + & 0.147\Delta U_{t-1} & - & 0.209\Delta_4 U_{t-3} & - & 0.446\Delta_4 U_{t-4} & + \\ (0.224) & & (0.146) & & (0.583) & & (0.115) & & (0.131) \end{matrix} +$$

$$0.390\Delta U_{t-5} - 0.079LEVEL + F(G) \begin{bmatrix} 0.258U_{t-1} + 1.01\Delta_4 U_{t-1} \\ (0.145) \quad (0.586) \end{bmatrix} + \varepsilon_t$$

$$\text{where } F(G) = \left( 1 + \exp \left[ \frac{-5.982 \begin{pmatrix} \Delta_4 U_{t-2} - (-1.06) \\ (3.373) \quad (0.252) \end{pmatrix}}{\hat{\sigma}(\Delta_4 U_{t-2})} \right] \right)^{-1}$$


---

*Long-run properties:*  $F(G) = 0 : \sum \phi_i = -0.118$ ;  $F(G) = 1 : \sum \phi_i = 0.895$ . *Diagnostic tests:*  $AIC: -2.10$ ;  $SBIC: -1.82$ ; *Standard error of residuals,*  $\hat{\sigma} = 0.332$ ; *Samples Standard deviation of*  $\Delta_4 U_{t-2} : \hat{\sigma}(\Delta_4 U_{t-2}) = 0.975$ ; *Autocorrelation 1-4:*  $F(4, 88) = 0.99[0.42]$ ;  $ARCH\ 4:$   $\chi^2(4) = 1.56[0.82]$ ; *Normality:*  $\chi^2(2) = 4.35[0.11]$ ; *Heteroscedasticity*  $F_{xixj} : \chi^2(40) = 38.75[0.53]$ ; *RESET test:*  $F(1, 91) = 0.15[0.70]$ ;

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Table 13: Parsimonious LSTAR model with annual change of unemployment as transition variable

Once again, the cycle asymmetry property arises, since the sum in absolute value of the autoregressive coefficients is lower(higher) when changes in unemployment are below(above) the threshold. However, in this case, when  $F(G) = 0$ , the sum is negative but higher than -1. This means that, although this regime is stationary, the adjustment to shocks is not smooth but with small “jumps” from

positive to negative values until it converges. This rather strange adjustment is, nevertheless, also present in the literature - see Skalin and Teräsvirta (2002). In any case, since all the remaining properties seem to be in line with the data and with the theoretical framework, we continue to analyse the model. The transition between regimes seems now faster ( $\hat{\gamma} = \frac{5.982}{0.975}$ ), and the sign of *Level* - the predominant level(s) at which wage bargaining takes place - is also negative, which is in line with the previous model.

Once again, the value of *Level* contributes to the equilibrium unemployment rate. In what relates to the Portuguese case, the minimum and maximum values for *Level* are 3 and 4, implying therefore that:

$$\begin{cases} \hat{u}_{Max} = 10.11 \\ \hat{u}_{Min} = 7.29 \end{cases}$$

The equilibrium value suggested by the baseline model is, again, between these two. The interpretation is rather similar to the previous model, which implies that the equilibrium unemployment level seems to depend on the value of *EPT*. Finally, the diagnostic tests do not indicate any misspecification problems.

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$$\begin{aligned} \Delta_4 U_t = & 0.298 - 0.021U_{t-1} + 1.197\Delta_4 U_{t-1} - 0.216\Delta_4 U_{t-3} - \\ & (0.233) \quad (0.023) \quad (0.067) \quad (0.118) \\ & 0.445\Delta U_{t-4} + 0.429\Delta_4 U_{t-5} - 0.044LEVEL + \\ & (0.133) \quad (0.098) \quad (0.045) \\ F(G) & \left[ \begin{array}{ccc} 1.93 & - & 0.444U_{t-1} & - & 0.565\Delta_4 U_{t-1} \\ (0.841) & & (0.192) & & (0.263) \end{array} \right] + \varepsilon_t \\ \text{where } F(G) = & \left( 1 + \exp \left[ \frac{-19.62 \left( \frac{\Delta_4 \log GDP_{t-4} - 4.453}{\hat{\sigma}(\Delta_4 \log GDP_{t-4})} \right) + 0.184}{(12.3)} \right] \right)^{-1} \end{aligned}$$


---

*Long-run properties:*  $F(G) = 0 : \sum \phi_i = 0.965$ ;  $F(G) = 1 : \sum \phi_i = 0.400$ . *Diagnostic tests:* AIC: -2.09; SBIC: -1.78; Standard error of residuals,  $\hat{\sigma} = 0.33$ ; Samples Standard deviation of  $\Delta_4 \log GDP_{t-4} : \hat{\sigma}(\Delta_4 \log GDP_{t-4}) = 2.616$ ; Autocorrelation 1-4:  $F(4, 86) = 1.10[0.37]$ ; ARCH 4:  $\chi^2(4) = 0.99[0.91]$ ; Normality:  $\chi^2(2) = 6.33[0.04]$ ; Heteroscedasticity  $F_{xixj} : \chi^2(48) = 45.29[0.58]$ ; RESET test:  $F(1, 90) = 0.02[0.89]$ ;

---

Table 14: Parsimonious LSTAR model with annual GDP growth rate as transition variable

Lastly, Table 14 proposes a parsimonious LSTAR model with annual GDP growth rate as transition variable and with *Level* as explanatory variable. The cycle asymmetry property arises, since the sum in absolute value of the autoregressive coefficients is lower (higher) when changes in unemployment are below (above)



the threshold. Moreover, when compared with the baseline model, both regimes seem to be less persistent. The transition between regimes seems now faster ( $\hat{\gamma} = \frac{19.62}{2.616}$ ), which is in line with the theoretical framework in which LMI influence the adjustment to shocks. It is interesting to note that the sign of *Level* is also negative. Notice that, since the transition variable is the annual GDP growth rate and not the annual change of unemployment, it is possible to analyse how this LMI influences the equilibrium level of the regimes, as follows:

$$\begin{cases} \hat{u}_1 = \frac{-0.298+0.044Level}{-0.021} \\ \hat{u}_2 = \frac{-0.298-1.93+0.044Level}{-0.021-0.444} \end{cases}$$

Indeed, taking into account the minimum/maximum values of *Level*, we have:

$$\begin{cases} \hat{u}_1 = \begin{cases} \hat{u}_{Max} = 7.90 \\ \hat{u}_{Min} = 5.81 \end{cases} \\ \hat{u}_2 = \begin{cases} \hat{u}_{Max} = 4.51 \\ \hat{u}_{Min} = 4.42 \end{cases} \end{cases}$$

As expected, the proposed values from the baseline model are between the maximum and minimum values within each regime. In other words, depending on the value of the LMI, we may end up with both lower or higher equilibrium unemployment rates. Finally, the diagnostic tests do not indicate any misspecification problems.

Interestingly, the impact on LMI on unemployment, suggested by the previous models, are in line with the main literature on labour economics. Regarding *EPT*, Blanchard and Portugal (2001) argue that this impact is ambiguous, whereas Pissarides (2001) shows that *EPT* does not increase unemployment if chosen optimally. On the other hand, in what relates to *Level*, Calmfors and Driffill (1988) states that there is an inverted U-shape relationship between unemployment and the level at which wage bargaining takes place. Since *Level* can assume values from 1 to 5, and the minimum value for Portugal is 3 (i.e., an intermediate value), positive variations in *Level* are capturing negative relationship side between these two variables, which is in line with our results.

Finally, Figure 6 presents a cross plot of all transition functions (vertical axis) against the correspondent transition variable. Figure 6.a) corresponds to the models with annual change of unemployment as transition variable and Figure 6.b) to the models with annual GDP growth rate as transition variable. As we stated before, the transition between regimes seems to be rather fast in all the presented models, and the only case where it appears to exist a non negligible

change between the baseline models and the ones with the LMI is the model which considers EPT - see Figure 6.a).

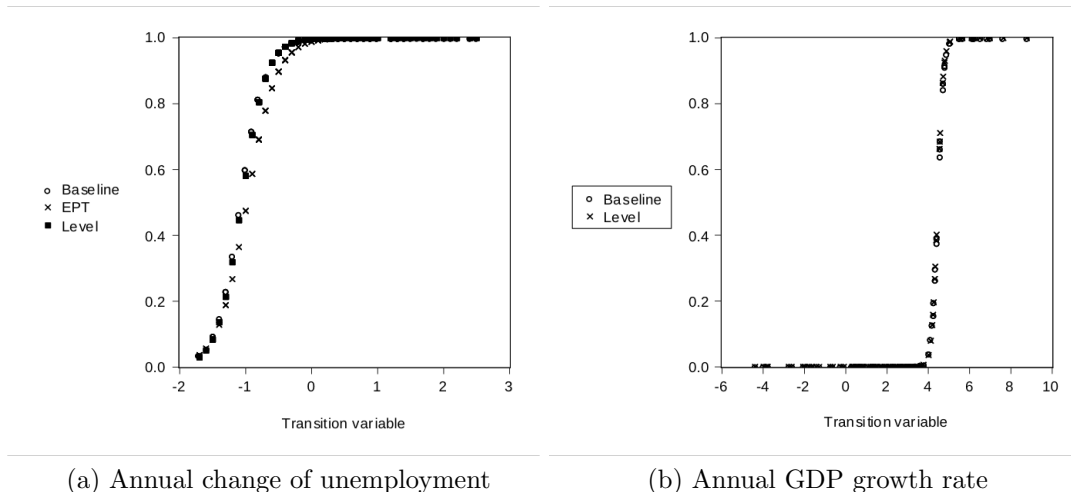


Figure 6: Cross plot of the transition function (vertical axis) against the transition variable. One dot represents at least one observation.

## 5 Conclusions

From the applied methodology and obtained results, six main conclusions can be drawn. First, the hysteresis hypothesis seems to be confirmed for the Portuguese unemployment rate, in line with Chang et al. (2005) and Lin et al. (2008). Second, unemployment behaviour is better described by a nonlinear model (LSTAR) rather than by an AR(5), using three types of transition variables: (a) annual change of cyclical unemployment (b) annual change of unemployment; and (c) annual GDP growth rate. Nevertheless, only the last two seem to correctly capture the cycle asymmetry behaviour and, in what relates to the latter, two unemployment regimes are suggested: a low regime with an equilibrium unemployment of 4.49% and a high regime with 7.11%. Third, from the beginning of the century, unemployment rate seems to be systematically in its high regime. Fourth, the transition between the two regimes appears to be rather fast. Fifth, LMIs seem to play an important rule in explaining the unemployment dynamics, affecting not only its regimes but also its equilibrium unemployment rate. Sixth, strong LMI appear to contribute to a lower unemployment rate. These results have strong implications in the design of labour market polices. As future work, we aim to (a) extend our methodology to other countries and (b) explore the differences between the LMIs among Europe and the OECD countries.

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## A Tests on structural breaks for Portuguese quarterly unemployment rate (1983:Q1-2013:Q4)

Specifications						
$z_t = \{1\}$	$q = 1$	$p = 0$	$h = 15$	$M = 5$		
Tests						
$SupF_T(1)$	$SupF_T(2)$	$SupF_T(3)$	$SupF_T(4)$	$SupF_T(5)$	$UDmax$	$WDmax$
69.02*	48.55*	92.13*	113.87*	165.52*	331.04*	649.06*
$SupF_T(2 1)$	$SupF_T(3 2)$	$SupF_T(4 3)$				
26.73*	18.25*	5.84				
Number of breaks selected						
Sequential	3					
LWZ	5					
BIC	5					
Estimates with two breaks						
$\hat{\delta}_1$	$\hat{\delta}_2$	$\hat{\delta}_3$	$\hat{\delta}_4$	$\hat{\delta}_5$		
7.68 *** (0.83)	-0.05 * (0.03)	1.48 (1.30)	3.61 *** (1.05)	6.02 *** (1.67)		
		$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$		
		1993Q3	2002Q2	2007Q4		

Table 15: Structural breaks: Bai and Perron (2003)

## B Correlations between LMIs variables

	Av. Nrr	Coord	EPT	Level	Ubdur	UD	U
Av.Nrr	1.000 ...						
Coord	-0.382 (0.847)	1.000 ...					
EPT	-0.900 (0.000)	0.320 (0.104)	1.000 ...				
Level	-0.157 (0.4252)	0.357 (0.057)	0.345 (0.078)	1.000 ...			
Ubdur	0.938 (0.000)	0.134 (0.497)	-0.752 (0.000)	-0.122 (0.535)	1.000 ...		
UD	0.156 (0.428)	-0.196 (0.309)	-0.736 (0.000)	-0.311 (0.100)	-0.051 (0.797)	1.000 ...	
U	0.156 (0.428)	-0.196 (0.309)	-0.736 (0.000)	-0.311 (0.100)	-0.051 (0.797)	0.105 (0.596)	1.000 ...

Table 16: Correlations between LMI variables and unemployment

Notes: Annual data (1983:2012). *p-values* are shown in square brackets.