Unemployment Hysteresis in Transitions Countries: Evidence from using stationarity panel tests with breaks*

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Abstract

This paper tests hysteresis effects in unemployment using panel data for Transition Countries covering the period 1992:1-2003:11. The tests exploit the cross-section variations of the series, and additionally, allow for a different number of endogenous breakpoints in the unemployment series. The critical values are simulated based on our specific panel sizes and time periods. The findings stress the importance of accounting for exogenous shocks in the series and give support to the natural-rate hypothesis of unemployment for the majority of the countries analyzed.

Key words: Hysteresis, panel unit root tests, structural break.

JEL classification: C22, C23, J64.

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1 Introduction

Enlargement is one of the most important challenges in the European Union (EU) agenda. The accession countries included in the enlargement process at present consist of ten Central and Eastern European Countries (CEECs), as well as Cyprus, Malta and Turkey. Since 1989, the process of transition has proceeded at a rapid pace. However, only eight out of the ten CEECs have fulfilled the so-called Copenhagen criteria set up in June 1993, and consequently, Bulgaria and Romania have been excluded from the May 2004 enlargement.

Focusing on the evolution of the labor markets, employment fell considerably in the CEECs during the transitional contraction period and continued to decline since then, despite transitory improvements in the economic growth pace from the middle of the nineties. Besides, there was a decrease in participation rates, which fell from the high levels typical of socialist economies. Unemployment exploded in the early transition years having a striking effect on poverty and social exclusion –see [?]. Since 1994, measured unemployment, based on labor market surveys following the ILO methodology[1], first decreased slightly below 10 per cent but increased again. However, in relative contrast to the overall pace of structural change in the transition countries, labor markets are characterized by very low mobility of workers across labor market strata, occupation and sectors ([?] and [?]).

The macroeconomic stabilization measures that these countries had to accomplish in order to meet the requirements for joining the EU, such as budgetary consolidations or inflation and exchange rate stabilizations are likely to have caused important shocks to output, prices and unemployment. Thus, unemployment is one of the key variables to facilitate the adjustment process through macroeconomic equilibrium. Moreover, with irrevocably fixed exchange rates, country-specific monetary conditions can no longer cush-

[1]The definition of and measurement of unemployment are neither very precise nor uniform among countries, so that a cross-country comparison of unemployment rates requires some adjustment to transform national measures into a reasonably standardized indicator. The standardized unemployment rates, which are based on labor market surveys, greatly improve comparability among countries. This measure, though, has some limitations as a measure of labor market slack, since it excludes discouraged workers, part-time employment, early retirement, government training and employment schemes and invalidity or disability schemes.

ion differences in cyclical positions nor help them to adjust to asymmetric shocks. Within a prospective enlarged euro area, if required, real exchange rate changes will have to be achieved by real wage changes directly, rather than indirectly via changes in the nominal exchange rate. The large rates of structural unemployment and the high regional concentration of unemployment suggest that labor market flexibility is not currently up to this requirement and, therefore, more geographic mobility would be needed (and expected). In the prospect of euro-area membership, the fulfillment of the Maastricht criteria will imply inflation rates in line with the 2% European Central Bank rate. Due to the real adjustment process involved, further employment destruction may be expected. In order to implement EU level policy measures to address the social problems associated, knowledge about the structural rate of unemployment and its shifting nature may be crucial for policy makers ([**?**]).

From a theoretical point of view, we can distinguish two main hypotheses relating unemployment and shocks. The first one, the so-called "natural" rate of unemployment or NAIRU. The concept of structural or "natural" rate of unemployment was first introduced by [**?**] and [**?**]. According to this approach, in the long-run the structural rate of unemployment is reached and hence there is no long-term trade-off between inflation and unemployment. However, in the short-term the Phillips Curve exists. From a statistical point of view, unemployment would be characterized as a mean reverting process, which means that the unemployment rate tends to revert to its equilibrium in the long run. According to the structuralist school the natural rate is endogenous and affected by market forces like any other economic variable (Pissarides, 1990, Layard et al., 1991) giving rise to autonomous movements of the natural rate due to changes either in real macroeconomic variables as real interest rates (Blanchard, 1999), rate of productivity growth (Pissarides, 1990), oil prices (Oswald, 1999) and stock prices (Phelps, 1999) or in the institutional framework such as the generosity of the unemployment-benefit welfare system, other forms of nonwage income, the family network, and the (consumption) tax wedge. The structuralist view would be in line with the existence of structural breaks of the steady-state path of a stochastic stationary process while hysteresis or persistence would be consistent with unit-root or near-unit root, processes, respectively. From a theoretical point of view, the slow adjustment process to that equilibrium is modelled by introducing real-wage rigidity through, for example, efficiency-wage or union behaviour models.

The second one, also known as the "hysteresis" hypothesis, states that shocks have permanent effects on the level of unemployment due to labour market rigidities as introduced by insider-outsider interactions (Blanchard and Summers, 1986) or human-capital effects (Layard et al., 1991), and, therefore, the level of unemployment can be characterized as a non-stationary process. It is worth to note that there is a crucial difference between the concepts of "hysteresis" and "persistence". Persistence implies a slow speed of adjustment towards the long run equilibrium level, and therefore, is a special case of the natural rate of unemployment hypothesis, as the series show mean reversion after all. From an econometric point of view it can be characterized by a near unit root process. If this is the case, macroeconomic policy would have long lasting but not permanent effects while, conversely, if hysteresis applies, the effects on unemployment are permanent. Sometimes the existence of persistence might be hiding changes in the level of the natural rate. This possibility has been pointed out by the structuralist view of the natural rate of unemployment (Phelps, 1994).

There is an extensive empirical literature on this hypothesis using time series, with mixed and sometimes counterintuitive results. The results of these studies applied to OECD almost uniformly fail to reject unit roots in the unemployment rates (with the exception of the US). The overwhelming evidence in favour of hysteresis was probably due to lack of power of the tests, pointing to the importance of either expanding the time span (and then allowing for discontinuities in the deterministic components as in Arestis and Mariscal (1999) and Papell et al. (2000)), or increasing the amount of information through panel data. More recently, there is a new generation of empirical papers using tests for unit roots in panels of countries trying to increase the power of the tests thanks to the increase of cross-section information, such as Song and Wu (1997, 1998) which strongly reject a unit root in the unemployment rate for US states using Levin and Lin test, and León-Ledesma (2002) which is able to reject hysteresis for the US but not for the EU using the IPS tests. The presence of structural breaks has been also taken into account in panel data (Strazicich et al. (2001), and Murray and Papell (2000)) However, the number of empirical studies is still very scarce for Transition countries. Exceptions are the papers by León-Ledesma and McAdam (2004) and Camarero, Carrion-i-Silvestre and Tamarit (2005), that reject the hysteresis hypothesis using univariate methods.

In this paper we contribute to the empirical literature in several respects. First, we apply both panel unit root and stationarity tests; in particular, those proposed by Im, Pesaran and Shin (1997, 2003)[2] and Maddala and Wu (1999) for the null of unit root, and Hadri (2000) tests for the null of stationarity. Second, we use two versions of each of these tests: the first one, imposing cross-section independence and, the second one, allowing for dependence and computing critical values by boostrap techniques. Third, we apply a new panel stationarity test incorporating multiple structural changes endogenously determined as proposed by Carrion-i-Silvestre et al. (2005), also accounting for cross-correlation in the residuals. Accounting for these two features (structural breaks and dependence) provide important power gains compared to the time series equivalent tests.

The remainder of the paper is organized as follows. Section 2 briefly describes the tests used in the paper, and the econometric results. Finally, in section 3 we report the main results and conclusions.

2 Empirical results

The main limitation for the analysis is the short span of the statistical information available for these new EU countries. The standard sources of statistical information such as the OECD, AMECO and EUROSTAT databases just offer a short sample of unemployment rates for these countries. In this Section we analyze the order of integration of the unemployment rates for all the CEECs countries acceding to the EU in May 2004. Due to the particularities of this group of countries, we have a constraint concerning the time span available for any economic variable. Thus, we have decided to use monthly data,

[2]IPS hereafter.

in order to increase the number of observations[3]. The monthly harmonized unemployment rates have been taken from EUROSTAT (Euroindicators) for the period 1998:12 to 2003:11. Then, we have applied backwards the growth rates of the monthly unemployment rates drawn from [?] to extend the database[4] that, at best, covers from 1991:1 to 2003:11[5]. **These harmonized unemployment rates are depicted in Figure 1.**

2.1 Panel unit root and stationarity tests. No breaks case

Thus, in order to test for hysteresis in the unemployment rate the empirical results presented in this section are organised in two groups. First, we test the null hypothesis of unit root using the t-bar ($\Psi_{\bar{t}}$) and LM-bar ($\Psi_{\overline{LM}}$) statistics in Im, Pesaran and Shin (1997, 2003), and the MW statistic in Maddala and Wu (1999) panel unit root tests, as well as the Hadri (2000) stationarity test.[6] Since these statistics are now well known, we address readers to the respective papers. Second, we apply the Carrion-i-Silvestre et al. (2005) test allowing for structural changes endogenously determined in a panel context, which improves largely the power of the time series test used in Papell et al. (2000). Additionally, this test allows for multiple number and type of breaks and accounts for cross-correlation in the residuals, solving the main drawbacks in panel studies above mentioned.

Before presenting these results we should introduce some comments on the deterministic specifications that are used along the paper. We should bear in mind that the rejection of hysteresis establishes that the unemployment rate evolves in a stationary way around the natural rate. Thus, the deterministic specification when testing both for the unit root hypothesis or for the stationarity hypothesis is the one given by a constant term. Although looking at the pictures of the variables in Figure 1 in the Appendix one could

[3] However, we are aware of the limitations of proceeding this way, since the increase of the frequency does not imply an increase of the long-run information.

[4] We thank Miguel León-Ledesma for kindly providing us the data.

[5] Specifically, for the Czech Republic the data spans from 1991:1 to 2003:10, Estonia (1995:5, 2003:11), Hungary (1991:3, 2003:11), Latvia (1994:1, 2003:11), Lithuania (1994:1, 2003:11), Malta (1997:3, 2003:10), Poland (1991:1, 2003:11), Slovakia (1991:1, 2003:11) and Slovenia (1992:1, 2003:11).

[6] When computing the panel stationarity test in Hadri (2000) we have estimated the long-run variance using the procedure in Sul, Phillips and Choi (2003), which reduces size distortions when stochastic processes are close to non-stationarity. Further details are given below and in Carrion-i-Silvestre and Sansó (2005).

decide to include a time trend in most of them, this specification would mask the fact that the unemployment rate might be experiencing a long transition between shifting natural rates.[7] This is pointed out in Papell et al. (2000) where it is mentioned that while a nonzero trend for unemployment does not make sense asymptotically, a slowly increasing natural rate could be represented by trend stationarity process in small samples.

Let us first focus on results without structural breaks. The results of the panel data unit root and stationarity tests applied to the unemployment rate are reported in Panel A of Table 1. Assuming that the individuals are cross-section independent, **all the tests mainly point to the presence of hysteresis in unemployment for the set of CEECs countries that has been analysed. Thus, the unit root hypothesis cannot be rejected by neither the IPS nor the MW tests. In addition, the test in Hadri (2000) strongly rejects the null hypothesis of stationarity. This conclusion is reached irrespectively of the deterministic specification.**

The assumption of cross-section independence is rarely found in practice, especially in a globalised economy where the shocks overpass the borders. However, these countries are in a process of opening-up. This is of special interest in our study, due to the inclusion in the panel data set of twelve EU countries, which in part are ruled by common governmental institutions. These facts question the validity of this assumption. In order to account for cross-section dependence, we have followed two approaches. Firstly, the independence assumption can be relaxed to allow for time-varying aggregate effects in the data. These effects can be removed by subtracting the cross-section mean from the data –see O'Connell (1997), and Levin, Lin and Chu (2002). **The results that are obtained after removing the cross-section mean are reported in Panel B in Table 1. The main drawback of this approach is that it assumes that the effect of the cross-section dependence is the same for all individuals.** In order to account for more general situations we have decided to compute the bootstrap distribution of the tests. This is the second approach that is followed here. The details of the bootstrap are given

[7] This is especially true in the case of France, New Zealand, Spain, Norway and Japan.

in Maddala and Wu (1999) with 2,000 replications for the bootstrap. Panel C in Table 1 reports the percentiles of interest of the bootstrap distribution.

Let us first focus on the results based on cross-section demeaning. Thus, except for the t-bar IPS statistic the null hypothesis of unit root cannot be rejected at the 5% level of significance, while stationarity is strongly rejected using the panel KPSS statistic. When we compare the statistics with the boostrap distribution we conclude that, except for the panel KPSS statistic that specifies a time trend, evidence points to non-stationarity.

In all, the results support the hysteresis hypothesis in unemployment rates, a conclusion that is robust to the presence of cross-section dependence. In general, this conclusion is also in accordance with the previous results in the literature. However, **it should be noted that evidence of stationarity is only found when the time trend is used with panel KPSS statistic. As pointed above, this deterministic specification can be masking the presence of structural breaks that might be shifting the natural rate.** This fact is not surprising as the natural rate depends on the fundamentals of the economies and these fundamentals change in accordance to the technological progress. Moreover, this contradiction between the unit root and stationarity tests can be thought to be an indicator of the presence of structural breaks –see Cheung and Chinn (1997).

2.2 Panel stationarity tests allowing for structural breaks endogenously determined

In order to account for this feature we proceed to compute the extension of the Hadri (2000) test for stationarity in panel data with multiple structural changes under the null hypothesis, which is proposed in Carrion-i-Silvestre et al. (2005). This framework allows for heterogeneity in several respects: multiple structural changes, multiple structural changes positioned at different unknown dates, and a different number of breaks for each individual. The details of this technique are described in the Appendix.

In order to detect the breaks, Carrion-i-Silvestre et al. (2005) suggest applying the procedure first proposed in Bai and Perron (1998). This consists of specifying a maximum number of breaks ($m^{\max}$), estimating their position for each $m_i \leq m^{\max}$, $i = \{1, ..., N\}$, testing for the significance of the breaks and, then, obtaining their optimum number and position for each series.

First, to estimate the dates of the breaks, they choose the argument that minimizes the sequence of individual SSR, as in Bai and Perron (1998). Some trimming would be necessary, that is commonly specified as $T_b^i \in [0.15T, 0.85T]$. Once the dates for the possible breaks have been estimated, then the number of optimal structural breaks should be selected for each i (that is, the optimal m_i). Bai and Perron (2001) compare two alternative procedures: information criteria (such as the Bayesian information criterion (BIC) and the modified Schwarz information criterion (LWZ) of Liu, Wu and Zidek (1997)) and the sequential computation of structural breaks, using pseudo F-type test statistics. They recommend using the LWZ criterion when the model includes trending regressors, whereas for non-trending ones the sequential procedure has better performance.

The results of the computation of the $LM(\lambda)$ test allowing for up to $m^{\max} = 5$ breaks, with the deterministic specification given by Model 1, are reported in Table 2. The number of breaks has been selected using the sequential procedure in Bai and Perron (1998). Panel A in Table 2 offers the individual information, *i.e.* the individual KPSS test, number of breaks and their position. In general, at least one structural break was detected by the sequential procedure in all the countries considered and, in six cases, we found up to four breaks. This finding may suggest that the analysis conducted in the previous Section would be wrong if these structural breaks were relevant for the analysis of the stochastic properties of the series. **The last two columns of Panel A show the 10 and 5% critical values computed by simulation of the individual KPSS tests with structural breaks. The null of stationarity cannot be rejected for the countries considered, with the only exception of Canada, Italy and New Zealand at the 5% level.**

If we combine the individual information to compute the $LM(\lambda)$ test in Panel A, we realise that the null hypothesis of stationarity cannot be rejected

when the test is computed using the homogeneous long-run variance estimate, although it is strongly rejected when using the heterogeneous long-run variance one. The same result is obtained when using the cross-section demeaned data –see Panel B in Table 2. However, this contradiction disappears when cross-section dependence is taken into account. Thus, the critical values drawn from the Bootstrap distribution indicate that the null hypothesis cannot be rejected at the 5% level –see Panel C in Table 2. Therefore, our results point to the absence of hysteresis in the unemployment rate of the OECD countries analysed.

There are important differences in our tests results compared with other empirical studies using panel techniques. Song and Wu (1998) do not consider the existence of breaks in their tests, whereas Papell et al. (2000) allow for multiple breaks but do not apply any formal panel data test. Strazicich et al. (2001) apply a *LM* ADF-type panel unit root test allowing for two breaks that does not account for the residual cross-correlation. Although they reject the hysteresis hypothesis for the panel, looking at the individual information that they also present, in only two cases the unit root can be rejected. The restriction in the number of breaks provokes, in our opinion, a misspecification problem of the deterministic component.

2.2.1 Explaining the breaks: the role of reforms in the transition economies

The last two columns of table 3 are devoted to summarizing the main results obtained by [**?**] for a similar group of countries. We should, however, emphasize that the results are not directly comparable. The unit root test they have applied allows for a break in a model for trending variables. In contrast, our one-break tests are applied to non-trended variables. Despite this difference, the results are similar to ours in the 1-break tests for the cases of the Czech Republic, Estonia and Hungary, and partially from our n-test results for Latvia and Poland. Finally, in column 5 we present [**?**] dates of regime change that can be derived from their application of the Markow-Switching methodology. Again, the changes in regime detected are coincident with these we found in the cases of the Czech

Republic, Estonia, Latvia, Lithuania, Slovakia and Slovenia. Significant discrepancies are only found in Hungary.

In 1989 the transition from centrally planned to market economies began embedding important common reforms: price liberalization accompanied by more disciplined fiscal and monetary policies; privatization of firms through different methods; a reform of the financial sector and, finally, an external sector reform (trade liberalization, currency convertibility and exchange rate regime choices). However, depending on the countries, the programs differed on several respects: wage controls; privatization programs (spontaneous privatization, using vouchers and restitutions or management-employment buy-outs); the choice of exchange rate system (many countries pegged their exchange rate, while others floated); finally, subsidies were removed at varying speeds. Overall, during this transition period a big amount of measures have been implement in all these countries but at a different pace in each of them. In fact, a key debate among policy makers has been the choice between the gradualist approach and a shock therapy ([**?**]). The argument for gradualism is that it avoids the output and employment decline associated with a shock therapy. In contrast, shock therapy involves a immediate economic adjustment to the market economy. In a nutshell, labor reallocation was deemed to occur mainly through unemployment, the single most important indicator of the speed of transition trajectories ([**?**]).

The labor markets of the former centrally planned economies on the eve of transition were characterized by full employment. This "full employment" was achieved at the cost of low wages and a large amount of hidden unemployment (e.g. about 30% according to some estimates). Employment was concentrated in heavy industries and in the public sector, with private initiative only being tolerated in the agricultural sector. Most Accession countries experienced substantial falls in GDP and wages at the initial stage of transition. The exceptions are the Czech Republic, that experienced a long period of low unemployment, and Estonia, that achieved significant labor reallocation from the beginning of the transition process. The transition period has been characterized by job shedding in the public sector, job creation in the private sector with an increased incidence of temporary (frictional) unemployment and a significant level of structural unemployment. The temporary mismatch between labor demand and supply has been due to the length of time taken to develop new private enterprises and the process of privatization.

Labor reallocation is a critical aspect of the transition process and because there is a significant variation in the timing of reforms across transition countries we find important asymmetries in the trajectories these countries have followed. Apart from this reason, there is a large body of theory mainly developed within the OECD framework that suggests that different types of social policy and labor market institutions influence greatly the distinct trajectories of adjustment in the new EU countries.

The role of institutional labor market rigidities is a matter of increasing concern. From the seminal papers by [?] and [?] there is an increasing empirical literature about the role that institutional factors play in determining the persistence of unemployment making clear their crucial importance not only for the determination of the structural unemployment rate but also for the speed of labor market adjustments[8]. Apart from shocks (variations in productivity, labor demand, real import price or real interest rates) and macropolicies, the longer-term patterns of unemployment tend to be dominated by shifts in the equilibrium rate. The speed of adjustment to that equilibrium will be affected by any variable which influences the ease with which unemployment individuals can be matched, and secondly, by any variable which tends to raise wages (despite excess supply in the labor market). Most of these variables reflect market institutions such as unemployment benefits or unions. There are four aspects of the unemployment benefit system that may influence the speed adjustment to the equilibrium: the level of benefits, the duration of entitlement, the coverage of the system and the strictness with which the system is operated. Of these, only the first two are available as time series for the CEEC countries that belong to the OECD.

According to [?] the transition process involves new job matches using workers with different skills that should generate an explosion of earning differentials at all levels, be-

[8] See [?], [?] and [?].

tween the public and private sectors, between and within firms as well as across regions. However, these inequalities may be mitigated by institutions imposing wage floors (e. g. unions, tax income policies, minimum wages and employment protection). These institutions are constraints to the adjustment process, generating more unemployment. However, the scarce empirical evidence existing up to date tends to signal a modest influence of the variables on the labor market in most transition economies. This effect was due to the lack of credibility of the old unions which had supported the communist regime, the intermittent use of the tax income policies (by 1995 many of the CEECs had abandoned these policies) and a lax use of the minimum wages (by 1996 minimum wages had fallen to about 30% of the average wage in all CEECs). Minimum wages were not binding in Hungary. However, minimum wages played some role indirectly as they served in all transition countries as a basis for calculating most social benefits (e. g. welfare, unemployment, and health benefits). Finally, the role played by employment protection regulation (severance pay and notice periods) seems to have been rather limited, especially compared to active policies (wage subsidies, direct job creation and schemes for school leavers) that have been pretty successful in some countries like the Czech Republic. The best studied item for the case of the transition countries has been the impact of unemployment programs (both active and passive) on the duration of unemployment and the probability of finding a job. Unemployment benefits were initially set at relatively high levels and provided in some cases for unlimited duration. However, as the number of beneficiaries was increasing rapidly, the conditions became more strict. The maximum duration was halved in Czechoslovakia and in Hungary and reduced to one year in Poland. The unemployment benefits (welfare assistance, disability benefits and sickness benefits) have played a non-standard function of relatively high importance in the transition countries as they have put "de facto" a floor to wage setting.

The inspection of the graphs provides also further evidence on this last issue. In the majority of the cases, the breaks are reflecting an increase in unemployment and, therefore, the associated coefficients are positive.

In all, the results point to the rejection of the hysteresis hypothesis and

are compatible with the structuralist theories as described by Phelps (1994) meaning that the majority of shocks to unemployment are temporary but, occasionally, and mainly associated with recessions, shocks can provoke a change in the level of the natural rate of unemployment.

3 Conclusions

In this paper we review the empirical validity of hysteresis in unemployment rates for a group of CEECs countries using annual data for the period 1992-2003. The hysteresis hypothesis can be easily tested in a framework based on unit root or stationarity tests. Therefore,

To summarize the results, the rejection of hysteresis in unemployment depends critically on the above mentioned characteristics of the tests. First, using panel unit root tests we cannot reject hysteresis in unemployment, even when allowing for cross-section dependence. Second, there is mild evidence in favour of the natural rate hypothesis with panel stationarity tests, homogeneous long-run variance and cross-section dependence. Finally, the results change dramatically when we also allow for structural breaks in the stationarity tests: hysteresis in unemployment is not only strongly rejected in the panel, but also in the individual country tests. Moreover, the dates of the breaks are consistent with the results in previous literature and support the structuralist view of unemployment meaning that temporary shocks have highly persistent but not permanent effects on unemployment. At the same time, structural factors can affect the natural unemployment rate and, therefore, unemployment would be stationary around a process that is subject to structural breaks.

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Table 1: IPS and Maddala and Wu (MW) panel unit root tests

	Panel A: Assuming cross-section independence				**Panel B:** Removing cross-section mean			
	Constant		Time trend		Constant		Time trend	
	Test	p-val	Test	p-val	Test	p-val	Test	p-val
$\Psi_{\bar{t}}$	0.826	0.796	0.044	0.518	-1.775	0.037	-0.906	0.182
$\Psi_{\overline{LM}}$	-1.448	0.926	-0.213	0.585	1.406	0.079	1.078	0.140
MW	22.829	0.975	30.018	0.819	47.872	0.131	50.376	0.086
Hadri (Hom.)	47.742	0.000	5.399	0.000	24.940	0.000	6.586	0.000
Hadri (Het.)	47.767	0.000	5.424	0.000	22.158	0.000	7.489	0.000

Panel C: Bootstrap distribution (allowing for cross-section dependence)

Constant

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-4.035	-3.397	-2.860	-2.264	2.481	3.479	4.452	5.737
$\Psi_{\overline{LM}}$	-2.991	-2.464	-2.024	-1.464	2.744	3.483	4.092	4.893
MW	9.645	13.406	16.921	22.096	63.065	69.929	77.858	85.841
Hadri (Hom.)	-3.236	-3.030	-2.814	-2.510	4.838	7.247	9.104	11.246
Hadri (Het.)	-2.840	-2.576	-2.377	-2.109	4.877	7.184	8.997	10.904

Time trend

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-4.708	-4.108	-3.636	-3.065	1.539	2.405	3.276	4.285
$\Psi_{\overline{LM}}$	-3.014	-2.296	-1.719	-1.090	3.272	3.879	4.399	5.064
MW	18.256	22.952	27.221	32.504	78.465	86.399	94.495	102.581
Hadri (Hom.)	-2.240	-1.875	-1.535	-1.008	4.866	5.964	7.204	8.557
Hadri (Het.)	-1.384	-1.075	-0.714	-0.259	5.266	6.206	7.517	9.058

Table 2: Panel KPSS tests and individual test. Sample 1956-2001 (T=46)

Panel A: Individual information

	Individual tests	m_i	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$	10%	5%
Australia	0.029	4	1974	1981	1989	1995	0.081	0.101
Austria	0.052	2	1961	1981			0.123	0.147
Belgium	0.035	4	1974	1980	1986	1992	0.083	0.104
Canada	0.115	3	1974	1981	1995		0.094	0.114
Denmark	0.036	3	1961	1974	1995		0.102	0.125
Finland	0.023	2	1975	1991			0.107	0.128
France	0.032	3	1974	1980	1991		0.087	0.106
Germany	0.032	4	1961	1974	1981	1992	0.061	0.071
Ireland	0.049	3	1974	1982	1995		0.091	0.110
Italy	0.216	3	1961	1974	1982		0.091	0.110
Japan	0.067	3	1974	1981	1995		0.094	0.114
Netherlands	0.049	4	1973	1980	1988	1995	0.076	0.094
Norway	0.037	4	1971	1981	1988	1995	0.068	0.082
New Zealand	0.340	3	1980	1988	1994		0.124	0.159
Spain	0.129	2	1974	1980			0.123	0.149
Sweden	0.086	1	1991				0.227	0.297
Switzerland	0.056	3	1961	1982	1991		0.099	0.123
United Kingdom	0.049	4	1974	1980	1987	1995	0.081	0.102
USA	0.089	3	1974	1986	1995		0.089	0.108

Panel Stationarity tests (Assuming cross-section independence)

	Test	p-val
Homogeneous	-0.385	0.650
Heterogeneous	3.216	0.001

Panel B: Removing cross-section mean

	Test	p-val
Homogeneous	-0.977	0.836
Heterogeneous	1.766	0.039

Panel C: Bootstrap distribution (allowing for cross-section dependence)

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
Homogeneous	0.885	1.171	1.464	1.950	5.811	6.642	7.331	8.291
Heterogeneous	2.055	2.488	2.833	3.245	8.134	9.268	10.200	11.543

Table 3: Structural breaks. Comparison of the different methods

	Bai-Perron n breaks in Camarero et al. (2005)	León-Led/McAdam 1 break (trend model)	León-Led/McAdam Markow Switching
Czech Rep.	1996:07 1998:07	1992:07 1998:04	1997-98
Estonia	1997:05 1999:03 2001:10	1998:10 2000:05	Multiple changes (1995, 1996, 1999...)
Hungary	2000:03	1992:11	Multiple changes
Latvia	1996:01 1998:09 2000:08 2002:06	1998:04	1998 2000
Lithuania	1995:05 1999:11 2002:01	1997:01	1998-1999
Malta	1992:11 1999:07 2001:11		
Poland	1992:11 1996:08 2000:01	1996:03	
Slovakia	n=1: 1998:12 n=2: 1993:01 n=2: 1999:02	1992:11	1998
Slovenia	1993:09 2000:04	1999:06	1994 1996 2000

Table 4: Dates of the breaks. Political and institutional events.

	Breaks and main events										
Countries	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002
Czech Rep.					**1996:07**		**1998:07**				
						Havel lost power					
						Post-transition recession					
							Restructuring				
Estonia						**1997:05**		**1999:03**		**2001:10**	
							Russian crisis→			Current account	
										deficit	
Hungary									**2000:03**		
				Early reform						ERM-2	
				Stabilization Plan (1995)					Employment recovery		
Latvia					**1996:01**		**1998:09**		**2000:08**		**2002:06**
	Tiny open economy										
	Exposed to external shocks						Russian crisis→				
Lithuania				**1995:05**				**1999:11**			**2002:01**
				Exchange			Russian crisis→				New peg
				rate peg						Faster privatization	
										Fiscal discipline	
Malta							**1998:02**	**1999:07**		**2001:11**	
	Tiny open economy									September 11^{th}	
	Exposed to external shocks									Tourism	
Poland	**1992:11**				**1996:08**				**2000:01**		
	Shock therapy for transition				Monetary tightening		Russian recession (↓GDP)				
Slovakia		**1993:01**						**1999:02**			
		Splits from					1998 elections				
		Czech Rep.									
Slovenia		**1993:09**							**2000:04**		
	1991: splits from Yugoslavia										

A Carrion-i-Silvestre, del Barrio and López (2005) panel stationarity tests with multiple breaks

These authors specify the following DGP under the null hypothesis of stationarity:

$$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \beta_i t + \sum_{k=1}^{m_i} \gamma_{i,k} DT^*_{i,k,t} + \varepsilon_{i,t} \tag{1}$$

with $DU_{i,k,t} = 1$ for $t > T^i_{b,k}$ and 0 elsewhere, $DT^*_{i,t,k} = t - T^i_{b,k}$ for $t > T^i_{b,k}$ and 0 elsewhere, where $\{\varepsilon_{i,t}\}$ are assumed to be independent across i –this assumption will be addressed below. This model includes the following elements: (i) Individual effects, that are in fact individual structural break effects (or shifts in the mean caused by the structural breaks), (ii) temporal effects if $\beta_i \neq 0$ and (iii) temporal structural break effects if $\gamma_{i,k} \neq 0$ (when there are shifts in the individual structural time trend).

This specification encompasses Model 1 in Perron and Vogelsang (1992) when $\beta_i = \gamma_{i,k} = 0$ and Model C in Perron (1989), that Carrion-i-Silvestre et al. (2005) call Model 2, when $\beta_i \neq \gamma_{i,k} \neq 0$. This specification has very convenient characteristics: (i) The structural breaks may have different effects on each individual time series (these effects are measured by $\theta_{i,k}$ and $\gamma_{i,k}$), (ii) these breaks can be located at different dates, because they do not impose the restriction $T^i_{b,k} = T_{b,k}, \forall i = \{1, ..., N\}$, and (iii) the individuals may have different numbers of structural breaks, so that $m_i \neq m_j, \forall i \neq j, \{i, j\} = \{1, ..., T\}$.

The test is formulated as in Hadri (2000), *i.e.*, the average of the individual KPSS statistic. The general expression takes the form:

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

where $S_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ denotes the partial sum process obtained form the OLS residuals of equation (1), and $\hat{\omega}^2 = N^{-1} \sum_{i=1}^{N} \hat{\omega}^2_i$, where $\hat{\omega}^2_i$ is a consistent estimate of the long-run variance of $\varepsilon_{i,t}$. The procedure that is applied to estimate ω^2_i is extremely important.

Thus, Caner and Kilian (2001) show that stationarity tests such as the KPSS statistic suffers from severe size distortion when the stochastic process is near to non-stationarity. Carrion-i-Silvestre and Sansó (2005) have shown that this size distortion can be reduced if the long-run variance is properly estimated. In this regard, Carrion-i-Silvestre and Sansó (2005) compare different ways to estimate ω_i^2 and suggest using the procedure described in Sul, Phillips and Choi (2003), which is also used in Carrion-i-Silvestre et al. (2005). In brief, their proposal bases on the application of a prewithened Heteroskedasticity and Autocorrelation Consistent (HAC) variance estimate, which in the first stage implies estimating an AR model for the residuals of (1):

$$\hat{\varepsilon}_{i,t} = \vartheta_1 \hat{\varepsilon}_{i,t-1} + \ldots + \vartheta_p \hat{\varepsilon}_{i,t-p} + \psi_{i,t}, \tag{3}$$

and obtaining the long-run variance of the estimated residuals in (3), which is denoted as $\tilde{\sigma}_{\psi i}^2$, through the application of a HAC estimator –for instance, Bartlett or Quadratic Spectral window– to control for the presence of heteroskedasticity. In the second stage the estimated long-run variance is recolored:

$$\hat{\omega}_i = \frac{\tilde{\sigma}_{\psi i}^2}{\tilde{\vartheta}(1)^2},$$

where $\tilde{\vartheta}(1)$ denotes the autoregressive polynomial $\tilde{\vartheta}(L) = 1 - \tilde{\vartheta}_1 L - \ldots - \tilde{\vartheta}_p L^p$ evaluated at one. In order to avoid the inconsistency of the test statistic, Sul, Phillips and Choi (2003) suggest using the following boundary condition rule to obtain the long-run variance estimate:

$$\hat{\omega}_i = \min\left\{ T\tilde{\sigma}_{\psi i}^2, \frac{\tilde{\sigma}_{\psi i}^2}{\tilde{\vartheta}(1)^2} \right\}.$$

The application of this rule ensures that the long-run variance estimate is bounded above by $T\tilde{\sigma}_{\psi i}^2$. In all computations the order of the AR model in (3) is chosen by the BIC information criterion specifying 5 lags as the maximum, and $\tilde{\sigma}_{\psi i}^2$ is obtained using the Quadratic spectral window as depicted in Sul, Phillips and Choi (2003).

Note that we do not require to assume homogeneity of the long-run variance across individuals, so that the expression (2) can include separate estimates for the long-run variance of each individual. The parameter λ denotes the dependence of the test on the dates of the break. The vector $\lambda_i = (\lambda_{i1}, ..., \lambda_{i,m_i})' = (T^i_{b,1}/T, ..., T^i_{b,m_i}/T)'$ indicates the relative positions of the dates of the breaks on the time period T. Finally, the normalized test statistic converges to a standard Normal distribution and turns out (according to Carrion-i-Silvestre et al. (2005) Monte Carlo results) to be more suited for panels with larger T compared to N.