Unemployment, Hysteresis and Transition

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Abstract: We quantify the degree of persistence in the unemployment rates of transition countries using a variety of methods benchmarked against the EU. In doing so, we will also characterize the dynamic behavior of unemployment in the CEECs during the past decade. In part of the paper, we will work with the concept of linear "Hysteresis" as described by the presence of unit roots in unemployment as in most empirical research on this area. Given that this is potentially a rather narrow definition, we will also take into account the existence of structural breaks and non-linear dynamics in unemployment in order to allow for a richer set of dynamics. Finally, we examine whether CEECs' unemployment presents features of multiple equilibria, that is, if it remains locked into a new level whenever a structural change occurs. Our findings show that, in general, we can reject the unit root hypothesis after controlling for structural changes and business cycle effects, but we can observe the presence of a high and low unemployment equilibria. The speed of adjustment is faster for CEECs than the EU, although CEECs tend to move more frequently between equilibria.

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1. Introduction.

One of the foremost features of the transition process of CEECs is the appearance of open unemployment hidden during the central planning regime. As reported in EBRD (2000) and IMF (2000), this phenomenon has had a deep impact on poverty and social exclusion experienced. This is partly due to the fact that the comprehensive social safety net left agents with little experience in dealing with the uncertainty and adversity associated with protracted unemployment. The labor-market reforms of the early 1990s, especially the reduction of unemployment benefits, did not seem to have the expected impact in reducing unemployment by improving matching (Boeri, 1997).[1] Employment expanded at a much slower pace than output, pointing to a high degree of persistence in unemployment, thus aggravating the social problems associated with the transition to a market economy. Furthermore, with the prospect of EU membership, accession countries will continue to pursue both product and labor market reforms that are likely to exert important shocks on employment (EBRD, 2000). This is especially true if labor hoarding is reduced by the introduction of foreign and domestic competition. Shocks are also likely to come about for some countries because of macroeconomic stabilization measures (i.e., budgetary consolidation, inflation and exchange rate stabilization) to meet the requirements for joining the EU.

In this paper, we quantify the degree of persistence in the unemployment rates of transition countries using a variety of methods. In doing so, we will also characterize the dynamic behavior of unemployment in the CEECs during the past decade. As far as we are aware, this is the first systematic attempt to describe the persistence pattern of aggregate unemployment for this set of countries. In part of the paper, we will work with the concept of linear "Hysteresis" as described by the presence of unit roots in unemployment as in most empirical research on this area. Given that this is potentially a narrow definition,[2] we will also take into account the existence of structural breaks

[1] For comprehensive reviews of labor market developments in Transition Economies see also EBRD (2000) and Vidovic (2001). See also Tichit (2000) for a comparative study of unemployment dynamics among Eastern European countries.

[2] For in-depth discussions of the concept and implications of Hysteresis in unemployment see Amable *et al.* (1995), Cross (1995) and Røed (1997).

and non-linear dynamics in unemployment in order to allow for a richer set of dynamics. Finally, we examine whether CEECs' unemployment presents features of multiple equilibria. That is, if it remains locked into a new level whenever a structural change occurs.

The question addressed is important for several reasons. First, it has important implications for social protection and labor market reforms, as well as macro-stabilization policy in the CEECs. The presence of hysteretic or highly persistent unemployment would imply that unemployment could become a long-lasting problem after radical macroeconomic and labor market policy reforms. Secondly, it helps to understand if the aggregate behavior of unemployment in our set of countries is consistent with recently developed models of labor markets in transition briefly described in the next section. Comparison with the persistence profile in EU countries could also help analyze the possible impact of common shocks. For instance, if unemployment were to be more persistent in the CEECs than in the EU, common negative shocks to both areas could increase migration pressures westwards, and common positive shocks reduce them. The paper also contributes to recent theory and empirical studies addressing the issue of Hysteresis in unemployment by carrying out our tests in a group of economies with a rapidly changing labor market.

In order to undertake our empirical analysis, we first work with the concept of Hysteresis as stemming from the presence of a unit or near-unit root in unemployment rates. We apply a battery of unit root tests on a set of 12 CEECs (benchmarked against an EU-15 aggregate) to test for the existence of random-walk behavior, quantify the degree of persistence and account for possible breaks in our sample and lack of power in our tests. Secondly, we use Markov Switching regressions to analyze persistence taking into account the possibility of a changing equilibrium unemployment due to breaks and large business-cycle fluctuations. This will, most importantly, allow us to work with a concept of Hysteresis as multiple equilibria in unemployment. In the next section we provide an overview of the evolution of unemployment in our sample of CEECs[3] and some theoretical models attempting to explain it.

[3] The CEECS in our sample comprise Poland, Romania, Slovenia, Croatia, Hungary, Bulgaria, Czech Republic, Slovakia, Estonia, Latvia, Lithuania and Russia.

2. Stylized facts and theoretical background.

The evolution of unemployment in Eastern Europe has showed a diversity of patterns depending on the particular transition conditions of each country. Figure 1 plots unemployment rates for a set of 12 Eastern-European transition countries. Overall, we can observe relatively high levels of unemployment during the past decade, which, in most cases, reach double-digit figures (exceptions are the Czech Republic, Estonia and Latvia). The first two show low unemployment levels during most of the 90s' and then a steady increase in the last few years of the sample. Notably, Russia shows a sharp increase in unemployment due to the 1998 crisis that starts to recover after 2000. For the rest of the countries, we can observe either an inverted L-shaped behavior or a slight recovery after 1995 followed by another sharp increase by the end of the decade.

HERE FIGURE 1

Unemployment in these countries arises as a consequence of the rapid process of structural change and as the inevitable consequence of labor-market reforms.[4] However, as Boeri and Terrell (2002) point out, more than the rate of employment destruction, it is the low rate of employment creation that has led to the existence of stagnant pulls of long-term unemployed. This is especially the case in CEECs, whereas Russia and the CIS countries have shown consistently lower levels of unemployment. This happens even when the output collapse in former Soviet Union countries has been far larger than in most CEECs. This lower elasticity of employment with respect to output (essentially labor hoarding) is one of the main differences in employment performance between these two groups of countries.[5] The difference is related to the fact that wage adjustment has been a more prominent feature of labor market dynamics in Russia whereas employment has been the main

[4] See Boeri (1997) for a review of labor market reforms in transition economies.

[5] Svejnar (1999) reports insignificant elasticities of employment to output for Russia and elasticities within the rage of 0.2 and 0.8 for CEECs.

adjustment variable in CEECs, pointing out to a higher degree of persistence of unemployment in the latter group (Boeri and Terrell, 2002).

The unprecedented process of structural change that shook CEECs' labor markets has not been absorbed as expected by the creation of new jobs in the private sector and the improvement of matching induced by more market-oriented labor-market policies (EBRD, 2000). This has led to the high unemployment observed in the CEECs together with persistent and long duration of unemployment spells. However, as argued by Boeri and Terrell (2002) and Boeri (2001), it is difficult to associate this persistence pattern with the flexibility of labor markets. This is because the traditional factors used to explain maladjustment such as unions, minimum wages, and employment protection legislation are either weak or effectively not implemented. For these authors, non-employment benefits acting as wage floors may have discouraged labor reallocation creating strong disincentive effects.

Theoretical models of multiple equilibria in transition labor markets have been developed by Aghion and Blanchard (1994), Garibaldi and Brixiova (1998) and Boeri (2001) amongst others. Aghion and Blanchard (1994) develop a model where, depending on agents' expectations, the transition economy could end up in a high unemployment equilibrium. In Boeri (2001), multiple equilibria can arise due to microeconomic lock-in effects owing to skill specificity of workers together with the search disincentives generated by non-employment benefits in the formal and informal sectors.[6] This pattern would generate the appearance of long duration spells of unemployment and regime shifts in aggregate unemployment. In many of these models the timing of reforms determine the unemployment equilibrium (high or low) towards which the economy fluctuates. Note that this high persistence will arise even in effectively highly non-regulated labor markets such as those in CEECs. These models point to Hysteresis in unemployment. However, this mechanism substantially differs from traditional models of persistence – such as Blanchard and Summers (1987) – based on insider-outsider effects or human capital loss.

[6] Garibaldi and Brixiova (1998) use a similar argument using a labor market transition with matching theoretical framework.

If these theoretical models are correct, we should expect either high levels of persistence in unemployment dynamics or frequent unemployment equilibrium changes on the face of shocks as those experienced by Eastern European countries. The first hypothesis has been traditionally tested on OECD countries by applying unit-roots tests to unemployment series as in Brunello (1990), Song and Wu (1997), Arestis and Mariscal (1999) and Leon-Ledesma (2002). The second hypothesis has been tested in Bianchi and Zoega (1998) and Jaeger and Parkinson (1994) to a set of OECD countries. Surprisingly, however, little effort has been done in studying aggregate unemployment dynamics in transition economies beyond mere descriptive analysis. We try to fill part of this gap by analyzing the persistence patterns of aggregate unemployment in Eastern Europe.

Let us formalize our framework. Consider the following *AR (K)* process for the unemployment rate (*y*),

$$y_t = \alpha + \sum_{k=1}^{K} A_k y_{t-k} + \varepsilon_t, \qquad \varepsilon_t \sim i.i.d\left(0, \sigma_\varepsilon^2\right) \tag{1}$$

Here, the "natural", mean or equilibrium rate to which unemployment reverts over time is $\bar{y} = \dfrac{\alpha}{1 - \sum_k A_k}$ assuming $\sum_k A_k < 1$. Nevertheless, as long as $\sum_k A_k < 1$ (and there are no intercept shifts, i.e., $\alpha = \alpha \, \forall t$) unemployment will be mean reverting. However, if $\sum_k A_k = 1$, unemployment follows a random walk and displays path-dependence (or pure *Hysteresis*[7]). Thus, shocks ε_t have permanent effects.[8] This is a particular cause of concern for transition countries since (as in our previous discussion) it is not unreasonable to suppose that they have been hit by a relatively high number of shocks (increased openness to trade, price liberalization, privitizations and the removal of subsidies, the decay of previous trading partners and appearance of

[7] As opposed to *partial* Hysteresis, $\sum_k A_k \approx 1 \Big| \sum_k A_k < 1$, Layard *et al.*(1991).

[8] Note that, for the purposes of our analysis, these can be supply, demand or nominal shocks. It is beyond the scope of this paper to identify the relevant shocks. Since we rely on univariate analysis, our intention is to describe the dynamic behaviour of unemployment facing "a shock."

new ones, etc). Testing for unit roots for the presence of pure linear Hysteresis provides an upper bound test of the hypothesis, given that this is an extreme case of path-dependence where any shock, large or small, matters. Given that unemployment rates are bounded, unemployment should be stationary for longer time spans. Hence, Hysteresis as a unit root should not necessarily be understood as a 'true' description of the underlying data generating process of unemployment but as a local approximation to it during a sample period. A less restrictive hypothesis considers Hysteresis as a process by which unemployment switches equilibria whenever sufficiently large shocks affect its actual value. That is, if only large shocks enter the long-run memory of the unemployment series because they generate changes in the 'natural' or equilibrium level of unemployment.

Conventional stationarity tests can verify the presence of such "unit roots". However, testing for non-stationarity (in our application) raises a number of non-trivial technical issues. First, we necessarily have a short span of data. Second, tests may have low power against precisely those structural breaks that we might expect to characterize the data (e.g., the α's and A's may be time varying)[9]. Third, if there are structural breaks, we must try to both date these and ensure that we distinguish them from normal business-cycle fluctuations. Finally, it is possible that unemployment takes – in contrast to equation (1) – some non-linear form. This paper systematically tries to overcome these difficulties to robustly identify persistence patterns in transition countries' unemployment. On the first point (small sample), we use (in addition to conventional tests) panel unit-root tests that exploit both the time-series and cross-sectional dimensions of the data. As regards structural break tests (second and third points), we use single-equation and panel structural-break tests as well Markov-Switching methods that endogenously search for and date structural breaks independent of normal cyclical fluctuations. Finally, our Markov-switching regressions control for any possible non-linearity in the unemployment process and allow for the analysis switching equilibrium unemployment as suggested both by theoretical models of labor markets in transition and, as already mentioned, by recent conceptualizations of unemployment Hysteresis.

[9] In our context, the most appealing form of break is an intercept break. This would be consistent with 'strutctural' explanations of the natural rate hypothesis. See Phelps (1994).

3. Testing for unit roots.

3.1. *Time-series tests.*

As mentioned earlier, a traditional testing procedure is to apply unit roots tests on the unemployment rate. The existence of a unit root would indeed imply that unemployment does not revert to its natural rate after a shock. Table 1 presents the four different unit root tests on the monthly, seasonally-adjusted unemployment series of our set of 12 transition economies plus the EU-15 aggregate[10]. Details on data sources and sample periods can be found in Appendix I. The tests carried out are the ADF test for the null of a unit root, the Kwiatkowski *et al.* (1992) LM test for the null of stationarity (KPSS) and the asymptotically most powerful DF-GLS tests of Elliott *et al.* (1997) (ERS) and Elliott (1999) for the null of a unit root.[11] We report the tests with and without a time trend, and also provide the estimated auto-regressive root for the ADF test together with the derived half-life for the shocks. Given that our data is monthly, it is not surprising to observe high roots implying a slow speed of reversion to the mean. The results indicate that, for the majority of the tests, we cannot reject the null of a unit root for most countries in the sample. The main exception is Bulgaria, where only the ERS DF-GLS test for the model with an intercept cannot reject the null of a unit root. The other three countries where the presence of a unit root is rejected by several tests are Poland, Hungary and Lithuania and, to a lesser extent, Romania. On the other hand, Croatia, Estonia, Slovenia, Slovakia, Russia and the EU, are shown to behave as unit root processes in most cases and, hence, have very large half-lives for the correction of shocks. For the Czech Republic most tests including a time trend also reject the null of a unit root.

Confidence intervals for the largest auto-regressive root of the ADF tests were also constructed following Stock (1991). The 90% confidence intervals are reported in

[10] We also performed our tests with an EU-12 aggregate with little change in our results. Details available.

[11] The main difference between the ERS and the Elliott (1999) tests is that the former assumes zero initial conditions for the process under both the null and alternative, while the latter draws the initial observation from its unconditional distribution under the alternative.

Table 2. Compatible with the previous results, only Romania and Bulgaria seem to lie within the unit interval. For the rest of the cases, the upper bound estimate of the largest root is higher than unity for at least one case. Note however, that for countries such as Hungary, Czech Republic and Lithuania, the lower bound is sometimes close to 0.6, implying a very fast adjustment to shocks with around 1 month half-life. The 0.0% confidence interval does not, in general, coincide with the point estimate in Table 1. This is because, as argued by Stock (1991), the local-to-unity distribution of the point estimate of the auto-regressive root is skewed and depends on nuisance parameters. Another aspect of relevance is that the confidence intervals are, with a few exceptions, reasonably tight given our short sample and number of observations.

HERE TABLE 1

HERE TABLE 2

The tests presented in Table 1, however, suffer from two important problems that could substantially reduce their reliability. First, as pointed out by Perron (1989), in the presence of a structural change, we could erroneously be favoring the existence of a unit root when the process is in fact stationary with a change of mean or trend. The second problem is the low power of these tests especially when the sample is small. Although we are dealing with series of around 120 observations, our sample period of about 10 years might reduce the power of our tests and, hence, bias the results towards the acceptance of the null of a unit root. We will attempt to deal with the latter shortcoming later when making use of panel unit roots tests.[12]

In order to illustrate the possible instability of the ADF regressions and the existence of structural breaks, we obtained recursive Chow stability tests of the auto-regressive form of the ADF test, *AR(p)*, with *p* being the maximum lag chosen for the unit root tests. Figure 2 reports the results. It is easy to see that, for several countries, the Chow test exceeds its 5% critical value for several observations (especially for the Czech Republic, Slovakia, Estonia, Russia and, to a lesser extent, Croatia). These results are

[12] We also checked for the possibility of an asymmetric adjustment of unemployment in expansion and slow-down periods by fitting a momentum threshold auto-regressive model (M-TAR) to our data. The results did not show significant asymmetries in unemployment dynamics except for the possible case of the Czech Republic, which showed a higher persistence in periods of unemployment reduction. This, however, did not change our previous conclusions about unit roots in this country. Details available.

not surprising, given that the structural change process suffered by these economies must be reflected on its labor market outcomes.

HERE FIGURE 2

In order to control for the presence of structural breaks on the ADF regressions, we carried out the Perron (1997) unit root test with endogenous search for structural change, based on the following ADF regression for the time series y_t:

$$y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t , \qquad (2)$$

where $DU_t = 1$ ($t > T_b$) 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, and k is the lag augmentation for correction of residual auto-correlation. This is Perron's (1989) 'innovational outlier model' that implies a change in the mean but not the slope of the ADF regression. This is the most likely case to occur in unemployment series because of changes in the 'natural' rate. The test for a unit root is performed using the t-statistic for the null hypothesis that $\alpha = 1$. The optimal search for the break date is carried out using two methods. The first finds T_b as the value that minimizes the t-statistic for testing $\alpha = 1$. In the second, T_b is chosen to maximize the absolute value of the t-statistic associated with the change in the intercept θ.[13] As is standard in structural break tests, we have limited the search of the break date for both methods excluding the first and last 10% sample observations.

HERE TABLE 3

Table 3 reports the results of the Perron (1997) test. We report the break date (T_b), the t-statistic for $\alpha = 1$ and the estimated auto-regressive root for both break search methods. The t-statistics are compared with the critical values for T = 100 provided by Perron (1997). For 11 out of 13 economies tested both methods gave the same break

[13] We chose this method instead of minimizing the t-statistic on θ to avoid imposing a priori assumptions on the sign of the change.

date (or very similar in the case of Slovakia).[14] The results for the unit roots test indicate that we can now reject the null of non-stationarity for 6 countries by at least one of the methods. The speed of adjustment is now substantially faster in all cases as reflected in lower values of the estimated root. Of our sample, only Poland gets close to the EU aggregate in terms of the calculated half lives. For some countries like Hungary or Russia, the half-life becomes close to 3 months. Thus, once we have controlled for structural breaks, Hysteresis in our set of transition countries appears to be less plausible.

3.2. *Panel tests.*

As mentioned earlier, because of our relatively short sample, traditional unit roots tests may suffer a lack of power. To solve this, several authors have proposed the use of panel unit roots tests that exploit both the time-series and cross-sectional dimensions of the data.[15] Several tests have been proposed to check whether the panel series have a unit root. Here we apply three of them; two of which – Im *et al.* (2002) and Chang (2002) – rely on panel versions of ADF regressions whilst the third, Sarno and Taylor (1998), is based on Johansen's Likelihood Ratio test for cointegration in a VAR.

The Im *et al.* (2002)(IPS) test is based on the ADF regression:[16]

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta y_{i,t-j} + \xi_{i,t} \text{ ,} \qquad (3)$$

where i = 1, 2, …, N, and t = 1, 2, …, T. IPS test the null of non-stationarity ($\rho_i = 0$ $\forall i$) against the alternatives H_A: $\rho_i < 0$, i=1, 2, …, N_1, $\rho_i = 0$, $i = N_1 + 1, N_1 + 2, \ldots, N$. Note that the IPS test does not assume that all cross-sectional units converge towards

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[14] An interesting feature is that the two break methods tend to give more breaks in the second half of the 90s. This is compatible with labor market research in CEECs that emphasizes the deterioration of unemployment around 1997-1999 and, as will be evident in section 4, with our results from Markov Switching regressions.

[15] See Baltagi and Kao (2000) for an overview.

[16] For simplicity we will ignore deterministic trends in the explanation of the tests.

the equilibrium value at the same speed under the alternative, i.e. $\rho_1 = \rho_2 = \ldots = \rho_N < 0$, and thus is a less restrictive test than previous panel tests such as Levin and Lin (1992). The IPS test is based on the standardized *t*-bar statistic:

$$\Gamma_t = \frac{\sqrt{N}[\bar{t}_{NT} - \mu]}{\sqrt{\nu}} \sim N(0,1) \tag{4}$$

where $\bar{t}_{NT}$ is the average of the N cross-section ADF(p_i) *t*-statistics. μ and ν are, respectively, the mean and variance of the average ADF(p_i) statistic under the null, tabulated by Im *et al.* (2002) for different *Ts* and lag orders of the ADF. Im *et al.* (2002) also show that under the null of a unit root Γ_t converges to a N(0,1) as $N/T \rightarrow k$ (k is any finite positive constant).

One of the problems of the IPS test is that it assumes that the different cross sections are distributed independently. One way to avoid this problem, as suggested by Im *et al.* (2002) is to subtract cross-sectional averages from the individual series. This, however, does not allow for more general forms of dependency. The test proposed by Chang (2002) tries to overcome this problem by using a nonlinear IV estimation of the individual ADF regressions using as instruments nonlinear transformations of the lagged levels. The standardized sum of individual IV *t*-ratios has a limit normal distribution. Here we used the following Instrument Generating Function as in Chang (2002):

$$F(y_{i,t-1}) = y_{i,t-1} e^{-c_i |y_{i,t-1}|} \tag{5}$$

Where c_i is proportional to the sample standard error of the first difference of y_{it}:

$$c_i = K T_i^{-1/2} s(\Delta y_{it}), \tag{6}$$

where K is a constant fixed at 5 as recommended in Chang (2002) for time dimensions larger than 25 observations.

The Sarno-Taylor (1998) test (ST) takes a different route based on Johansen's (1992) Maximum Likelihood method to determine the number of common trends in a system of unit root variables. We can represent a *k*-dimensional vector auto-regressive (VAR) process of *p*-th order as:

$$\Delta Y_t = \mu + \Theta_1 \Delta Y_{t-1} + \ldots + \Theta_{p-1} \Delta Y_{t-p+1} + \Pi Y_{t-p} + \varepsilon_t \,, \qquad (7)$$

where μ is a $(k \times 1)$ matrix of constants, Y_t is a $(k \times 1)$ random vector of time series, Θ_i are $(p \times p)$ matrices of parameters, and Π is a $(k \times k)$ matrix of parameters whose rank contains information about long-run relationships between the variables in the VAR. If Π has full rank ($rank(\Pi) = k$) then all variables in the system are stationary. Hence, the ST test has as a null H_o: $rank(\Pi) < k$ and as alternative H_A: $rank(\Pi) = k$, which can be implemented using Johansen's (1992) Likelihood Ratio test. That is, it tests if one or more of the system variables is non-stationary against the alternative that all the variables are stationary. This is a more restrictive test than LL and IPS because it will reject the null *if at least one* of the series in the panel has a unit root.

The results from these three tests are presented in Table 4. We have carried out the test for three different groups. The first one contains all the transition economies. The second excludes Bulgaria, since this was the only economy in which nearly all time series tests rejected non-stationarity. Given that the null of the IPS and Chang (2002) tests is that *all* cross-sections have a unit root, they would clearly be affected by the inclusion of a stationary series. The third group contains all economies except Estonia, Latvia and Lithuania for which data only starts in 1994M1 and shortens the time-series component of the panel. As the IPS test loses power if there is substantial cross-sectional correlation in the panel, we also applied the tests to each series adjusted by subtracting the cross-sectional average. Overall, the results show that the unemployment series are stationary. Only the null of the test on unadjusted data and an intercept for the 11 countries group seems to indicate the presence of a unit root. The IPS test rejects the null in all cases but two and the Chang (2002) and TS test, the most restrictive, reject the null in all cases.[17]

[17] We compared the ST test to a $\chi^2(1)$ adjusted by a factor $T/(T - p \cdot k)$ as recommended by Sarno and Taylor (1998) to account for finite sample bias.

HERE TABLE 4

Finally, given the evidence on the likely importance of structural breaks, we combine panel unit roots tests with endogenous break search tests by using the Murray and Papell (2000) (MP) test. This test can be considered a combination of the Perron (1997) test and the Levin and Lin (1992) panel unit roots test. It assumes that the auto-regressive coefficient of all cross-sections is the same and that the date of break is also common between cross-sections. It allows for heterogeneity in the intercept and the lag augmentation of the ADF equation and accounts for cross-sectional correlation by estimating the panel by SUR methods. The break date is found as the one that minimizes the *t*-statistic for testing $\alpha = 1$ as in Perron's (1997) Method I. The results of this test are reported in Table 5. We chose the lag augmentation of each unit to be the same found for the ADF tests and, again, applied the test for the 3 groups considered in previous panel tests. The results, again, strongly reject the null of non-stationarity at the 99% level, and the auto-regressive roots are found to be of the order of 0.9. The dating of the break in the second half of the 1990s is not surprising, given the rapid deterioration of unemployment in many countries during this period, and the results obtained using the Perron (1997) test for individual countries.

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The overall picture shows that unemployment dynamics in transition economies during the last decade have not been characterized by a linearly hysteretic behavior. Once we control for the impact of structural change, the low power of time series tests, or both, we can reject a random walk in unemployment. Although there is still a high level of persistence in countries such as Croatia, Slovenia, Estonia or Latvia, on average, it is lower than that for the EU aggregate. The lock-in effects that theory models describe at the micro level do not appear to have derived from a random walk behavior.

4. Markov Switching Analysis.

Despite the relevance of the unit root analysis in the previous section, our analysis has been confined to testing for a strong version of Hysteresis that assumes that every shock will have permanent effects on the level of unemployment. However, following Amable *et al.* (1995) and Røed (1997), Hysteresis is best associated with the existence of multiple equilibria in unemployment dynamics as mentioned earlier. Importantly, our previous analysis of unit roots makes a number of assumptions, which we might now like to relax or reconsider. First, the unit-root, structural-break tests used – being essentially supremum tests – might be considered as biased towards finding a break even if one does not exist. Secondly, this is particularly problematic if the data (as we might suppose) is characterized by both business-cycle fluctuations and possibly structural breaks. Third, the break implicit in the analysis of unit roots of the previous section assumes that either unemployment reverts to a constant level or to an average characterized by sudden changes. Unemployment, however, is more likely to adapt smoothly to an infrequently changing average or 'natural' level of unemployment. That is, that unemployment is subject to changes in regimes due to microeconomic factors such as those described in theoretical models of labor markets in transition economies. For these reasons, and to add an extra layer of robustness to our previous results, we move on to analyze the persistence profile of unemployment using Markov switching regressions. This will allow us not only to test for Hysteresis with a changing average level of unemployment, but also to analyze the frequency of regime changes and the behavior of unemployment in each of these regimes. Another advantage of the technique is that it accounts for non-linearities in the trend unemployment function accruing not only from structural breaks but also from normal business cycle fluctuations.

The Markov switching model for m regimes (or states) – where $m \in [2, \infty)$ – can be represented by equation (8) where y_t (the unemployment rate) is regressed on an intercept (v) and auto-regression of length I with a residual (u) with variance Σ – all of which might be state dependent (denoted by s_t):

$$y_t = \upsilon(s_t) + \sum_{k=1}^{K} A_k(s_t) y_{t-k} + u_t\,,\; s_t = 1,\ldots,m\,,\; u_t \sim \mathrm{N}\left(0, \Sigma(s_t)\right) \tag{8}$$

Although most popularly found in business-cycle applications, Markov models have also been applied to employment phenomena: Eaton (1970), Schager (1987), Ciecka *et al.* (1995), Bianchi and Zoega (1998), Akram (1998). The notable characteristic of such models is the assumption that the unobservable realization of the state, s_t, is governed by a discrete-time, discrete-state Markov stochastic process. This is defined by the transition probabilities:

$$\Pr(s_{t+1} = j \mid s_t = i) = \rho_{ij}, \sum_{j=1}^{m} \rho_{ij} = 1, \forall i, j = 1,...,m \tag{9}$$

Thus, s_t follows a Markov process with the transition probabilities matrix, P:

$$P = \begin{bmatrix} \rho_{11} & \rho_{12} & \cdots & \rho_{1m} \\ \rho_{21} & \rho_{22} & \cdots & \rho_{2m} \\ \vdots & \vdots & \ddots & \vdots \\ \rho_{m1} & \rho_{m2} & \cdots & \rho_{mm} \end{bmatrix} \tag{10}$$

Defining the number of states (m) is among the most difficult aspects of Markov-switching (see discussions in, inter alia, Garcia, 1998, Garcia and Perron, 1996). Here, we use two state identification methods (see Appendix II). First, using kernel density estimation methods, we use the number of modes in the density function as an indicator for the number of states. Second, we use Likelihood criteria.

Equation (8) represents the general case and allows all components (means\intercepts, auto-regression and variances) to be state dependent. Though we found a mixture of all these elements, notably, the data suggested intercept over mean dependency. This is also straightforward to motivate; since we are dealing with a labor market – rather than, say, a spot financial market – we consider it more plausible that the mean should slowly and gradually adjust to a new level (from one transition to another) rather than as an immediate mean jump.[18]

[18] The data also strongly rejected the Markov-switching-in-mean model; details available.

We estimate using the EM algorithm (Hamilton, 1990) and follow Hamilton's (1989) classification method by assigning an individual observation x_t to the state m with the highest "smoothed" probability: $m^* = \arg\max_{m} \Pr\left(s_t = m \middle| x_T, x_{T-1}, ..., x_1\right)$. To derive standard errors for the estimates of equation (8) we bootstrap.[19] In addition, we provide Likelihood Ratio tests to confirm state-dependent variances.[20] Not unsurprisingly given our sample coverage, we find essentially only two states in the data. Exceptions are Poland, Romania and Croatia, for whom we model one [21]. Table 6 presents country estimates of the summed auto-regressive parameter *A(L)*, transition probabilities ρ_{ii}, proportion measures ξ_i, state error variances σ_i^2 and state means, $\bar{u}_i$. First of all, we see that – excluding Latvia and second-state Czech Republic – having controlled for different states (i.e., business-cycle fluctuations and/or structural breaks) all countries have stationary processes for their unemployment rates. [22] As before, the country with the highest level of persistence – and thus the slowest adjustment to a shock – is the EU15 with Russia and second-state Czech Republic relatively close by. In many cases, we can see that there has been a rather unbalanced state dependence. For example, most countries (excepting Bulgaria, the Czech Republic) spend around two thirds of their time in one state. States are also highly persistent: once in state *i* the probability of remaining there is around 0.8 and upwards. (The exception to this appears to be Hungary where there has been considerably more switching between states). Notably, in those cases where there exists state-dependent variance, high unemployment generally accords with high variance.

TABLE 6 HERE

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[19] Davidson (2001), Ehrmann *et al.* (2001).

[20] Ang and Bekaert (1998) provide Monte Carlo evidence on the power of LR tests in a Markov context.

[21] Thus, for these single-state cases, the stationarity tests already reported remain the measure of their hysteretic properties.

[22] Ang and Bekaert (1998) suggest that state-switching models with a unit (or near unit) root process in one of the states remain stationary as long as there is at least one strictly stationary state. It is precisely this property which allows state-switching models to capture the near unit-root persistence in unemployment data.

The country with the highest effective level of persistence statistically is the Czech Republic. This is because we cannot reject the null of a unit root in state two.[23] However, state one, where the Czech economy spends nearly 47% of the observations, presents a very low auto-regressive root. These two states identify the rapid process of labor market deterioration suffered by the Czech economy during the late 1990s. Another important result is that, for the majority of cases, and in line with previous unit root tests, we reject the null of a random walk behavior. Furthermore, unemployment mean rates across states appear relatively well separated (e.g., the Czech Republic has an average unemployment rate of 3.0% in state 1 and 6.8% in state 2). With the exception of the EU-15 and Russia – where the spread is marginal – our results lend strong support to the notion of multiple equilibria.

As we know, states captured by Markov-switching methods can be both business-cycle fluctuations (recessions and expansions) as well as structural breaks. A concept related to the latter is an *absorbing* state: a state which, once entered, is never exited. One might also consider locally- (or semi-) absorbing states, whereby the process resides in one state for a "sufficiently" long time. An absorbing state occurs when the Markov chain becomes reducible – i.e., from $t = t_1,..., t^{upper}$, we have p_{ii} = 1. The absorbing state would set $t^{upper} = T$ and a semi-absorbing state might define some sufficient distance $t^{upper} - t_1$. This definition of structural break as a (semi) permanent change of state is related to the existence of Hysteresis defined as a system with multiple equilibrium. Once unemployment suffers a rapid increase or decrease, it tends to stay in the new state (lock-in). This is probably a closer definition of the Hysteresis arising in theoretical models of labor markets in transition economies.

Examining the smoothed probability for each country (Figure 3), most countries have spent unusually long periods in one state.[24] The Czech Republic spent the early

[23] The Markov Switching model also suggests a unit root in the case of Latvia – although this derives more from the imprecision of the standard errors than a high point value.

[24] The comparison between absorbing states and time-series structural breaks is by no means exact. The former essentially verify a break when there is complete degeneracy (i.e., there is no further exist from state) whilst the latter may be more commonly thought to register a structural break during the transition away from a previous state; that is to say as $\rho_{ii} \to 1$. Despite this, comparing time-series (as earlier examined) and Markov-switching structural break detection methods may be a useful cross-checking exercise.

sample (up until around 1996) in the first (low unemployment) state followed by a transition to a high-unemployment one. We might therefore tentatively suggest a structural break around 1997-98 (as Table 3 suggests). The same can be said for Lithuania with a likely break in 1998-1999 (although Table 3 picks up the earlier break of 1997:1). Slovakia appears to have spent most of its time in the (low-unemployment) first state but from 1998 onwards appears to head permanently into a high unemployment state (Table 3 tends to pick up the break around late 1992). Latvia appears to have entered a high unemployment state in the immediate aftermath of the Russian crisis (as also indicated by Table 3) but recovered by around mid 2000.

HERE FIGURE 3

These results, hence, show that, for several transition countries, unemployment follows a multiple equilibrium pattern. More concretely, the shocks that affected unemployment during the last years of the past decade seem to have moved these economies towards a high-unemployment equilibrium. Hysteresis, although not manifested, in general, as a linear random walk process behavior, seems to take the form of multiple equilibria especially for countries such as the Czech Republic, Lithuania and Slovakia. This lock-in pattern behavior is supportive of recent models of transition in labor markets such as Boeri (2001).

5. Conclusions.

In this paper we have undertaken a systematic analysis of the dynamic behavior of unemployment in transition economies benchmarked against the EU-15 aggregate. We tested for the existence of hysteretic features in their labor markets making use of both unit roots tests and Markov switching regressions. Our findings show that, in general, we can reject the unit root hypothesis after controlling for structural changes and business cycle effects, but we can observe the presence of a high and low unemployment equilibria towards which the economy fluctuates after sufficiently large shocks.

When compared with the behavior of aggregate unemployment dynamics in the EU during the past decade, we can see that transition countries' unemployment shows a faster speed of adjustment and larger changes in unemployment equilibria across regimes. Exception to this pattern would be Croatia, whose unemployment behavior is best described as a linear unit root or near-unit root process, and Latvia, where unemployment seems to follow a random walk and also regime changes. For the rest of the countries the level of persistence is relatively low, which is consistent with the existence of less regulated labor markets. Moreover, for several countries we find that changes in unemployment regimes tend to be highly absorbing. That is, once unemployment shifts towards a new regime, it tends to remain locked into it or, at least, remain there for a long period of time. Notable cases of lock-in in a high unemployment regime during the final years of the 1990s are the Czech Republic, Lithuania and Slovakia. We can thus conclude that unemployment dynamics in Eastern Europe are characterized by a switching unemployment equilibrium towards which actual unemployment reverts quicker than in the EU. This pattern is supportive of recent theoretical models of the labor market in transition countries.

These results have important implications for labor market reforms, as well as macro-stabilization policy in the CEECs. Standard progressive macroeconomic stabilization policies do not appear to have a long lasting impact on unemployment, at least not longer than what the EU experience reveals. However, deeper reforms of both labor and goods markets – which might constitute "large" shocks – that are likely to continue in the CEECs should take into account the possibility of having a long lasting impact on the equilibrium level of unemployment.

Appendix I: Data Sources

Country Series	Data Source	Sample
Poland	Central Statistical Office of Poland.	Jan. 1991 – June 2001
Romania	National Commission for Statistics.	Dec. 1991 – Apr. 2001
Slovenia	Central Bank of Slovenia.	Jan. 1992 – May 2001
Croatia	Statistical Office of Croatia.	Jan. 1992 – May 2001
Hungary	Central Statistical Office of Hungary.	May 1991 – Aug. 2001
Bulgaria	WIIW, Eastern Europe Economy.	Jan. 1991 – June 2001
Czech Republic	WIIW, Eastern Europe Economy.	Jan. 1991 – May 2001
Slovak Republic	Slovak Statistical Office.	Jan. 1991 – May 2001
Estonia	OECD Main Economic Indicators.	May 1993 – May 2001
Latvia	Latvijas Statistiskas/Monthly Bulletin	Jan. 1994 – May 2001
Lithuania	Lithuanian Department of Statistics	Jan. 1994 – May 2001
Russia	Goskomstat/Russian Economic Trends.	Jan. 1992 – Mar. 2002.
EU-15	EUROSTAT	Jan. 1991 – Dec. 2000

Note:
WIIW = Wiener Institut für Internationalen Wirtschaftsvergleich.

Appendix II—Identifying Markov-Switching State Number

Defining the number of states (m) is among the most difficult aspects of Markov-switching. Often, for instance, a state number is imposed on the data, or the prior implicit in the exercise (such as a two-state business-cycle model) is used. Here, however, we directly test for state number. We use two approaches. First using Kernel density estimation methods. Second, Likelihood criteria. Given the relatively short span of data, these tests remain indicative. Where there is conflict between the tests, we favor the lower state case.

Bootstrap Multi-Modality tests and Density Estimation Techniques

The numbers of modes (or peaks, bumps) that underlie the data are taken to indicate the number of states relevant for the Markov-Switching representation. Multi-Modality techniques have been used substantially in Statistical fields but also in Economics, particularly in the income-distribution and income-convergence literature (e.g., Quah, 1997).[25] The multi-modality tests used are based on kernel density estimation (Silverman, 1986) and bootstrapping (Efron and Tibshirani, 1993). For $m=m^*$ (where m^* indicates the number of modes in the data) the frequency distribution of a series, x_t, can be expressed $f(x)=\sum_{j=1}^{m^*} p_j g_j\left(x;\mu(s_j),\sigma^2(s_j)\right)$ where the p_j's are mixture proportions, $\sum p_j = 1$ and g_j are uni-modal densities with first and second moments, μ,σ^2. By kernel methods $f(x)$ is estimated non-parametrically by: $f(x)=\frac{\sum_{t=1}^{T}\Gamma(\bullet)}{Th}$ where Γ is the Gaussian kernel, T denotes the sample size and h ($h>0$) is the bandwidth parameter. Silverman (1986) defines the critical bandwidth, $h_{crit}(m)$ as the smallest possible h producing a density with less than or equal to m modes; thus for $h_{crit}(m)>h^*$, the density has greater than or equal to $m+1$ modes. Specifically, if the series has n modes, then $h_{crit}(n-1)$ should be "large" because substantial smoothing is required to generate an n-mode density. Thus, this technique

[25] This state-selection method in economics has been applied and discussed in, for instance, Akram, (1998), Bianchi and Zoega (1998), Fernandes (1998), Pittau and Zelli (2001).

forms a natural hypothesis-testing framework. Although how large is "large" is defined by the bootstrap: a sample is taken of the original series (with replacement) and transformed to have the same first and second moments. *P-values* for $h_{crit}(m)$ are obtained by generating a large number of samples from $f_{crit}(m)$ and relying on the proportion of the samples for which $h^{*}_{crit}(m) > h_{crit}(m)$, where $h^{*}_{crit}(m)$ is the smallest h for which a density with m modes is produced from the bootstrapped equivalent values of x (x^*). The *p-values* (which allows conventional inference) are generated from $\delta^{-1}\sum_{j=1}^{\delta} I_{m,j}$, where I is a dummy variable defining whether $f_{h_{crit(m)}}(x^*)$ has greater than m modes ($\Rightarrow$I=1) etc and δ is the number of bootstrap replications (we set δ = 10,000). However, it is well known – Silverman (1983), Izenmen and Sommer (1988), Hall and York (2001) – that the Silverman test tends to underestimate true rejection regions and accordingly probability values higher than conventional ones (e.g. 0.025, 0.05, 0.10) are typically used. Consequently while reporting *p-values* in table 1A below, we tend to rely on the Silverman h^* statistic – a robust estimate of which is determined by $0.9AT^{-0.2}$, where A is the minimum of the standard deviation of the series and its interquartile range divided by 1.34.

Table 1A gives both the h^* statistic and the *p-values*. As we have said the precise cut-off point for deciding on rejection-region *p-values* is controversial and thus we focus on and crosscheck with the Silverman statistic. The decision rule is to search until $h^* > h_{crit\_m}$ then choose $m^* = m$. Figure 1A graphs the resulting densities using the critical bandwidths consistent with the country series' uni, bi- or tri-modality. To illustrate, the Silverman test suggests 2 states s in the Czech Republic data and indeed the Czech density graph visually identifies those 2 modes well.

Complexity Penalized Likelihood Criteria

Following Psaradakis and Spagnolo (2002), an estimate of the true state number (m^*) is given using the following penalized complexity penalty function:

$$\hat{m}^* = \arg\max_{m\in[1,m^{Max}]}\{\ln L(\theta, X) - C_T \dim(\Theta)\}$$

where $L(\theta, X)$ and θ are the likelihood function and parameter vector respectively for the model estimated on the data *X* with parameter dimension, $\dim(\Theta)$, and sample size *T*. The choices of the constant C_T include 1 (Akaike Information criteria) and $\frac{\ln T}{2}$ (Bayesian Information). Table 2A shows our results (for max$\{m^*\}$ = 4).

Results

As can be seen, agreement between the Kernel density and likelihood results is close. In 9 out of the 13 country cases there is exact agreement between the AIC, BC and the Silverman statistic. In the case of the Slovak Republic, the BC, Silverman and AIC suggest respectively 2,3 and 4. We choose 2 states on the grounds of degrees of freedom, the marginal rejection of the Silverman statistic in choosing 3 over 2 states and the (unreported) poor performance of a three-state regression. On Russia, the Silverman and BC suggest 2 states as against 4 for the AIC. Again, we choose the more plausible 2-state case. Similarly, for the EU-15, we choose the lower two-state model identified by the Silverman statistic. The most notable case, however, is Croatia. All state-identification methods suggest 2 states. For Croatia it has not proved feasible to model a Markov process (details available). This may be related to our earlier finding of non-stationarity. More generally, the data would not appear mean reverting over any sample – the data seems to be trending downwards and then upwards either side of the mid-1990s – which, as we know, militates against the fitting of Markov process (e.g., Ang and Bekaert, 1998).

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FIGURES

Figure 1. Unemployment Rates.

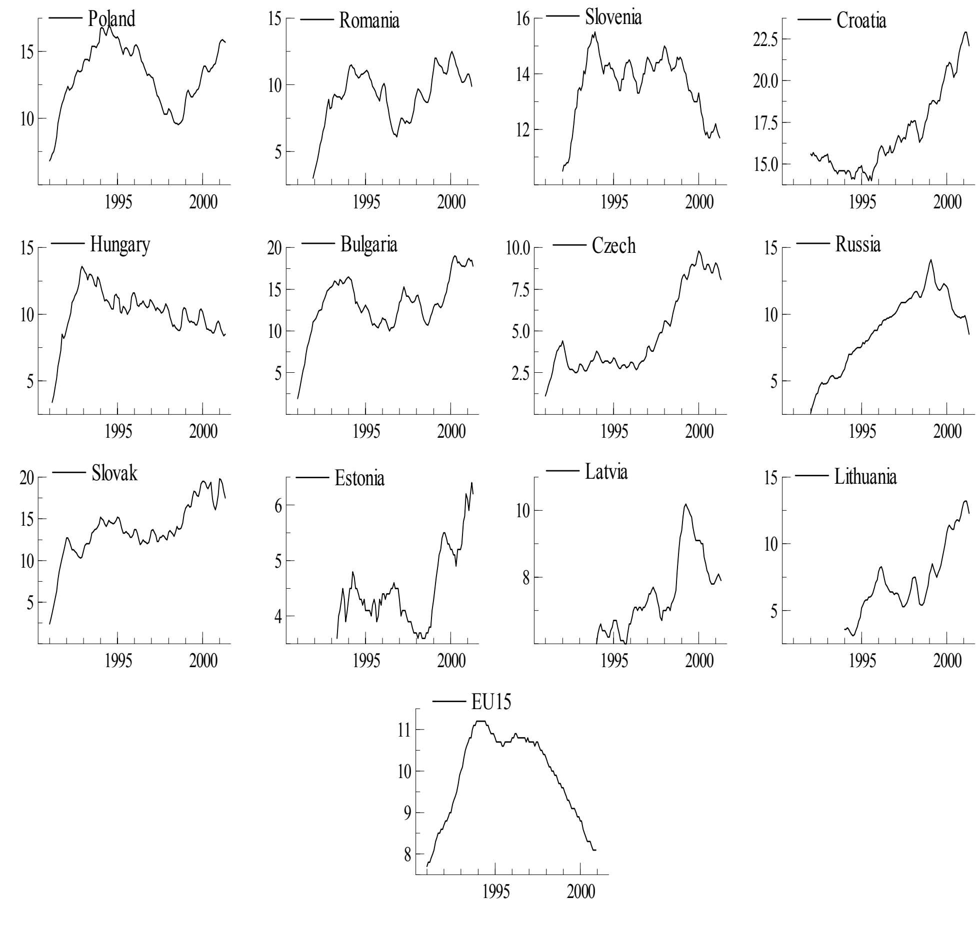


Figure 2. Break point recursive Chow instability test

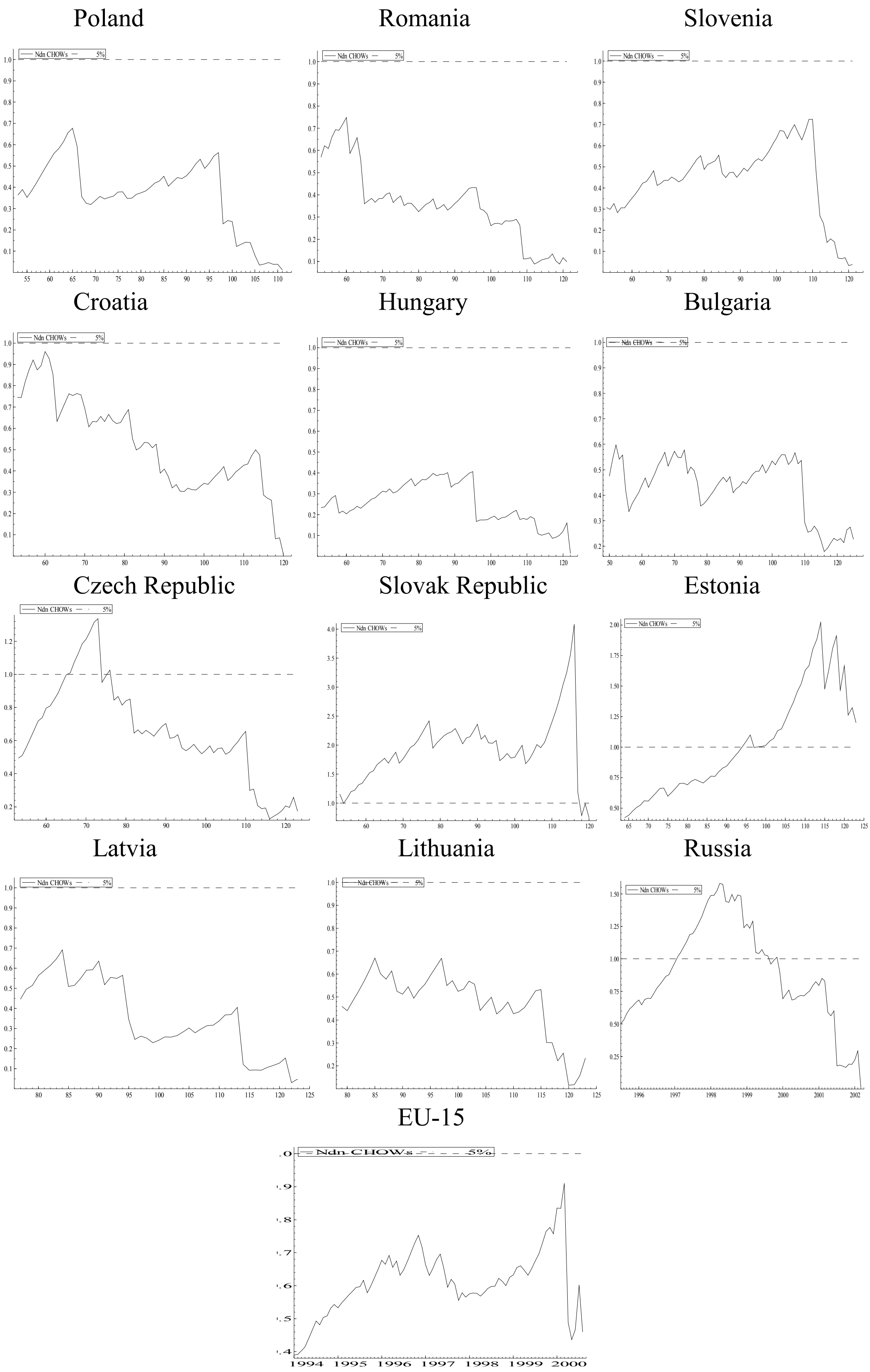
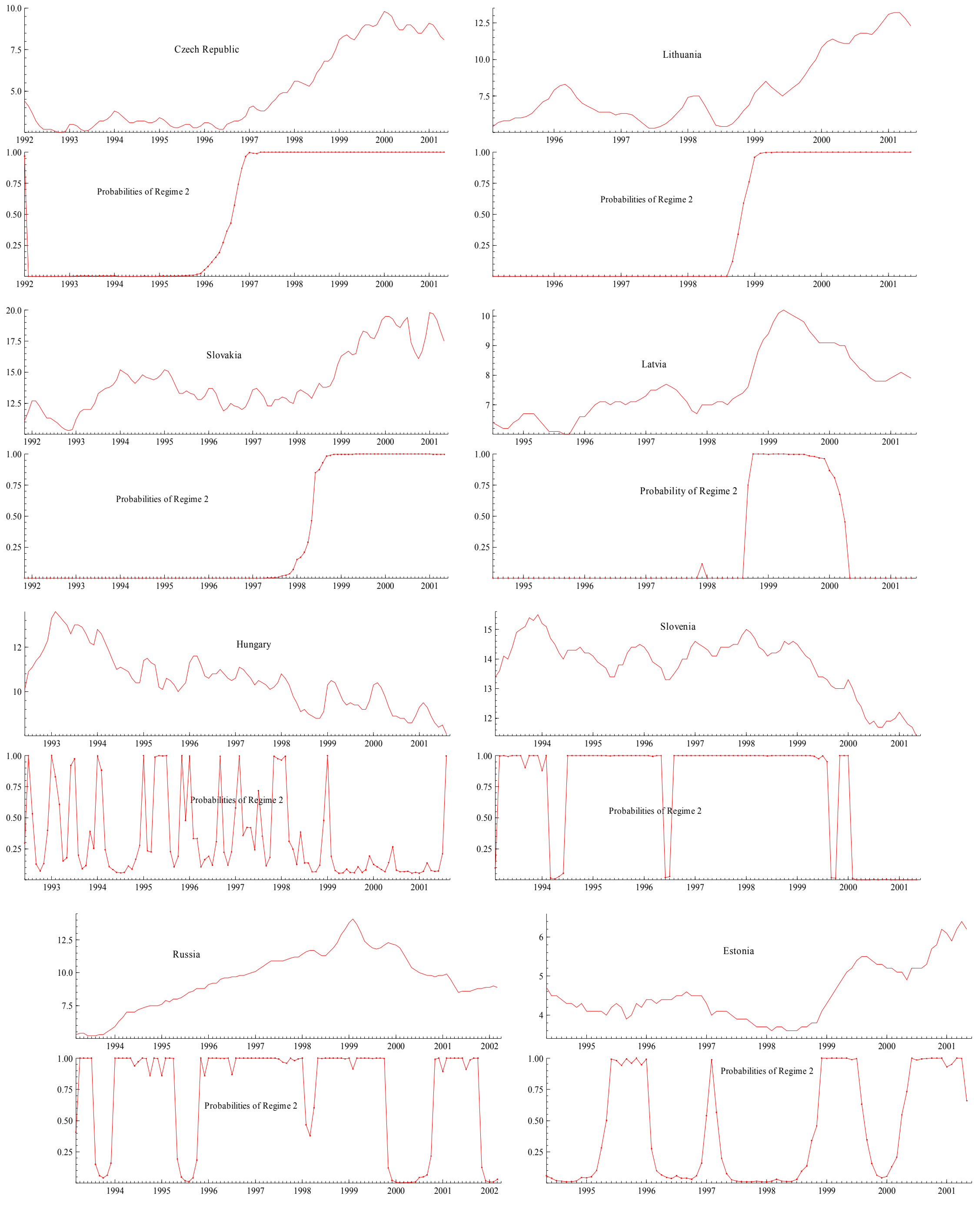
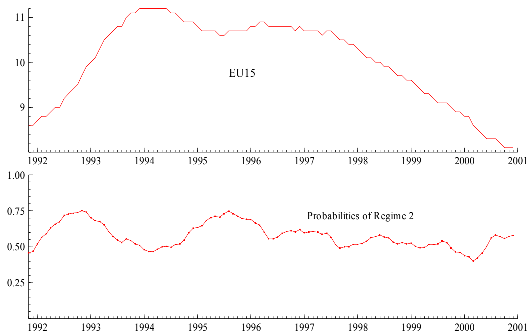


Figure 3. Unemployment Rates and Smoothed Probabilities.




EU15
Probabilities of Regime 2
11
10
9
1.00
0.75
0.50
0.25
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1993
1994
1995
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1997
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2000
2001

Figure 1A. Density Graphs

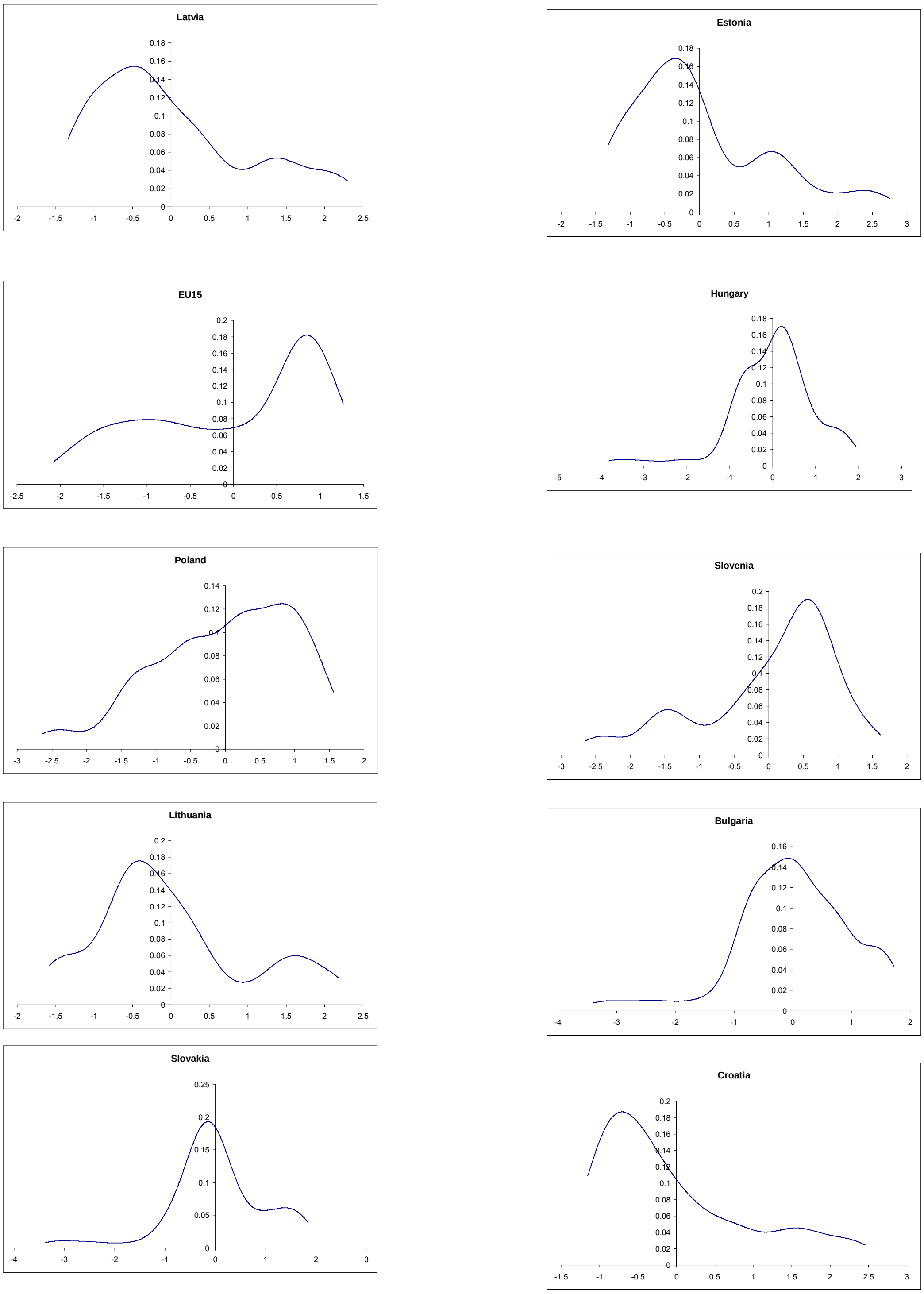
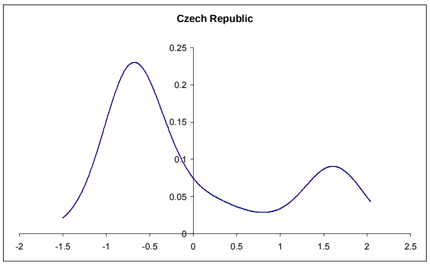
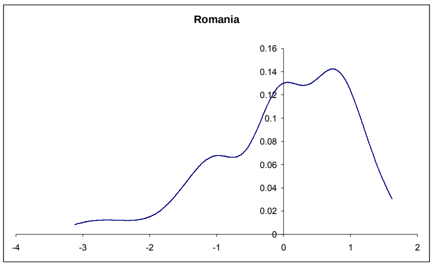


Figure 1A (Cont.): Density Graphs





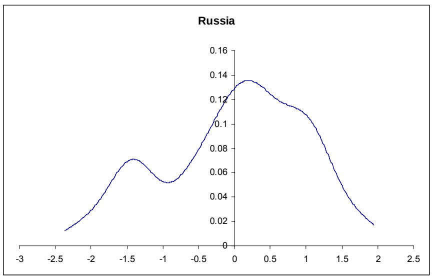


Table 1. Unit root tests

		ADF				KPSS		ERS		Elliott (1999)	
		Intercept model		Trend and intercept model		Intercept model	Trend model	Intercept model	Trend model	Intercept model	Trend model
Country series	No of Lags	T-ratio	Estimated root {half life}	T-ratio	Estimated root {half life}	η_{μ}	η_{τ}	DF-GLS	DF-GLS	$DF\text{-}GLS_u$	$DF\text{-}GLS_u$
Poland	12	-2.848	0.973 {25.32}	-2.363	0.977 {29.79}	0.1518*	0.1580	-2.432	-3.206*	-2.801*	-3.214*
Romania	10	-3.386*	0.945 {15.05}	-3.648*	0.947 {13.80}	0.3210*	0.1116	-2.120	-1.895	-1.927	-2.070
Slovenia	11	-2.324	0.944 {12.03}	-2.752	0.928 {9.28}	0.1904*	0.1820	-1.291	-1.367	-1.649	-1.367
Croatia	12	0.235	1.019 {n.a.}	-2.348	0.941 {11.40}	0.8212	0.2373	-0.694	-1.345	-0.611	-1.821
Hungary	12	-2.118	0.950 {13.51}	-4.527*	0.841 {4.00}	0.2112*	0.1398*	-2.260	-2.921*	-3.059*	-2.901
Bulgaria	12	-3.523*	0.947 {12.73}	-4.379*	0.928 {9.28}	0.3790*	0.1080*	-2.812*	-3.195*	-3.464*	-3.846*
Czech Rep	12	-0.897	0.995 {138.28}	-4.697*	0.949 {13.24}	0.8479	0.2218	-3.118*	-4.172*	-0.912	-4.109*
Slovak Rep	12	-0.782	0.988 {57.41}	-2.054	0.947 {12.72}	0.7864	0.1201*	-1.893	-2.226	-0.501	-2.251
Estonia	7	-0.932	0.971 {23.55}	-1.616	0.966 {20.04}	0.5896	0.2313	-1.892	-1.892	-1.191	-1.885
Latvia	2	-1.638	0.986 {49.16}	-2.283	0.950 {13.55}	2.0493	0.1866	-2.277	-2.385	-1.620	-2.364
Lithuania	12	-2.001	0.974 {26.31}	-4.673*	0.897 {6.38}	0.6341	0.1324*	-2.533	-4.188*	-2.033	-4.112*
Russia	12	-1.606	0.989 {62.67}	0.479	1.014 {n.a.}	0.8217	0.2013	-1.187	-1.312	-1.533	-1.461
EU-15	6	-1.687	0.989 {62.67}	-2.0411	0.986 {49.16}	0.4285*	0.4215	-1.576	-1.716	-1.898	-1.891

Notes to Table 1:
ADF is the Augmented Dickey-Fuller test, KPSS is the Kwiatowski *et al.* (1992) test and ERS is the Elliott *et al.* (1997) test.

* Indicates rejection of the null of a unit root at the 5% level for the ADF, DF-GLS and DF-GLS$_u$ tests respectively and not rejection of the null of stationarity for the KPSS test at the 95% level.

The half life was calculated as - [Ln(2)/Ln(ρ)], where ρ is the auto-regressive root of unemployment in the ADF test, and is expressed in months.

The number of lags was chosen using a General-to-Specific criterion.

Table 2. 90% Confidence intervals for the auto-regressive root in ADF test.

Country series	Intercept Model		Trend Model	
	90% interval	0.0% interval {half life}	90% interval	0.0% interval {half life}
Poland	(0.812, 1.001)	0.895 {6.25}	(0.879, 1.028)	0.969 {22.01}
Romania	(0.737, 0.935)	0.833 {3.79}	(0.726, 0.956)	0.832 {3.77}
Slovenia	(0.867, 1,015)	0.940 {11.20}	(0.838, 1.023)	0.928 {9.28}
Croatia	(1.004, 1.034)	1.016 { n.a. }	(0.880, 1.028)	0.971 {23.55}
Hungary	(0.886, 1.019)	0.956 {15.40}	(0.592, 0.841)	0.710 {2.02}
Bulgaria	(0.729, 0.929)	0.827 {3.65}	(0.624, 0.869)	0.741 {2.31}
Czech Rep	(0.972, 1.030)	1.010 { n.a. }	(0.579, 0.829)	0.698 {1.93}
Slovak Rep	(0.978, 1.031)	1.011 { n.a. }	(0.908, 1.031)	1.013 { n.a. }
Estonia	(0.970, 1.030)	1.010 { n.a. }	(0.946, 1.034)	1.017 { n.a. }
Latvia	(0.926, 1.025)	0.996 {172.94}	(0.887, 1.029)	0.979 {32.66}
Lithuania	(0.896, 1.020)	0.965 {19.45}	(0.583, 0.832)	0.701 {1.95}
Russia	(0.942, 1.027)	1.005 { n.a. }	(1.017, 1.041)	1.024 { n.a. }
EU-15	(0.919, 1.025)	0.992 {86.30}	(0.905, 1.033)	1.014 { n.a. }

Notes: Confidence intervals calculated using Stock's (1991) method. Half-lives calculated as in Table 1.

Table 3. Perron (1997) tests on unemployment series

Country Series		Break Search Model I			Break Search Model II		
	Period	Break date	T-ratio	Estimated root {half life}	Break date	T-ratio	Estimated root {half life}
Poland	91:01 – 01:05	1996:03	-4.372	0.946 {13.49}	1996:03	-4.372	0.946 {13.49}
Romania	91:12 – 01:04	1993:11	-5.032**	0.915 {8.80}	1996:01	-4.537	0.880 {6.42}
Slovenia	91:12 – 01:04	1999:06	-4.578	0.835 {4.84}	1999:06	-4.578	0.835 {4.84}
Croatia	91:01 – 01:05	1999:01	-4.017	0.880 {6.42}	1999:01	-4.017	0.880 {6.42}
Hungary	91:05 – 01:08	1992:11	-5.947*	0.680 {2.80}	1992:11	-5.947*	0.680 {2.80}
Bulgaria	91:01 – 01:05	1999:02	-5.28*	0.907 {8.10}	1999:02	-5.28*	0.907 {8.10}
Czech Republic	91:01 – 01:05	1998:04	-6.781*	0.891 {7.01}	1992:07	-3.789	0.961 {18.42}
Slovak Republic	91:01 – 01:05	1992:11	-3.580	0.908 {8.18}	1992:12	-3.477	0.909 {8.26}
Estonia	93:05 – 01:05	2000:05	-3.431	0.912 {8.52}	1998:10	-2.343	0.911 {8.43}
Latvia	94:01 – 01:05	1998:04	-4.195	0.850 {5.26}	1998:04	-4.195	0.850 {5.26}
Lithuania	93:01 – 01:05	1997:01	-8.153*	0.791 {3.96}	1997:01	-8.153*	0.791 {3.96}
Russia	92:01 - 01:04	1998:08	-6.379*	0.732 {3.22}	1998:08	-6.379*	0.732 {3.22}
EU-15	91:01 – 00:12	1992:05	-3.800	0.967 {21.66}	1992:05	-3.800	0.967 {21.66}

Notes: * and ** indicate rejection of the null at the 5% and 10% significance level respectively. Critical values from Perron (1997) Table 1. Half life calculated as in Table 1.

Table 4. Panel unit-roots tests results on CEECs.

		Im, Pesaran and Shin (1997)	Chang (2002)	Taylor-Sarno (1998)
Unadjusted 12 countries	Intercept	-0.962	-1.636**	11.490*
	Trend	-3.809*	-2.795*	14.750*
Adjusted 12 countries	Intercept	-2.202*	-	-
	Trend	-3.381*	-	-
Unadjusted 11 countries	Intercept	-0.461	-1.404**	11.955*
	Trend	-3.327*	-2.663*	13.122*
Adjusted 11 countries	Intercept	-1.698*	-	-
	Trend	-2.738*	-	-
Unadjusted 9 countries	Intercept	-4.245*	-1.711*	7.891*
	Trend	-3.931*	-2.754*	10.147*
Adjusted 9 countries	Intercept	-3.510*	-	-
	Trend	-2.855*	-	-

Notes:
Estimation periods for 12 and 11 countries are 1994M1-2001M4.
Estimation period for 9 countries is 1992M1-2001M4.
12 countries includes all the database.
11 countries excludes Bulgaria, which was shown not to have a unit root in ADF tests.
9 countries exclude Estonia, Latvia and Lithuania due to their shorter time series.
For the IPS and Chang tests we used the lags chosen from the ADF tests. We used 4 lags for TS test because of lack of degrees of freedom to estimate a larger lag structure in a sensible way. However, the TS test results are not sensible to the inclusion of up to 6 lags. The critical values for the TS test ($\chi^2(1)$) have been adjusted by a factor $T/(T - p \cdot k)$ as recommended by Sarno and Taylor (1998), where p is the lag of the VAR and k is the number of countries. Results for the TS test with adjusted data are not possible to obtain, because the adjustment method would obviously lead to multicollinearity in the VAR.
Results for the Chang test with adjusted data are not reported given that the test controls for cross-sectional dependence.
* and ** indicate rejection of the null of a unit root at the 5% and 10% level respectively.

Table 5. Murray-Papell break panel unit root test.

	Period	Break date	T-ratio	Estimated root {half-life}
12 countries	1994:01 – 2001:04	1998:09	-12.115*	0.899 {6.51}
11 countries	1994:01 – 2001:04	1998:05	-10.065*	0.920 {8.31}
9 countries	1991:01 – 2001:04	1996:03	-9.978*	0.939 {11.01}

Notes:
11countries includes all countries in the sample except Bulgaria.
9 countries excludes Estonia, Latvia and Lithuania.
The critical values are given by Murray and Papell (2000). For N = 10 and T=100, the 1% critical value is –8.658 and for T = 50 and N = 10 it is –9.056. * indicates rejection of the null at the 5% level.

Table 6. Markov Switching Results

Country Series	$A_1(L)$ (1)	$A_2(L)$	$H_o:\sigma_1^2=\sigma_2^2$ (2)	$\sigma^2_{u,1}$	$\sigma^2_{u,2}$	$\bar{u}_1$ (3)	$\bar{u}_2$	ρ_{11} (4)	ρ_{22}	ξ_1 (5)	ξ_2
Slovenia	0.82259 (0.03674)		9.6762 [0.0019]	0.0038 (0.007985)	0.0183 (0.007314)	12.215	14.196	0.8290	0.9416	0.2546	0.7454
Hungary	0.9393 (0.0028)		42.5960 [0.0000]	0.0092 (0.0053)	0.1345 (0.0066)	9.0721	11.742	0.7665	0.5604	0.6531	0.3469
Bulgaria	0.90152 (0.0141)		1.3212 [0.2504]	0.0724 (0.0018)		11.580	15.783	0.9424	0.9474	0.4776	0.5224
Czech Republic	0.6902 (0.1522)	0.9785 (0.1406)	2.2744 [0.1315]	0.0098 (0.0009)		3.0277	6.7857	0.9796	0.9821	0.4669	0.5331
Slovak Republic	0.92926 (0.02635)		16.672 [0.0000]	0.0600 (0.027763)	0.2631 (0.028043)	13.314	17.902	0.9873	0.9999	0.6774	0.3226
Estonia	0.93685 (0.01528)		7.2956 [0.0069]	0.0043 (0.00168)	0.0356 (0.00223)	3.8029	5.9310	0.9054	0.8817	0.5558	0.4442
Latvia	0.93546 (0.099006)	0.82406 (0.3227)	12.240 [0.0005]	0.0126 (0.003415)	0.0027 (0.000365)	7.1191	9.3035	0.9822	0.9351	0.7851	0.2149
Lithuania	0.8488 (0.008531)	0.90615 (0.00795)	4.3888 [0.0362]	0.0228 (0.000111)	0.0111 (0.000114)	6.2260	8.9376	0.9784	1.000	0.5947	0.4053
Russia	0.9793 (0.0036)		27.306 [0.0000]	0.0003 (0.0005)	0.0315 (0.0005)	9.9165	9.5087	0.8083	0.9254	0.2801	0.7199
EU-15	0.9872 (0.0047)		0.0006 [0.9805]	0.0029 (0.0002)		9.1294	10.112	0.9521	0.8553	0.7513	0.2487

Notes: (1) $A_i(L)=\sum_k A_k \middle| m=i$. (2) LR tests distributed as a $\chi^2(1)$. (3) Mean unemployment rate in i^{th} state: $\bar{u}_i$. (4) $\rho_{ii}=\Pr(s_{t+1}=i \mid s_t=i)$. (5) Proportion of time in i^{th} state: $\xi_i=\frac{n_{t,i}}{T}$, $\sum_{i=1}^{m}\xi_i=1$ where T=sample size, $n_{t,i}$=number of observations in i^{th} state and, $\sum_t\sum_{i=1}^{m}n_{t,i}=T$ Standard errors in ()'s, p-values in []'s. In each case, there are as many intercepts as states.

Table 1A. Multi-Modality Tests

Country Series	Critical Bandwidths			p-values				
	h_{crit_1}	h_{crit_2}	h_{crit_3}	*m=1*	*m=2*	*m=3*	*m**	*h**
Poland	0.325	0.275	0.265	0.486	0.313	0.050	1	0.343
Romania	0.325	0.310	0.130	0.596	0.151	0.995	1	0.350
Slovenia	0.445	0.315	0.160	0.082	0.061	0.590	2	0.350
Croatia	0.385	0.235	0.215	0.350	0.436	0.155	2	0.350
Hungary	0.555	0.300	0.260	0.204	0.374	0.212	2	0.343
Bulgaria	0.405	0.320	0.275	0.525	0.160	0.275	2	0.343
Czech Republic	0.379	0.175	0.170	0.460	0.446	0.049	2	0.343
Slovak Republic	0.540	0.355	0.195	0.087	0.095	0.719	3	0.297
Estonia	0.395	0.355	0.175	0.193	0.021	0.588	2	0.360
Latvia	0.420	0.255	0.195	0.199	0.287	0.325	2	0.367
Lithuania	0.535	0.270	0.160	0.009	0.092	0.591	2	0.356
Russia	0.420	0.255	0.165	0.063	0.2744	0.6388	2	0.343
EU-15	0.525	0.205	0.185	0.053	0.631	0.217	2	0.345

Table 2A. Complexity Penalized Likelihood Criteria

Country Series	Complex Penalized Likelihood	
	AIC	**BIC**
	$m^* \in [1,4]$	
Poland	1	1
Romania	1	1
Slovenia	2	2
Croatia	2	2
Hungary	2	2
Bulgaria	2	2
Czech Republic	2	2
Slovak Republic	4	2
Estonia	2	2
Latvia	2	2
Lithuania	2	2
Russia	4	2
EU-15	3	3