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Unemployment dynamics and NAIRU estimates for CEECs : A univariate approach*

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Abstract

In this paper we test for the hysteresis versus the natural rate hypothesis on the unemployment rates of the EU new members using unit root tests that account for the presence of level shifts. As a by product, the analysis proceeds to the estimation of a NAIRU measure from a univariate point of view. The paper also focuses on the precision of these NAIRU estimates studying the two sources of inaccuracy that derive from the break points estimation and the autoregressive parameters estimation. The results point to the existence of up to four structural breaks in the transition countries NAIRU that can be associated with institutional changes implementing market-oriented reforms. Moreover, the degree of persistence in unemployment varies dramatically among the individual countries depending on the stage reached in the transition process.

JEL Classification: C22, C23, E24

Keywords: Hysteresis, NAIRU, Structural Breaks, Unit Root, CEEC

Resum

En aquest article es contrasta la histèresi front la hipòtesi de la taxa natural d'atur pels nous membres de la UE emprant contrastos d'arrel unitària que tenen en compte la presència de canvis estructurals. Adicionalment, el treball duu a terme l'estimació d'una mesura de la NAIRU emprant tècniques univariants. Al llarg del treball es farà èmfasi en la precisió d'aquestes estimacions de la NAIRU analitzant dues possibles fonts d'imprecisió derivades de l'estimació dels punts de trencament i de l'estimació dels paràmetres autoregressius. Els resultats evidencien l'existència de fins a quatre canvis estructurals que afecten el valor de la NAIRU pels països en transició, canvis estructurals que poden ser associats a canvis institucionals ocasionats per la implementació de reformes orientades cap a una economia de mercat. Finalment, el treball mostra com el grau de persistència en la desocupació varia radicalment entre els països considerats segons el nivell assolit en el procés de transició.

1. Introduction

Enlargement is one of the most important challenges in the European Union (EU) agenda. Although it is not the first time that the EU is admitting countries with lower levels of economic development than the existing members, three characteristics make the present enlargement unique. The first difference concerns their nature as planned socialist economies that started a process of transition to private market economies; the second difference is the income gap between new and old members and third, but not less important, the potential for labor migration after the accession. Due to income differentials and geographical proximity to core EU countries, the current enlargement could result in higher migration flows than in previous occasions.

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 Commission (2001). 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 (Boeri, Burda and Köllö (1998) and Huber (1999)).

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 cushion 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

[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.

structural rate of unemployment and its shifting nature may be crucial for policy makers (De Grauwe (1975)).

From an economic policy point of view, the approach that explains the inflation-unemployment relationship dates back to the 60's and 70's. The concept of structural or "natural" rate of unemployment was first introduced by Friedman (1968) and Phelps (1968). 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. An important element of this approach is the concept of a nonaccelerating inflation rate of unemployment, or NAIRU, defined for the first time by Modigliani and Papademos (1975) as an unemployment rate (or range of unemployment rates) that produces a stable rate of inflation serving as the empirical counterpart of the unobserved "natural rate of unemployment" –see Franz (2003). Supply shocks and inflation expectations are also commonly felt to be important determinants of inflation, and the NAIRU may change over time depending on the structure of the economy and government policy. Although this concept is no longer very popular among academic economists, both, the NAIRU and the theory of the inflation-unemployment relationship on which it is based still receives much attention from press and among economic policy makers (Espinosa-Vega and Russell (1997)). In particular, the OECD has been developing a research program to provide accurate measures of the NAIRU. Within this framework, the OECD defines three different NAIRU concepts, which vary by the timeframe they cover (Richardson, Boone, Giorno, Meacci, Rae and Turner (2000)). The first one, or "short-term NAIRU", is the rate of unemployment required to stabilize the inflation rate at its current level in the next period. This short-term concept is more volatile by definition as it requires a level of NAIRU that will provide an inflationary offset to any impact from short-term supply shocks, expectations, and possible speed limit/persistence effects. This paper focuses on the short-term NAIRU which is potentially the most appropriate to help policymakers to assess inflationary developments in the short-term (Estrella and Mishkin (1998) and King (1999)).

A second NAIRU concept is the medium-term one. This NAIRU is the equilibrium rate that unemployment converges to in the absence of temporary supply shocks and the dynamic adjustment of inflation to previous shocks is completed. Finally, a third definition is the long-term equilibrium unemployment rate which is equivalent to the long-term steady state, as the NAIRU has fully adjusted to all long and short-term supply and policy shocks.

As Szeto and Guy (2004) state, of these three definitions only the first two can be estimated; the short-term NAIRU can be directly estimated and the medium-term NAIRU can be estimated by controlling adequately for short term influences (when possible).

In this paper we analyze the persistence pattern of aggregate unemployment for the Acceding Countries into the European Union, testing for hysteresis in their unemployment rates, and computing a simple measure of short-term NAIRU. This goal is important for several reasons. First, gaining insight about the degree of persistence helps to assess the effects of labor market reforms as well as macro-stabilization policies undertaken by this countries; second, it can be used as a tool to ascertain asymmetries in labor markets across the new EU members and between them and the rest of the EU in the future, especially on the way to a prospective enlarged monetary union; third, the NAIRU may still have an indicative role for policy makers.

From a methodological point of view, we contribute to the empirical literature in three respects. First, we take into account the small sample problem common to all empirical analyses of transition economies by applying the M unit root tests by Ng and Perron (2001) when testing for hysteresis. Secondly, we address the issue of structural changes typical of all post-communist countries in Central and Eastern Europe, as recently stressed by Fidrmuc and Tichit (2004). For this purpose, we use tests that allow us to endogenously determine the potentially multiple structural breaks. This is an important feature as previous studies on transition economies found important differences among the countries depending on the temporal pace of the reforms. Finally, we compute an improved measure of both point and confidence intervals of the NAIRU, based on the previous structural change analysis.

The plan of the paper is as follows. In Section 2 we offer an overview of the univariate approach adopted for measuring the NAIRU. Section 3 reports the results of the applied research focusing, firstly, on the hysteresis versus natural rate hypothesis testing and, secondly, on how precisely the NAIRU is estimated. We also compute the half lives associated with the obtained NAIRU as a measure of persistence. Finally, Section 4 presents the main conclusions.

2. The univariate approach to measure the NAIRU: A brief overview

According to McAdam and McMorrow (1999) there are two broad modelling approaches to measure the NAIRU, namely the expectations-augmented Phillips curve approach, which singles out a series of labor markets variables as potential empirical determinants of the NAIRU, and the univariate method, where the time series properties of the macro variables are used to determine the NAIRU.

The former approach establishes the NAIRU at the point where a stable Phillips-curve relationship exists between the deviation of unemployment from the NAIRU and unexpected inflation. Within this context it is possible to distinguish two variants, the single equation inflation approach and the multiple equation wage-price model (*e.g.* the Bargaining model). This approach corresponds more closely to a measure of the long-run equilibrium rate of unemployment. Although these structural models can provide a strong theoretical framework, they do not allow specific estimates of the NAIRU and are subject to several criticisms[2] (Richardson et al. (2000)). Firstly, there is no agreement about the appropriate theoretical model (e.g. long-run effects of changes in real interest rates, taxation or productivity growth on real wages and equilibrium unemployment); second, there is a lack of consensus on specification issues (number and identity of explanatory variables and its linear-non linear and/or symmetric-asymmetric nature)[3]; third, there is a statistical identification problem typical of systems of equations where the determined variable enters as explanatory variable at the same time, creating the well known endogeneity problems. Finally, there are many difficulties to obtain reliable data about institutional variables (unemployment benefits, employment protection, degree of unionization...) which have an increasing importance to determining structural unemployment (Blanchard and Wolfers (2000), Belot and Van Ours (2001)).

Therefore, as an alternative to multivariate approaches that look for the level and determinants of the NAIRU, the univariate approach is the one that we apply in this paper. We think that the above mentioned flaws of the structural approach are especially acute in the particular case of the transition countries. Note that the degrees of freedom in the estimations would be seriously compromised by the scarcity and reliability of the time-series information available.

The univariate approach is essentially statistical, with the underlying assumption being that unemployment reverts to its mean or natural rate over time. Thus, previous to any empirical research aiming at measuring the NAIRU, the

[2] Refinements of the empirical specifications led Gordon (1997) to summarize them in terms of three factors: expectations/intertia, the unemployment gap, and supply shocks, being important to distinguish between temporary (changes in real import prices or changes in real oil prices) and long-lasting supply shocks (level of real interest rates, tax wedge, demographics...). The former are expected to revert to zero over the long horizon and are particularly important for monetary policy, while the latter may permanently alter the NAIRU, so that inflation will vary until unemployment adjusts.

[3] This point is especially relevant when the objective is to apply the same specification accross many countries (as it would be the case in the present paper).

hypothesis of hysteresis in unemployment should be discarded. The economic literature has experienced an increasing debate on testing the natural rate paradigm versus the hysteresis hypothesis of the unemployment rate. Most of the empirical evidence has focused on developed countries, mainly because these countries suffered more intensively the effects of the oil crises, which meant an increase in both the level and the persistence of the observed unemployment rate.

The *natural rate theory* assumes that the unemployment rate evolves around an equilibrium level (or natural rate) describing stationary fluctuations. This equilibrium level is set depending on fundamentals of the economy such as labor productivity, technology, world real interest rates and the real exchange rate, to mention a few –see Friedman (1968) and Phelps (1967, 1968, 1994). In contrast, the *hysteresis hypothesis* implies that the shocks affecting the unemployment rate have a permanent effect on the variable so that it never attains its equilibrium level. In this extreme case, unemployment is not anchored by structural variables, but will instead reflect the cumulative effect of all past shocks to the economy, including those to demand. Thus, this hypothesis can be formulated in terms of unit root testing, with perfect hysteresis associated with the non-rejection of a unit root.

The decomposition of the unemployment rate into "cyclical" and "structural" unemployment is important for analytical and policy reasons. The structural unemployment rate ($\bar{U}$) is defined as the equilibrium rate of unemployment, or the rate at which there is no tendency emanating from the labor market or inflation to either increase or decrease, that is why it is also called the NAIRU.

In practice, the distinction between structural and cyclical unemployment is complicated due to the existence of either unemployment persistence effects or "hysteresis". The concept of hysteresis can be easily explained using a conventional Phillips curve:

$$\pi_t = \beta(L)\pi_t - a(U_t - \bar{U}_t), \quad (1)$$

where π_t is the rate of inflation, $\beta(L)$ the lag operator, U_t the actual rate of unemployment and $\bar{U}_t$ the structural rate.

In this basic version, $\bar{U}_t$ is identical to the NAIRU but if we introduce the possibility of *full hysteresis*, this equation changes to:

$$\pi_t = \beta(L)\pi_t - a[U_t - \beta(L)U_{t-1}], \quad (2)$$

and $\bar{U}_t$ will no longer be uniquely defined, while the NAIRU will equal the lagged unemployment rate. As a consequence, there will be a permanent trade-off between inflation and unemployment, but the long-term equilibrium rate of unemployment will not be permanent anymore and will change as a result of cumulative past shocks to the economy –see Blanchard and Summers (1987).

An intermediate variant, including both a stable $\bar{U}_t$ and *partial hysteresis* (or "*persistence*") effects, can also be specified as follows:

$$\pi_t = \beta(L)\pi_t - a(U_t - \bar{U}_t) - b[U_t - \beta(L)U_{t-1}]. \quad (3)$$

Introducing a persistence element into a Phillips curve results in a deviation between $\bar{U}_t$ and the NAIRU. The latter will now be a linear combination of $\bar{U}_t$ and the lagged actual unemployment rate:

$$NAIRU_t = \frac{a}{(a+b)}\bar{U}_t + \frac{b}{(a+b)}\beta(L)U_t. \quad (4)$$

This opens the possibility that inflation increases even if $U_t > \bar{U}_t$, which will happen if the actual unemployment rate drops too fast. This is why the partial hysteresis effect is also referred to as "speed-limit" effect. $\bar{U}_t$ and the NAIRU will be equal only in long-term equilibrium, when $\bar{U}_t$ equals $\beta(L)U_t$. It is therefore justified to refer to $\bar{U}_t$ and the NAIRU as the "long-run" and "short-run" structural (or equilibrium) rates of unemployment, respectively. The distinction between $\bar{U}_t$ and the NAIRU is rarely made in practice in the context of policy formulation, reflecting the difficulties of measuring either concept with precision (IMF (2001)). However, it is crucial to bear in mind this distinction when interpreting the results of the present paper, where the estimation has focused on the short-term NAIRU. Note that if there is evidence of inertia, this can delay the adjustment of the short-term NAIRU, keeping it closer to the actual unemployment rate than to the long-run NAIRU for prolonged periods. Again, this implies the possibility of "speed-limit" effects, so that if the actual unemployment rate is well above the long-run NAIRU, it will only be possible to close the gap slowly without increasing inflation. Divergences between short and medium-long NAIRU can be great when persistence effects are strong. Departure between both variables can be substantial and prolonged due to the relatively weak effects of the unemployment gap on inflation. At the same time stimulative policy actions can be postponed to avoid short term inflationary "speed limit" effects.

The validity of the natural rate hypothesis is based on two assumptions – see Friedman (1968) and Phelps (1967, 1968). Firstly, the uniqueness of the equilibrium level of unemployment and its independence from monetary variables in the steady-state. Secondly, actual unemployment tends to return to the natural rate given that expectations tend to correct themselves sooner or later. The rejection of either of these two assumptions would imply refuting the natural rate hypothesis. While initially the absence of theories explaining the determination of the natural rate meant that in practice it was taken to be constant, subsequent developments have attempted to explain the reasons behind changes across economies and over time. Among the structural factors influencing the natural rate are the productivity level and its growth, energy prices, international trade, union power, and normative traditions –for a discussion of these issues see Bianchi and Zoega (1998). Note that the experience of the acceding countries in transition from communism to capitalism suggests important changes in the fundamentals of the economies, affecting among others the labor market. This casts serious doubts on the empirical validity of the natural rate theory and therefore, an alternative specification that accounts for slow adjustments toward a shifting natural rate could be more realistic.

Most of the empirical literature focused on testing for hysteresis or persistence in unemployment has been based on analyzing the sum of the coefficients in the autoregressive process representing the unemployment rate. Values close to but lower than one were associated with partial hysteresis, that is, strong persistence. In contrast, perfect or pure hysteresis exists when the sum of the coefficients is equal to one. It should be stressed that only in the latter case is the natural rate hypothesis violated, given that even in cases of strong persistence unemployment slowly converges to the natural rate. However, it is sensible to state that in such a case the difference is negligible –see Bianchi and Zoega (1998). However, the fact that the effects of the level of unemployment on inflation are relatively weak and slow-acting, instead of non-existent, makes an important difference with relevant implications for the relationship between actual unemployment, the NAIRU and the associated unemployment gap.

In this context, unit root tests have been widely applied to unemployment rates –see Blanchard and Summers (1987), Decressin and Fatás (1995) and Bianchi and Zoega (1998)–, whose findings favor, in general, the hypothesis of hysteresis versus the natural rate, as the null was not rejected. Mostly the evidence supports hysteresis in the EU economies and the natural rate in the US and Nordic countries –see Papell, Murray and Ghiblawi (2000). However, those conclusions are based on unit root tests that under the alternative assume a constant, unique, natural rate of unemployment. Unfortunately, these test statistics are not robust to the presence of structural breaks –see Perron (1989) and Montañés and Reyes (1998). In order to overcome this limitation, new contributions have allowed for a more flexible specification of the deterministic component in unit root tests. For instance, Arestis and Biefang-Frisancho Mariscal (1999) and Papell et al. (2000) apply unit root tests that allow for structural breaks in the unemployment rates of samples of OECD countries. For the majority of the countries analyzed the null of pure hysteresis (unit root) is rejected in favor of the alternative of stationarity around a changing equilibrium rate. Papell et al. (2000) conclude that such a finding seems to be more in accordance with the structuralist theories of unemployment. The institutional characteristics of the Accession countries recommend to account for the presence of structural breaks when assessing the stochastic properties of the unemployment rates, provided that a misspecification error of the deterministic function would lead to infer wrong conclusions and, hence, to apply spurious economic policy advice.

Although the univariate approach is purely statistical (by contrast to the structural modelling approach) and offers indicators of trend unemployment that are consistently estimated across countries, they suffer from a number of practical drawbacks (Richardson et al. (2000)). First, the estimated indicators are often not very well correlated with inflation and are difficult to extrapolate even in the short-term. Second, they tend to be least reliable at the end of the sample, that is precisely the period of most interest for policy issues. This problem can be partially solved by adding forecasts at the end of the data sample. Third, as most of the filters behave like simple moving averages, they tend to perform poorly if a sudden change in the unemployment rate occurs. This problem has been tackled allowing for different means of the unemployment rate across the sample (Staiger, Stock and Watson (1997)). Finally, it is difficult to judge the degree of precision of the results. In order to overcome these problems, in this paper we use more sophisticated estimation techniques that help us to allow for structural changes and measure the statistical uncertainty surrounding the NAIRU estimates, calculating confidence intervals.

Moreover, recent contributions aiming at improving the estimation of the NAIRU using a univariate approach have emphasized its time-varying nature, modelling the variable either as a deterministic function of time or as an unobserved stochastic process, evaluating the uncertainty surrounding these estimates –see King, Stock and Watson (1995), Staiger et al. (1997) and Gordon (1997). This is the approach followed in this paper.

3. 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. Thus, in order to test for hysteresis in the unemployment rate we have applied the *M* unit root tests proposed by Ng and Perron (2001) that offer suitable properties in finite samples. Additionally, due to the institutional changes experienced by these countries during the period studied, we consider the existence of one or two structural breaks endogenously determined using the tests of Perron and Vogelsang (1992), and Montañés and Reyes (1997) as well as the tests by Lumsdaine and Papell (1997), respectively. In a second stage, following the strategy of Papell et al. (2000), we apply the procedure developed by Bai and Perron (1998) to the stationary series determined in the previous phase to obtain the number of structural breaks with some confidence intervals that are used to estimate the NAIRU in the transition countries. For this purpose, we apply the methodology developed by Staiger et al. (1997) which is based on univariate models.

The combination of these elements allows us to obtain both point and confidence intervals estimates of the NAIRU that offer a picture of the precision achieved in the calculations. This is a relevant question since it has been argued that the NAIRU has been traditionally measured quite imprecisely, casting doubts on the role that these estimates should play in discussions of monetary policy – see Staiger et al. (1997). As an additional test for hysteresis versus persistence, we compute the half lives associated with the obtained NAIRU values as well as modified Phillips Curves using the obtained short-run NAIRU estimates.

3.1 Hysteresis or natural rate of unemployment in the acceding countries?

In this Section we analyze the order of integration of the unemployment rates for all the countries acceding to the EU in May 2004, with the exception of Cyprus. 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, in order to increase the number of observations[4]. 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 León-Ledesma and McAdam (2004) to extend the database[5] that, at best, covers from 1991:1 to 2003:11[6]. These harmonized unemployment rates are depicted in Figure 1.

In a first step of the analysis, the hysteresis hypothesis is tested through the application of the unit root tests to each time series. In this Section we use

[4] 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.

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

[6] 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).

Table 1: M-class Unit root tests

	Z_α	MZ_α	MZ_t	MSB	P_T	MP_T	ADF
Czech Rep.	0.099	0.153	0.120	0.782	47.111	38.205	0.117
	(0.789)	(0.789)	(0.792)	(0.859)	(0.856)	(0.851)	(0.800)
Estonia	-2.543	-2.480	-1.094	0.441	12.059	9.776	-1.121
	(0.353)	(0.351)	(0.310)	(0.465)	(0.410)	(0.376)	(0.307)
Hungary	-0.736	-0.678	-0.512	0.755	38.739	29.508	-0.517
	(0.623)	(0.624)	(0.576)	(0.842)	(0.805)	(0.779)	(0.584)
Latvia	-3.035	-2.979	-1.220	0.410	10.106	8.224	-1.226
	(0.302)	(0.298)	(0.254)	(0.406)	(0.347)	(0.310)	(0.262)
Lithuania	-2.415	-2.338	-0.992	0.424	12.854	9.919	-1.002
	(0.374)	(0.373)	(0.364)	(0.436)	(0.437)	(0.386)	(0.371)
Malta	0.108	0.349	0.217	0.622	40.598	27.768	0.054
	(0.850)	(0.863)	(0.861)	(0.801)	(0.861)	(0.832)	(0.842)
Poland	0.203	0.269	0.209	0.779	49.897	38.932	0.182
	(0.808)	(0.810)	(0.815)	(0.856)	(0.869)	(0.855)	(0.817)
Slovakia	0.228	0.292	0.321	1.097	92.367	70.830	0.295
	(0.813)	(0.815)	(0.842)	(0.973)	(0.968)	(0.961)	(0.844)
Slovenia	-1.828	-1.769	-0.893	0.505	17.019	13.175	-0.923
	(0.440)	(0.440)	(0.402)	(0.568)	(0.534)	(0.489)	(0.397)

P-values in parentheses.

the so-called M-tests proposed in Ng and Perron (2001) and based on the GLS detrending procedure by Elliot, Rothenberg and Stock (1996) due to their better small sample performance in terms of empirical size and power. We also present in Table 1 the feasible point optimal test statistic, P_T, by Elliot et al. (1996) and the GLS detrended ADF test or ADF^{GLS}.

These are the test statistics that we have computed, in the first stage of the analysis, to test for the unit root hypothesis (hysteresis) in unemployment. The deterministic specification consists of a constant term ($p = 0$) that is consistent with the natural rate paradigm. In order to take into account the presence of autocorrelation we have set $k_{\max} = int\left(12\left(T/100\right)^{1/4}\right)$ using the $MAIC\left(k\right)$ criterion to choose the lag augmentation. Instead of using the asymptotic critical values in Ng and Perron (2001), we have applied the finite sample p-values from the estimated P-value functions in Carrion-i-Silvestre (2003). The results are reported in Table 1, which shows that there is strong evidence in favor of the unit root hypothesis for all countries –the p-values are in parentheses.

However, based on the visual inspection of the unemployment rates in Figure 1, we should consider the possible presence of structural breaks that may have affected the acceding countries unemployment rates. Thus, the natural rate might be experiencing a slowly transition between shifting natural rates, movements probably due, in turn, to changes in the fundamentals of the economies. It is well known that the presence of these breaks can provoke the spurious non-rejection of the null hypothesis in standard unit root tests–see Perron (1989) and Montañés and Reyes (1998). Thus, in order to account for this possible problem, we have computed a group of ADF-type unit root test allowing for structural breaks.

These tests can be specified using a general regression equation:

$$y_t = \mu + \beta t + \sum_{i=1}^{n} \theta_i DU_{i,t} + \sum_{i=1}^{n} \gamma_i DT_{i,t} + \sum_{i=1}^{n} d_i D\left(T_b\right)_{i,t} + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + \varepsilon_t, \tag{5}$$

with $DU_{i,t} = 1$ and $DT_{i,t} = t$ for $t > T_{b,i}$, 0 elsewhere; $D(T_b)_{i,t} = 1$ for $t = (T_{b,i} + 1)$ and 0 elsewhere, and where $T_{b,i}$ defines the i-th $(i = 1, ..., n)$ structural change. As before, testing for the null hypothesis of hysteresis in unemployment rate means that $\beta = 0$ and $\gamma_i = 0, \forall i = 1, ..., n$, in ((5)). From this general specification, several test statistics have been proposed, that differ mainly on the method that is applied to choose the break points and the formulation of the alternative hypothesis.

First, we deal with the situation in which there is just one structural break. One of the most used and popular unit root test statistic that takes into account the presence of a structural break is the one proposed by Perron and Vogelsang (1992). These authors suggest estimating the break date through the minimization of the sequence that is obtained after the ADF test is computed at all possible break dates. We denote this test as t_{α}^{*}, *i.e.* the ADF test that estimates the break date as

$T_b^* = \arg\min_{T_b \in (k+1,T)} t_{\hat{\alpha}}\left(T_b, k\right)$, where $t_{\hat{\alpha}}$ is the pseudo t-ratio for α in ((5))[7]. While Perron and Vogelsang (1992) deal with non-trending variables allowing for one structural break both under the null and alternative hypotheses –henceforth, we denote their specification as Model An– Zivot and Andrews (1992) derive the limit distribution of the test when it is applied to trending variables, but allowing for a structural break just under the alternative hypothesis[8]. The tests in Zivot and Andrews (1992) are for the joint null hypothesis of unit root without structural break and the alternative of stationarity with one structural break.

Lumsdaine and Papell (1997) extend the analysis in Zivot and Andrews (1992) to account for two structural breaks in trending variables using three deterministic specifications known as Models AA, AC-CA and CC, *i.e.* a model that accounts for two structural breaks that affect the level (Model AA), a model that accounts for two level shifts but just one slope shift (Model AC-CA), and a model that permits the two break dates to shift both the level and the slope (Model CC). In addition, Carrion-i-Silvestre, Sansó and Artís (2004) follow Lumsdaine and Papell (1997) and define a test statistic with two breaks for non-trending variables –that we call Model AAn.

Alternative proposals, that we are also considering here, focus on the statistical significance of the dummy variables when estimating the date of the breaks. We denote $t^*_{\alpha,|\theta|}$ the ADF test where the break point is chosen so that it minimizes the absolute value of the statistical significance test for the level shift, *i.e.* $T_b^* = \arg\max_{T_b \in (k+1,T)} \left|t_{\hat{\theta}}\left(T_b, k\right)\right|$, being $t_{\hat{\theta}}$ the pseudo t-ratio for θ in ((5)). Note that the maximization of the absolute value avoids imposing the sign of the break effect. If the analyst is willing to assume some break effect's direction, then the break point can be estimated through either the minimization –if the break effect

[7] Note that this test selects the break points that most favor the alternative hypothesis of stationarity, which implies that if the null hypothesis cannot be rejected at T_b^* it would not be rejected at whatever else date.

[8] The three specifications that defined are known as Models A, B and C, depending on if the break affects the level, the slope or both, respectively.

Table 2: Unit root tests with one level shift

	t^*_α	T^*_b	$t^*_{\alpha,\lvert\theta\rvert}$	T^*_b	$F^*_{\alpha,\theta}$	T^*_b
Czech Rep.	-3.845	1997:02	-1.921	1992:04	17.343^a	1992:04
Estonia	-4.042	1998:10	-4.042	1998:10	8.198	1998:10
Hungary	-4.466^c	1999:08	-2.191	1992:11	10.765^c	1992:11
Latvia	-2.647	1995:08	-1.819	2002:07	3.830	2002:07
Lithuania	-5.106^b	1998:12	-5.106^a	1998:12	13.105^b	1998:12
Malta	-2.916	2001:06	-2.916	2001:06	4.269	2001:06
Poland	-3.823	1998:07	-3.823	1998:07	9.716^c	1998:07
Slovakia	-4.621^b	1998:07	-4.621^b	1998:07	10.731^c	1998:07
Slovenia	-5.746^a	2000:02	-5.746^a	2000:02	16.508^a	2000:02

The critical values for the t^*_α test for $T = 100$ are -5.33, -4.58 and -4.27 for the 1, 5 and 10% level of significance, respectively, and are obtained from Perron and Vogelsang (1992). The critical values for the $t^*_{\alpha,|\theta|}$ test are -4.92, -4.38 and -4.09, respectively, and are obtained by direct simulation. Finally, the ones for the $F^*_{\alpha,\theta}$ test are 9.47, 10.61 and 12.99 for the 10, 5 and 1% level of significance, respectively, for a sample size $T < 75$, and 9.58, 10.84 and 13.48 respectively, for a sample size of $75 < T < 125$, and 9.72, 11.05 and 13.90, respectively, for $T > 125$.

is negative– or maximization –if it is positive– of the sequence of $t_{\hat{\theta}}$ tests. This increases the power of the testing procedure, although it introduces a prior belief.

Finally, Montañés and Reyes (1997) propose using the information linked to the statistical non-significance of the deterministic terms under the null hypothesis of unit root, as a mechanism to increase the power of the unit root testing. Specifically, it can be shown that the step dummy in ((5)) disappears under the null hypothesis but it is statistically significant under the alternative. Therefore, it is possible to design a test statistic for the joint null hypothesis of $\alpha = 1$ and $\theta = 0$ in ((5)). We denote this test as $F^*_{\alpha,\theta}$, where $T^*_b = \arg\max_{T_b \in (k+1,T)} F_{\hat{\alpha},\hat{\theta}}\left(T_b, k\right)$.

The results obtained from the application of the unit root tests allowing for one level shift are presented in Table 2, where the order of the autoregressive correction has been chosen with the *t*-sig criterion in Ng and Perron (1995) with $k_{\max} = 12$. It should be noted, first, that all these tests provide consistent estimates of the break dates. In addition, the unit root hypothesis is rejected for six out of the nine countries that we have analyzed when we allow for one level shift.

The same qualitative conclusion is reached when two breaks are considered – see Table 3. The unit root tests presented in Tables 1 and 2 have been specified with a deterministic component given by a constant term, the most adequate model according to the natural rate hypothesis. However, in practice, most time series

Table 3: Unit root tests with two structural breaks

	Model AAn			Model AA		
	Test	$T^*_{b,1}$	$T^*_{b,2}$	Test	$T^*_{b,1}$	$T^*_{b,2}$
Czech Rep.	-6.230^b	1997:03	1998:05	-6.668^b	1998:05	2001:08
Estonia	-5.631^b	1998:11	2001:07	-6.020^c	1998:11	2001:07
Hungary	-5.368^c	1997:11	2000:05	-5.797	1992:08	2000:05
Latvia	-4.562	1998:05	2002:07	-4.331	1996:01	1998:08
Lithuania	-5.135	1999:01	2003:04	-6.042^c	1998:01	2002:12
Malta	-3.220	1999:05	2001:10	-7.269^b	1999:12	2000:06
Poland	-5.909^b	1996:04	1998:08	-6.136^c	1992:12	1996:04
Slovakia	-6.226^b	1992:12	1998:10	-5.822	1992:12	1998:10
Slovenia	-6.283^b	1993:03	2000:03	-6.311^b	1993:03	2000:03
	Model AC-CA			**Model CC**		
	Test	$T^*_{b,1}$	$T^*_{b,2}$	Test	$T^*_{b,1}$	$T^*_{b,2}$
Czech Rep.	-6.789^b	1998:05	1994:09	-8.011^b	1997:03	1999:02
Estonia	-6.115	1998:11	2001:09	-6.131	1998:03	2000:09
Hungary	-6.050	2000:09	1992:10	-6.277	1992:10	2000:10
Latvia	-7.022^b	1998:09	2002:10	-6.789^c	1998:09	2002:08
Lithuania	-7.391^b	2002:02	1998:03	-7.450^b	1998:01	2001:08
Malta	-6.443^c	2002:06	2000:08	-6.472	1999:12	2001:07
Poland	-6.681^b	1994:01	1996:06	-6.368	1994:02	1996:04
Slovakia	-5.762	1992:12	1998:12	-4.980	1992:12	1998:10
Slovenia	-6.394^c	1993:03	2000:05	-6.175	1993:03	2000:03

The critical values for Model AAn are taken from Carrion-i-Silvestre et al. (2004), while for Models AA, AC-CA and CC they have been obtained from Lumsdaine and Papell (1997). b, c, stand for rejection of the null hypothesis of non-stationarity at 5 and 10% level of significance, respectively.

exhibit a trending behavior. Although this contradicts the natural rate paradigm in the limit, we should not discard the possibility of a time trend in the model for a short period of time. This fits those cases of a slowly increasing natural rate –see Papell et al. (2000).

Next we estimate the models allowing for a linear trend and more than one break (model AA without a time trend in Carrion-i-Silvestre et al. (2004), that we call model AAn, and models AA, AC and CC as in Lumsdaine and Papell (1997)). Compared to the results in Table 2, we find in this case stronger evidence against the unit root hypothesis, since it is possible to reject the null hypothesis of hysteresis for all the countries considered in, at least, one of the specifications. Therefore, we have found evidence of stationarity using test statistics that consider the presence of structural breaks, a result that supports the shifting natural rate hypothesis on the unemployment rate.

3.2 Shifting NAIRU in the Accession countries. Political reforms and structural breaks

In this section we give one step further and estimate the value of the NAIRU using a univariate approach. This approach, in contrast to other alternatives, based on the Phillips curve or on theories of the labor market, assumes that, over medium to long horizons, the unemployment rate reverts to its natural rate. Thus, univariate data on unemployment can be used to extract an estimate of the NAIRU as a local mean of the series. We have chosen this option in order to avoid an additional source of uncertainty related to the variety of feasible models explaining the NAIRU and the fact that many of the explanatory variables involved in the NAIRU calculations (such as inflation expectations or the degree of excess demand) are non-observable.

Here we follow the method in Staiger et al. (1997) who derive the NAIRU estimates from the specification given by

$$U_t - \bar{U}_t = \beta\left(L\right)\left(U_{t-1} - \bar{U}_{t-1}\right) + \varepsilon_t, \tag{6}$$

where $\bar{U}_t = \bar{U}_i$ for $T_{b,i-1} < t \leq T_{b,i}$, and $\beta\left(L\right)$ denotes the dynamics. This approach assumes a shifting natural rate of unemployment, which means that, in the long-run, the unemployment rate reverts to the path that describes the breaking-mean model. Therefore, an estimate of the NAIRU can be obtained as a local mean of the time series. Following the developments in Staiger et al. (1997), ((6)) can be expressed as

$$U_t = \mu + \theta_i DU_{i,t-1} + \beta\left(L\right)U_{t-1} + \varepsilon_t, \tag{7}$$

which implies a NAIRU, *i.e.* local mean, of

$$\bar{U}_i = \frac{\mu + \theta_i}{1 - \beta\left(1\right)}$$

for $T_{b,i-1} \leq t < T_{b,i}$, $i = 1, \ldots, n$. In the previous Section we have concluded that the unemployment rates can be characterized as stationary processes once the analysis has accounted for up to two structural breaks. However, we are aware of the limitations associated to the deterministic specifications considered above. First of all, the fact of just allowing for up to two structural breaks restricts the kind of models to face. Second, some of the trending patterns that have been highlighted above can be masking the presence of more than two structural breaks. In order to overcome these drawbacks, we proceed to the estimation of both the number of breaks and break dates as in Bai and Perron (1998), that is a suitable method to detect discontinuities in stationary processes but that has proven to provide consistent estimates of the break dates even for non-stationary variables.

3.2.1 Estimation of the number and position of the structural breaks

Briefly speaking, Bai and Perron (1998) suggest applying the following strategy. First, once a maximum number of breaks $(n^{\max})$ has been defined, we should obtain the global Sum of Squared Residuals (SSR) using a dynamic optimization algorithm. As a result, we have $n^{\max}$ sets of time periods which are taken as the estimates of the break dates. At this point, the goal is the estimation of the number of breaks $(\hat{n})$. For this purpose, Bai and Perron (1998) design a sequential testing procedure for non-trending variables that relies on testing for the presence of structural breaks, first, globally and, second, following a specific-to-general principle. For the case of trending variables, they suggest using the LWZ information criterion (Liu, Wu and Zidek (1997)).

This procedure allows not only to obtain an estimate of the number and position of the structural breaks, but also to calculate confidence intervals for the estimated break dates. This is especially relevant in our particular application, as we are interested in assessing the precision of the break point estimates due to the fact that it might affect, in turn, the precision of the NAIRU estimates. Table 4 reports the date and number of the estimated break points for the unemployment rates when a maximum of $n^{\max} = 5$ structural breaks have been allowed. The 90 and

Table 4: Point and confidence interval estimates for the date of the breaks

	$\hat{n}$	CI	$T^*_{b,1}$	$T^*_{b,2}$	$T^*_{b,3}$	$T^*_{b,4}$
Czech Rep.	5%		1996:07	1998:07		
		95%	(95:10, 96:09)	(98:1, 98:07)		
		90%	(95:12 ,96:09)	(98:2, 98:07)		
Estonia	5%		1997:05	1999:03	2001:10	
		95%	(97:03, 97:07)	(99:01, 99:02)	(01:08, 02:02)	
		90%	(97:04, 97:07)	(99:01, 99:02)	(01:08, 02:01)	
Hungary	5%		2000:3			
		95%	(00:02, 00:04)			
		90%	(00:02, 00:04)			
Latvia	5%		1996:01	1998:9	2000:08	2002:06
		95%	(95:10, 96:02)	(98:06, 98:09)	(00:06, 01:01)	(02:05, 02:06)
		90%	(95:10, 96:02)	(98:06, 98:09)	(00:06, 00:12)	(02:05, 02:06)
Lithuania	5%		1995:05	1999:11	2002:01	
		95%	(95:02, 95:06)	(99:5, 99:10)	(01:11, 02:09)	
		90%	(95:03, 95:06)	(99:5, 99:10)	(01:11, 02:09)	
Malta	5%		1998:02	1999:7	2001:11	
		95%	(98:05, 98:09)	(98:1, 99:8)	(01:09, 02:05)	
		90%	(98:06, 98:09)	(98:5, 99:8)	(01:09, 02:04)	
Poland	5%		1992:11	1996:08	2000:01	
		95%	(92:09, 93:08)	(96:05, 96:11)	(00:02, 00:07)	
		90%	(92:09, 93:05)	(96:05, 96:10)	(00:03, 00:07)	
Slovakia	5%		1998:12			
		95%	(98:10, 00:06)			
		90%	(98:10, 00:01)			
	10%		1993:01	1999:02		
		95%	(92:10, 95:07)	(98:08, 99:02)		
		90%	(92:11, 94:10)	(98:09, 99:01)		
Slovenia	5%		1993:09	2000:4		
		95%	(93:03, 93:12)	(00:01, 00:08)		
		90%	(93:04, 93:11)	(00:02, 00:07)		

The second column reports the significance level that has been used for the sequential estimation of the break points. It refers to the 5% level of significance, but the same results are achieved when working at the 10% level (Slovakia being the only exception, as indicated in the table). The third column offers the confidence intervals at the 95% and 90% of probability. The maximum number of structural breaks is set to $n^{max} = 5$.

95% confidence intervals for the break dates are also presented, and are computed using a heteroscedasticity and autocorrelation robust estimate of the variance of the disturbance term. These results show that Hungary and Slovakia have suffered one level shift, the Czech Republic and Slovenia two structural breaks, Estonia, Lithuania, Malta and Poland three breaks and, finally, four structural breaks are detected in the case of Latvia. Note that for Malta we have estimated three level shifts, which can explain why we had to include a time trend in the specification of the unit root hypothesis testing conducted in the previous Section. Looking at the narrow confidence intervals obtained, we can conclude that the break points have been estimated quite precisely.

As the rest of the analysis is based on the structural changes estimated using the Bai and Perron (1998) method, we present in Table 5 (columns 1 to 3) a comparison of the estimated break dates obtained using the different unit root tests[9] in the previous section, with the breaks obtained with the Bai and Perron (1998) procedure. The main conclusion that can be drawn from the table is that they are broadly compatible. As Bai and Perron (1998) argue, their procedure provides consistent estimates of the break-dates independently of the stationarity nature of the variables. In the majority of the cases, when there are multiple changes, the 1-break and 2-break tests are capturing part of them or, at least, are detecting an instability period in the variable. This improvement in the deterministic specification explains the higher power of these tests if compared with those that do not account for instabilities, as they are able to reject the null of non-stationarity (either with one or two breaks) for all the countries involved.

As an example of this compatibility, in the case of Slovenia the unit root tests with one break estimate it in 2000:02. When we allow for two breaks, they appear

[9] The first column includes the resulting break-dates obtained using tests allowing for one break. We present two dates when the tests gave different results. The second column corresponds to the test with two breaks in the levels in the models with non-trending regressors (AAn) and with trending regressors (AA). The third column reports the results of the Bai and Perron methodology to consistently estimate structural breaks.

Table 5: Structural breaks. Comparison of the different methods

	1 break tests	2 break tests	n breaks	León-Led/McAdam 1 break (trend model)	León-Led/McAdam Markow Switching
Czech Rep.	1992:04 1997:02	AAn: 1997:03 AAn: 1998:05 AA: 1998:05 AA: 2001:08	1996:07 1998:07	1992:07 1998:04	1997-98
Estonia	1998:10	1998:11 2001:07	1997:05 1999:03 2001:10	1998:10 2000:05	Multiple changes (1995, 1996, 1999...)
Hungary	1992:11 1999:08	AAn: 1997:11 AAn: 2000:05 AA: 1992:08 AA:2000:05	2000:03	1992:11	Multiple changes
Latvia	1995:08 2002:07	AAn: 1998:05 AAn: 2002:07 AA: 1996:01 AA: 1998:08	1996:01 1998:09 2000:08 2002:06	1998:04	1998 2000
Lithuania	1998:12	AAn: 1999:01 AAn: 2003:04 AA: 1998:01 AA: 2002:12	1995:05 1999:11 2002:01	1997:01	1998-1999
Malta	2001:06	AAn: 1999:05 AAn: 2001:10 AA: 1999:12 AA: 2000:06	1992:11 1999:07 2001:11		
Poland	1998:07	AAn: 1996:04 AAn: 1998:08 AA: 1992:12 AA: 1996:04	1992:11 1996:08 2000:01	1996:03	
Slovakia	1998:07	1992:12 1998:10	n=1: 1998:12 n=2: 1993:01 n=2: 1999:02	1992:11	1998
Slovenia	2000:02	1993:03 2000:03	1993:09 2000:04	1999:06	1994 1996 2000

in 1993:03 and 2000:03. Finally, the Bai and Perron (1998) method places them (also 2) in 1993:09 and 2000:04. If we allow for the 5% confidence intervals (already reported in table 4) all the results are highly in line.

The last two columns of table 5 are devoted to summarizing the main results obtained by León-Ledesma and McAdam (2004) 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 León-Ledesma and McAdam (2004) 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.

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

Before going any further in our estimation of the NAIRU, it is worth to shed some light on the general or systemic characteristics of the transition process that are on the ground of some of the breakpoints found in our analysis as well as other specific shocks or circumstances that affected the different economies in a heterogenous way. Eight of the new member states have a particularly interesting economic background. 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 (Roland (2000)). 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 (Boeri and Terrell (2002)).

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.

Among the EU Accession countries, the three Baltic countries differ in important ways from the remaining five central European members of this group. In addition, transition economies display an enormous diversity in terms of physical and population size, level of development (as measured by GDP per capita), natural resource endowment, and cultural and historical background, greatly complicating intercountry comparison. Countries' actual transition experience has differed enormously, with respect to both policies implemented and results achieved to date. The reason for the differences include the country's initial conditions, the external environment (notably external shocks), and the specific policies pursued during the transition (IMF (2001)).

From the point estimates of the breakpoints presented in Table 4, a first broad clustering of the countries in two groups can be made in terms of the number of breaks. The first one consists of the small countries (Baltic countries and Malta) plus Poland, that display from 3 to 4 structural changes. The second group includes medium-size economies (Czech Republic, Hungary, Slovakia and Slovenia) that only experience one or two breaks. Thus, before proceeding to estimate the short-term NAIRU, in the next subsection, we will also describe the countries' characteristics that may help to explain the timing of the breaks.

Firstly, the Baltic countries have had high growth but, at the same time, relatively high unemployment. This can be partially explained by the fact that they were all affected by the Russian economy crisis in 1998 (see Table 6). Estonia's economy is small, just 0.4% the size of France's economy and 2.3% of France population[10]. Because of the country's small size and the open nature of its economy, it is extremely vulnerable to external shocks. After its independence

[10] The details about political events and the size and population of the transition countries has been mainly obtained from Network (2003).

in 1991 and an adjustment process, unemployment was decreasing (see the 1997 break). However, the Russian crisis pushed it upwards over 12% in 1999. This economy can be characterized by its flexibility, so that the recovery was rapid and the new structural change in 2001 led the unemployment back to previous levels.

Latvia's economy is tiny compared to most EU-15 and EU Accession countries. Its GDP is just a 0.6% of France's and 3.9% of its population. Latvia's GDP was the lowest of the 10 countries accepted for membership in 2004. As in the previous case, its small size and openness leaves this economy at the mercy of external shocks. In fact, Latvian economy was the most tightly linked to Russia, because of the structure of Latvian exports. Of the four structural changes found in the case of Latvia, the first two are associated with increases in the short-term NAIRU, whereas the two last show an improvement in the labor market. As in Estonia, the 1998 structural break can be linked to the Russian crisis. In such an open economy, persistent current account deficits have been a bottleneck to Latvia and have caused higher volatility in this country's growth and employment rates.

Lithuania's economy is the biggest among the Baltic states, but still is a small open economy and shares with the others the same external constraint. After experiencing one of the largest output declines of the former soviet republics at the beginning of the nineties, the economy started growing in 1995, after the currency's peg to the dollar in 1994. However, in 1999-2000, Lithuania suffered significantly more than Estonia's or Latvia's in the aftermath of Russia's financial crisis. The rapid unemployment increase is captured in the 1999:11 structural break and, although the third break in 2002 is reflecting a recent recovery, the unemployment rate remains comparatively high. Lithuania has some structural characteristics, such as a relatively high minimum wage and restrictions on hiring part-time employees, which will continue to make unemployment rates difficult to reduce. The above mentioned recovery in 2002 can be possibly due to the measures undertaken in 2001 aiming at liberalizing the labor market as well as a faster privatization process and a sound fiscal discipline before the new peg to the euro in February 2002.

The last small country, Malta, is 0.3% of France's GDP and 0.7% of population. However, on a per capita basis, Malta ranks among the richest new EU members. Due to the country's small size, its large dependence on tourism revenues and the increasing open nature of its economy, makes it very susceptible to external shocks. The three structural breaks found can be associated to global shocks, such as the Russian and Latin American crises at the end of the nineties, but specially the geopolitical tensions in the Middle-East and the downturn in world tourism in the aftermath of September 11th. Therefore, Malta's greatest risks lie in the country's overdependence on tourism revenues and the lack of diversity of its manufacturing sector.

Finally, the case of Poland deserves special attention, as it is the only big country that has experienced three structural changes, according to our results. Although Poland is by far the largest of the EU accession countries (more than three times the size of the Czech Republic, the second largest economy among them) and, consequently, less exposed to external shocks that the small transition countries, the strategy followed during the transition period explains a great deal of its unemployment high levels and changes. Roughly, the size of the Polish economy in terms of GDP is 14% that of France. However, in terms of GDP per capita it ranks the sixth among the CEECs. Partially because of its size Poland features the least open economy among the EU accession countries. During most of the 1990s Poland was considered to be the undisputed leader among the European transition economies implementing a radical "shock therapy" to its economy in 1990. This early increase in unemployment is captured in the first structural change found in our analysis (1992:11) and it is also evident in Figure 1. Therefore, Poland was the first country in the region to come out of the transition recession, reporting positive growth rates in GDP already in 1992. At the same time, it was the first country to regain the pre-transition level of GDP in 1997. A second structural change is found in 1996:08, that can be associated to the effects of the output growth in employment. However, this expansion in 1994-97 led to substantial external imbalances and relatively high inflation. Growth in domestic demand was also supported by rather lax fiscal policy on the part of the Social Democratic government. Consequently, the National Bank of Poland tightened monetary conditions in an attempt to adjust the economy. Unfortunately for Poland, this policy adjustment coincided with the Ruble crisis in Russia in 1998 and the slowdown in growth in Poland's main export markets in the EU, such as Germany. As a result of that, Poland's economy growth also diminished to 1.0-1.3% rates. The third structural change found in our analysis (2000:01) is consistent with this recession.

While inflation was clearly the weakest point in Poland's economy in the early stages of transition, unemployment is now, by far, the most important medium term problem facing the Polish economy. Much more substantial changes to the labor market structure and the employment taxation and regulations will be necessary to reduce the unemployment problem more decisively[11].

The second group of countries exhibiting one or two breaks are all of them medium size economies.

First, Slovakia ranks fourth in terms of population and GDP among the 10 Accession countries, but behind the other Visegrad countries. Slovakia has the same number of inhabitants than Denmark, but only about one eighth of its size in terms of GDP. Compared with France, Slovakia has only 9% of its population and 2% of its GDP. Historically, Slovakia was much more rural than the neighboring Czech Republic. Many of the firms built in Slovakia during the socialist period were dedicated to the production of heavy industry and weapons to export to the

[11] This point will be more extensively documented in section 3.3 where the persistence results and the institutional discussion are presented.

Soviet Union. Once trade with the USSR collapsed after 1989, Slovakia was in a difficult position. Unemployment quickly surged, moreover, observers were skeptical about the future of the Slovakian economy after their split from the Czechs in 1993. Slovaks launched coupon privatization in the early nineties, but the program stopped after the split, being replaced by the so-called "crony capitalism", where firms were sold to domestic allies of the ruling parties at rock-bottom prices. That approach ended with 1998 Parliamentary elections, when a pro-western government consisting of a broad range of parties took control of the country, saving it from imminent collapse. The two structural changes detected using Bai and Perron method are consistent with these facts: the first one is found in 1993:01, whereas the second appears at the end of 1998-beginning of 1999.

Second, Slovenia's economy is 1.5% of the size of the France's and its population 3.3% of the number of French inhabitants. However, Slovenia's per capita GDP ranked as the highest among the EU accession countries: only the Czech Republic and the Hungary's GDP equal half of Slovenia's level. Slovenia's small size and relative ethnic homogeneity have minimized political problems, allowing the government to concentrate on supporting economic growth and reform. Emerging from the former Yugoslavia in 1991, the country's ties to Western Europe (and specially Italy and Germany) laid the ground for rapid economic development. Services, and more specifically tourism, is a key industry in the country. However, the largest single sector remains manufacturing. Although it suffered from low productivity and relatively poor international competitiveness, the government sponsored a wide program of employment retraining and administrative limits on wages that boosted employment in the late nineties. As evident from the corresponding graph in Figure 1, the two structural breaks found in 1993:09 and 2000:04 are associated with reductions in unemployment. In fact, the present rate is currently near its lowest historical level, around 6%, with most industrial restructuring nearly finished and production projected to expand.

Third, the Czech Republic is among the largest and richest of the EU accession countries. The country ranks second after Poland and closely followed by Hungary. The Czech Republic comprises some of the historically wealthiest and most industrialized territories in Europe. Although many local firms lost their competitiveness during 40 years of communism, the country has retained certain advantages in terms of brand-names and manufacturing tradition. The country has been revitalized since 1989, attracting foreign tourists and investment. In the early 1990s, the country was leader in economic reforms promoting liberalization. Nonetheless, by 1996 many flaws had been revealed, as the pace and the strategy of the reform (coupon privatization and lack of regulation) were inadequate. This lack of effective reforms kept unemployment unnaturally low. It was not until former prime minister Vaclac Klaus and his allies lost power in late 1997 that major restructuring took place. The two structural changes found in the empirical analysis (1996:07 and 1998:07) show the rapid increase in unemployment in the second half of the nineties.

Finally, Hungary's GDP is about 4.5% of France's and, in terms of population, it is similar to the Czech Republic. This country started the transition process in a strong position, as the communist government in the 1980s had already installed some basic features of a market oriented economy, such as tax incentives in special economic zones, that attracted highly qualified work force and a rapid privatization process. After the 1995 stabilization plan, "Bokros program", the last years of the 1990s, under the coalition of Socialists and Free Democrats, have achieved an improvement of the external position of the economy, accelerating liberalization. The central bank, in cooperation with the government, employed a crawling-peg exchange rate regime, that was abandoned in May 2001 in favor of entering the ERM-2. The policy change and the advance of reforms is reflected in a slow recovery of employment, as depicted in the corresponding graph and captured in the structural change found in 2000:03.

3.2.3 OLS and Median Unbiased NAIRU estimates

Once the break-points have been detected, we can calculate the NAIRU both for the point and confidence interval estimates. The order of the autoregressive model in ((7)) has been set using the LWZ information criterion allowing up to $k_{\max} = 12$ lags. In order to save space and to avoid repetitions, we report in Table 7 the NAIRU values using the point estimates of the break points, along with the estimations that are obtained using the lower and upper bound of the 95% confidence interval for the break dates –the results for the 90% confidence interval are also available upon request. This information provides us a first insight on how precisely the NAIRU is estimated, something that it has not been addressed in previous empirical literature. Looking at these estimates we have to conclude that, in general, the NAIRU is estimated with high precision given the small discrepancies among the estimated NAIRUs. Less precise estimates are obtained for Slovakia when only one level shift is allowed. However, the precision of the NAIRU estimates for this country increases when two level shifts are included, which means that the latter is the best specification for Slovakia. The NAIRU estimates are depicted in Figure 1 and compared to the original variables.

Despite the good results in terms of precision of the NAIRU estimates discussed above, these NAIRU univariate-based estimates are a biased approximation due the procedure used in the computations[12]. In order to correct for this distortion, Andrews (1993) suggests computing the exact median-unbiased (MU)

[12] It is well known that the estimation of autoregressive specifications like ((6)) or ((7)) by OLS produces biased estimates of the parameters and, consequently, a biased estimated NAIRU. Furthermore, the closest a stochastic process is to non-stationarity, the higher is the bias distortion. In this case, the process is close to being non-stationary and, as the least squares estimator minimizes the regression residual variance, it will tend to make the data-generating process appear to be more stationary than it actually is by forcing the autoregressive parameter away from unity.

Table 6: 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**	Havel lost power Post-transition recession	**1998:07** Restructuring				
Estonia						**1997:05**	Russian crisis→	**1999:03**		**2001:10** Current account deficit	
Hungary				Early reform Stabilization Plan (1995)					**2000:03** Employment recovery	ERM-2	
Latvia	Tiny open economy Exposed to external shocks				**1996:01**		**1998:09** Russian crisis→		**2000:08**		**2002:06**
Lithuania				**1995:05** Exchange rate peg			Russian crisis→	**1999:11**		Faster privatization Fiscal discipline	**2002:01** New peg
Malta	Tiny open economy Exposed to external shocks						**1998:02**	**1999:07**		**2001:11** September 11^{th} Tourism	
Poland	**1992:11** Shock therapy for transition				**1996:08** Monetary tightening		Russian recession (↓GDP)		**2000:01**		
Slovakia		**1993:01** Splits from Czech Rep.					1998 elections	**1999:02**			
Slovenia	1991: splits from Yugoslavia	**1993:09**							**2000:04**		

estimates for AR(1) processes. The goal is to substitute the OLS estimate for its median unbiased counterpart, as the latter does not favor any particular outcome. Thus, this bias correction delivers an impartiality property to the decision-making process because there is an equal chance of under or overestimating the autoregressive parameter. This issue is of special interest when measuring persistence in time series, given that for positive AR(1) processes the estimate of the autoregressive parameter is downward biased. Moreover, an unbiased estimate of the autoregressive parameter will allow us to calculate an unbiased scalar estimate of persistence in the next sub-section.

The direction of the bias is not clearly defined for models of higher order than AR(1), although the estimate of the sum of the autoregressive coefficients tends to exhibit a downward and large bias –see Andrews and Chen (1994). Once again, this is extremely important for our purpose since the sum of these AR coefficients is involved in the computation of the NAIRU. Thus, if precise NAIRU estimates have to be obtained, the estimation bias should be accounted for. For this purpose, we can apply the procedure proposed by Andrews and Chen (1994), who extend the approach in Andrews (1993) to $AR(p)$ processes, and design an iterative process to *approximate* median-unbiased estimators for models such as the ones in ((6)) or ((7)). Their procedure only provides approximate MU estimates due to the use of estimators rather than true parameters in one stage, and due to the use of pseudorandom numbers –see Andrews and Chen (1994) for further details[13].

In addition and more importantly, the application of these techniques produces confidence intervals for the parameters of interest, *i.e.* the sum of the AR coefficients, which provides further information to the analyst on the accuracy of the estimates. Thus, interval estimation addresses the low-power problem associated with unit roots, by informing us of whether we are failing to reject

[13] In brief, the iterative procedure starts with the OLS estimation of

$$u_t = \mu + \theta_i DU_{i,t-1} + \alpha u_{t-1} + \gamma\left(L\right)\Delta u_{t-1} + \varepsilon_t,$$

Table 7: NAIRU. Point and 95% confidence interval estimates

NAIRU. Point estimates for $T^*_{b,i}$

	Regime				
	1st	2nd	3rd	4th	5th
Czech Rep.	3.25	5.61	8.08		
Estonia	9.96	9.03	12.13	9.75	
Hungary	7.63	5.63			
Latvia	10.07	11.50	14.02	12.90	10.91
Lithuania	6.72	10.23	15.56	13.14	
Malta	5.87	6.41	6.82	7.70	
Poland	12.14	15.42	13.58	18.50	
Slovakia	12.50	18.35			
	11.05	12.65	18.40		
Slovenia	7.57	7.13	6.12		

NAIRU. Lower bound for the 95% CI of $T^*_{b,i}$

	Regime				
	1st	2nd	3rd	4th	5th
Czech Rep.	3.21	4.55	8.03		
Estonia	9.97	8.96	12.15	9.77	
Hungary	7.63	5.63			
Latvia	10.07	11.39	14.02	12.87	10.76
Lithuania	7.11	9.92	15.27	13.07	
Malta	5.99	6.72	6.71	7.74	
Poland	12.03	15.75	13.35	18.59	
Slovakia	12.30	18.41			
	10.80	12.44	18.26		
Slovenia	7.17	7.21	6.15		

NAIRU. Upper bound for the 95% CI of $T^*_{b,i}$

	Regime				
	1st	2nd	3rd	4th	5th
Czech Rep.	3.18	5.71	8.08		
Estonia	9.83	9.40	11.57	9.96	
Hungary	7.60	5.60			
Latvia	10.26	11.73	13.65	12.85	10.64
Lithuania	7.11	10.30	15.06	12.74	
Malta	6.15	6.68	7.02	7.79	
Poland	13.59	15.05	14.02	18.88	
Slovakia	16.01	16.50			
	12.42	12.93	18.39		
Slovenia	7.77	6.97	6.21		

For Slovakia we present the NAIRU that is obtained with one and two level shifts in the first and second entry for this country, respectively.

the null because it is true or because there is too much uncertainty about the true value of the autoregressive parameter. All these elements allow us to assess the precision with which the NAIRU values have been approximated. Table 8 presents these OLS and MU estimates of α. Note that the $\hat{\alpha}_{MU}$ estimates are always larger than the $\hat{\alpha}_{OLS}$ ones, which reflects the downward bias associated to the OLS estimation method. The consequence of working with a higher α is the increase in the NAIRU estimates –this becomes evident from the comparison of Tables 7 and 8. Moreover, notice that the confidence intervals that have been computed for the $\hat{\alpha}_{MU}$ reveal that, in general, non-stationarity is found at the upper bound. The case of Slovakia is remarkable, since we have obtained that $\hat{\alpha}_{MU} = 1$ when we only allow for one structural break. However, when taking into account two level shifts the $\hat{\alpha}_{MU}$ drops below one. This suggests that it is better to consider two level shifts for this country, as commented above.

3.2.4 Modified Phillips Curves

According to Richardson et al. (2000) the short-run NAIRU indicator is probably of the greatest relevance to the operation of monetary policy. By construction the short-run NAIRU gap will be closely correlated with contemporaneous or predicted changes in inflation. The short-run NAIRU can be seen as a useful synthesis of information concerning current inflationary pressures. However, the usefulness as a forecast for future inflation is limited for several reasons. First, differences between NAIRU and short-run NAIRU are likely to

which is a reparametrization of ((7)) with $\alpha = \beta(1)$. Treating $\hat{\gamma}^1(L)$ as if they were the true values of $\gamma(L)$, the procedure computes the bias-corrected estimator of α, denoted as $\hat{\alpha}^1_{MU}$, as the value that has as a median $\hat{\alpha}^1_{OLS}$. Then, conditional on $\hat{\alpha}^1_{MU}$ we obtain the estimates of $\hat{\gamma}^2(L)$. After that we conduct another iteration, assuming that $\hat{\gamma}^2(L)$ are the true values of $\gamma(L)$ and computing a second bias-corrected estimator of α, $\hat{\alpha}^2_{MU}$, as the value that has as a median $\hat{\alpha}^2_{OLS}$. The iterations for $j = 1, 2, \ldots$ continue after some convergence criterion is achieved– due to the computational cost we have specified as the convergence criterion that the difference between two consecutive iterations should be $\left|\hat{\alpha}^{j+1}_{MU} - \hat{\alpha}^{j}_{MU}\right| < 0.01$. Andrews and Chen (1994) report simulation results concluding that this approximation provides accurate median-unbiased estimates.

Table 8: MU-based NAIRU estimates

NAIRU. Point estimates for $T^*_{b,i}$

			95% CI for $\hat{\alpha}_{MU}$		NAIRU estimates using $\hat{\alpha}_{MU}$. Regimes				
	$\hat{\alpha}_{OLS}$	$\hat{\alpha}_{MU}$	Lower	Upper	1st	2nd	3rd	4th	5th
Czech Rep.	0.950	0.976	0.932	1	3.42	7.21	8.33		
Estonia	0.797	0.836	0.748	0.934	9.91	9.28	12.04	9.71	
Hungary	0.920	0.945	0.887	1	7.65	5.56			
Latvia	0.810	0.846	0.768	0.931	10.19	11.68	13.87	12.92	10.77
Lithuania	0.904	0.926	0.880	0.974	7.28	10.38	15.53	13.20	
Malta	0.841	0.967	0.807	1	9.94	7.39	7.24	8.64	
Poland	0.971	0.988	0.960	1	10.58	13.50	15.50	17.75	
Slovakia	0.939	0.961	0.911	1	12.79	18.28			
	0.923	0.944	0.898	1	11.59	12.96	18.46		
Slovenia	0.934	0.970	0.907	1	9.17	6.77	6.03		

NAIRU. Lower bound for the 95% CI of $T^*_{b,i}$

			95% CI for $\hat{\alpha}_{MU}$		NAIRU estimates using $\hat{\alpha}_{MU}$. Regimes				
	$\hat{\alpha}_{OLS}$	$\hat{\alpha}_{MU}$	Lower	Upper	1st	2nd	3rd	4th	5th
Czech Rep.	0.964	0.998	0.949	1	6.00	23.50	13.00		
Estonia	0.808	0.845	0.756	0.985	9.96	8.99	12.18	9.70	
Hungary	0.921	0.944	0.884	1	7.75	5.59			
Latvia	0.872	0.925	0.846	1	10.07	11.51	14.28	12.75	10.36
Lithuania	0.927	0.954	0.904	1	7.85	10.04	15.59	13.00	
Malta	0.873	0.977	0.833	1	13.30	15.30	7.35	9.65	
Poland	0.964	0.983	0.949	1	11.53	15.35	14.24	18.71	
Slovakia	0.927	0.946	0.897	1	12.43	18.41			
	0.925	0.956	0.896	1	11.75	12.61	18.43		
Slovenia	0.905	0.934	0.873	1	7.36	7.20	6.12		

NAIRU. Upper bound for the 95% CI of $T^*_{b,i}$

			95% CI for $\hat{\alpha}_{MU}$		NAIRU estimates using $\hat{\alpha}_{MU}$. Regimes				
	$\hat{\alpha}_{OLS}$	$\hat{\alpha}_{MU}$	Lower	Upper	1st	2nd	3rd	4th	5th
Czech Rep.	0.944	0.966	0.925	1	3.26	6.62	8.09		
Estonia	0.875	0.917	0.838	1	9.77	9.41	11.48	9.93	
Hungary	0.922	0.950	0.888	1	7.76	5.56			
Latvia	0.835	0.876	0.798	0.957	10.22	11.68	13.60	12.80	10.63
Lithuania	0.926	0.947	0.906	0.996	7.49	10.55	15.21	12.72	
Malta	0.870	0.989	0.841	1	14.91	10.18	8.73	10.73	
Poland	0.974	0.994	0.963	1	15.17	11.33	18.00	19.17	
Slovakia	0.983	1	0.973	1	∞	∞			
	0.938	0.969	0.917	1	13.23	13.58	18.52		
Slovenia	0.950	0.987	0.934	1	11.23	6.23	6.31		

be most marked for those economies characterized by strong persistence effects (such as some transition economies). Secondly, the short-run NAIRU gaps have restrictions to be used to forecast future inflation because they are based on the current short-term NAIRU estimated, observed unemployment rates and the unrealistic assumption that there are no future supply shocks and that medium term NAIRU remains constant.

Bearing the above mentioned limitations in mind, we have used the short-run NAIRU values resulting from the MU estimates obtained in the previous sections to draw a modified Phillips curve as in Hahn (1996). In Figure 2 we present, for all the countries considered in the analysis, these modified Phillips curves with the percentage gap in unemployment rates (the difference between the actual and the NAIRU rate of unemployment) in the horizontal axis and consumer prices inflation in the vertical axis –the data is drawn from the International Financial Statistics CD-ROM of the IMF. According to the theory, we expect a negative relationship between the two variables provided that the expectations correspond to a short-run Phillips curve. Changes in the stance of monetary policy and, therefore, in the agents' expectations, may shift the curve upwards or downwards. Along a short-run Phillips curve, high (low) inflation rates are associated with current unemployment below (above) the short-term NAIRU.

This is the picture that results from the majority of the accession countries in Figure 2: the relation between the inflation rate and the unemployment gap is clearly negative. Moreover, with the exception of Malta, where the inflation rate is very low for the whole period and the sample shorter than in the other cases, the short-run Phillips curves have been moving downwards, as the monetary authorities gradually were able to reduce the high inflation rates experienced at the beginning of the transition period.

Although the process is more acute in the Baltic countries, all the countries analyzed exhibit the same pattern: they depart in the early 1990s from high levels of inflation that are progressively reduced. Their relation with the unemployment gap describes a zigzag behavior. However, once low levels of inflation are achieved, the unemployment rate gap does not seem to affect so strongly the inflation rates.

From the picture, the short-run Phillips curves cross several times, in all the countries in the sample, the vertical line drawn at zero. Along this line the current rate equals the short-run NAIRU, so that this means that this works as an attractor or equilibrium value. This would validate the natural rate hypothesis and, indirectly, the adequacy of the techniques we have applied to compute the short-term NAIRU.

3.3 Measuring persistence

Shocks persistence has been extensively analyzed in the macroeconomic literature primarily related to output, the labor market –unemployment rates– and prices –deviations from the purchasing power parity (PPP). There are several concepts of persistence associated with the use of different techniques applied to its measure, although the common practice consists of summarizing the persistence in a scalar –see Cochrane (1988), Andrews and Chen (1994) and Murray and Papell (2002).

The most extended approach is the one that relies on the computation of the half-life of a shock. This definition measures persistence as the number of time periods required for a unit impulse to dissipate by one half, *i.e.* denotes the median lag. The popularity of this approach is mainly due to the ease of computation. Thus, when the process is an AR(1) the half-life (HL) can be easily obtained as $HL = \ln(0.5) / \ln(\alpha)$, where α denotes the autoregressive parameter. However, this approximation to the HL does not work for AR processes of order higher than one, as it does not account for the dynamics of the stochastic process. Instead, the HL should be computed by the impulse response function (IRF).

In our case, we will concentrate on the computation of the half-life of a shock affecting unemployment based on the MU estimates. As Cashin, McDermott and Pattillo (2004) point out, the MU point and interval estimates of the HL can be interpreted as follows: first, when determining whether the variable suffers transitory (finitely persistent) or permanent (infinitely persistent) shocks, the selection rule chooses transitory if the bias-corrected half-life is finite and chooses permanent if the bias-corrected half-life is infinite; second, the width of the confidence interval allows assessing the level of uncertainty about the true persistence of the shocks.

In Table 9 we present the results of the computation of the measure of persistence based on the MU estimates obtained in the previous Section. In this case, instead of using the alphabetical order, as in previous tables, we have ordered the countries from those that exhibit more rapid adjustment, at the top, to these for which adjustment is sluggish. This ordering will later ease the interpretation of the results. As before, we have also focused on its sensitivity to the break points estimates. Thus, the first column of Table 9 reports the results obtained using the point estimates of $T^*_{b,i}$, while the second and third columns correspond to the lower and upper bounds that define the 95% confidence interval for $T^*_{b,i}$. The computations are also carried out using the lower and upper bounds of the 95% confidence interval for the α parameter, which in turn provide a confidence interval for these measures of persistence. The half-lives are, with the exception of Slovakia with one level shift, finite. According to the selection rule explained above, the unemployment rate would suffer just transitory shocks. In addition, when looking at the half-lives for the point estimates of $T^*_{b,i}$ we conclude that they are below two years. With the exceptions of the Czech Republic, Malta, Poland and Slovenia, this result is robust when we compute the half-life using the lower and upper bounds of the break point estimates. The medians of HL are 7.86, 7.54 and 9.06 for the point, lower and upper bounds of $T^*_{b,i}$ estimates,

Table 9: Half-life using MU estimates

	Point estimates for $T^*_{b,i}$	95% Confidence interval for $T^*_{b,i}$	
		Lower bound	Upper bound
Latvia	2.52	4.69	2.92
	(1.85, 4.76)	(2.61, ∞)	(2.03, 7.22)
Estonia	2.88	3.19	5.53
	(1.92,7.31)	(2.04, 34.86)	(2.82, ∞)
Lithuania	3.84	5.56	4.80
	(2.79, 8.23)	(3.35, ∞)	(3.26, 47.34)
Hungary	7.86	7.77	8.59
	(4.10,∞)	(4.00, ∞)	(4.12, ∞)
Slovakia	8.37	6.47	∞
	(3.99, ∞)	(3.74, ∞)	(11.45, ∞)
	5.76	7.54	10.05
	(3.43, ∞)	(3.59, ∞)	(4.17, ∞)
Czech Rep.	12.72	72.30	9.06
	(4.83, ∞)	(6.01, ∞)	(4.47, ∞)
Slovenia	14.43	6.10	32.29
	(4.68, ∞)	(3.43, ∞)	(6.19, ∞)
Poland	17.85	14.22	31.41
	(8.04, ∞)	(6.99, ∞)	(8.39, ∞)
Malta	17.88	27.49	52.87
	(2.50, ∞)	(2.93, ∞)	(3.01, ∞)
MEDIAN	7.86	7.54	9.06
	(3.43, ∞)	(3.43, ∞)	(4.12, ∞)

The second column provide the HL estimates using the break points estimates, while the third and fourth columns correspond to the HL estimates using the lower and upper bounds, respectively, of the 95% confidence interval for the break points estimates. The confidence interval for the HL appears between parentheses.

respectively, which indicate that deviations from the NAIRU correct quite rapidly. These results crucially depend on the differences in the labor market institutions that have been set up during the transition period.

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 Layard and Jackman (1991) and Scarpetta (1996)

Table 10: Main characteristics of the unemployment benefit system

	Employment conditions	Waiting period	Payment rate %	Duration (months)
Czech Republic	12 months in 3 years	7 days	50% 40% after 3 months	6
Estonia	12 in 24 months	7 days	50% for first 100 days 40% thereafter	180 days, longer in case of long insurance
Hungary	12 months in 4 years	—	65%	12
Latvia	9 months in 12 months	?	50% for 6 months, then depending on employment and duration	9 months
Lithuania	24 months in 3 years	7 days	19-34%	6 months in 12 months
Malta	?	?	Flat rate	150 days
Poland	365 days in 18 months	1 day	Flat rate	18 max (lower in areas with low unemployment)
Slovenia	12 in last 18 months	?	70% of average wage in last 12 months for 3 months, 60% afterwards	3-24 months depending on length of employment history
Slovakia	24 months in 3 years	—	50% 45% after 3 months	6 or 9 months depending on length of employment history

Extracted from Burger (2003).

Table 11: Minimum wages

	Minimum wage as % of average wage	Year	Monthly minimum wage, January 2003	
			in	in PPP
Czech Rep.	34.0%	2001	199	389
Estonia	28.2%	1999	138	264
Hungary	40.0%	2001	212	384
Latvia	40.0%	2000	116	239
Lithuania	40.0%	2001	125	252
Malta	74.0%*	2001	535	752
Poland	38.0%	2000	201	351
Slovakia	38.5%	2000	118	265
Slovenia	58%	target value	451	668
EU-15	45.0%-50%		Min: 416	Min: 543

Extracted from (Burger, 2003). * in percentage of average net rates.

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[14]. 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 Boeri et al. (1998) the transition process involves new job matches using workers with different skills that should generate an explosion of earning differentials at all levels, between 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

[14] See Elmeskov and Scarpetta (1998), Blanchard and Wolfers (2000) and Nickell and Quintini (2002).

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.

Going back to Table 9, we can see that there is a link between the speed of adjustment from a short-term NAIRU to the next one and the labor market institutions. In table 10 we summarize the main characteristics of the unemployment benefit systems of the accession countries, whereas in table 11 we present the minimum wages. From the comparison of the three tables it becomes evident that those countries that enjoy more social benefits and higher minimum wages are those which exhibit the largest half-lives. Examples of this behavior are Poland, Malta, Slovenia and the Czech Republic. Take, for example, the unemployment and social assistance paid when unemployment insurance is exhausted, that just amounts to 15% of the average wage in the Baltic countries, over 20% in Hungary, and about 33% in the Czech and Slovak republic. This matches exactly the ordering found when computing the half-lives in Table 9, up to the particular case of Poland (which is at the bottom of the table), that has the most generous benefit scheme no matter the indicator used.

4. Conclusions

In this paper we address two questions related to the behavior of the unemployment rate in the EU Acceding countries. First, we test for hysteresis versus natural rate hypothesis; second, once we assume that the time series properties of the variables are compatible with a changing natural rate, we use a univariate approach to measure the NAIRU as the local mean of unemployment in-between structural changes. In addition to these two policy-oriented goals, we devote some space to the discussion of the sources of uncertainty that produce imprecision in the estimates of the NAIRU. This is a key issue that should put a word of caution before the NAIRU estimates are used in policy-making. A conclusion of the analysis, also found in previous very comprehensive studies, such as Staiger et al. (1997), is the risk of obtaining imprecise estimates of the NAIRU, even when the analyst applies the state of the art statistical techniques.

Concerning the first of the questions, the application of the standard GLS-class of unit root tests leads to conclude in favor of hysteresis in the unemployment rates. However, when allowing for the presence of up to two structural breaks in unit root testing the conclusions reverse: the null hypothesis of hysteresis can be rejected for all the countries analyzed. Thus, the empirical evidence points to the fulfilment of the shifting natural rate paradigm. This is in accordance with the experience of these countries in the past decade, since their economies suffered the transition process from communism to the capitalist system.

In a second stage, our analysis has allowed to approximate a measure of the NAIRU that characterizes these economies. The applied statistical techniques pay special attention to two of the sources of uncertainty mentioned in the literature. First, the uncertainty arising from not knowing the parameters of the model, that we address by computing confidence intervals that provide reliable and accurate measures of this imprecision. Second, the uncertainty related to the stochastic nature of the NAIRU, that we treat by allowing for breaks that occur at unknown dates and that we consistently estimate. In addition, we calculate the NAIRU not only for the point estimates derived from the models, but also for the confidence intervals. We also report the results using two alternative estimation methods, the downward-biased OLS estimator and the median-unbiased estimator. Correcting for the OLS bias produces larger autoregressive parameters and affects the NAIRU estimates. However, independently of the estimation method, the main conclusion is that these NAIRU univariate-based estimates are measured quite precisely.

The empirical difficulties and inaccuracies related to NAIRU measurement are well documented in the literature. Although this limits its usefulness as a policy tool in terms of macroeconomic policy-making, it provides relevant information that can still be used jointly with a range of other indicators. Additionally, measuring the NAIRU can be potentially relevant for the microeconomic analysis of the labor markets. Thus, NAIRU estimates for the Accession countries can be used as a measure of cross-country differences in the functioning of labor markets.

From the estimation of the structural breaks and the sum of the autoregressive coefficients, we can draw several features describing unemployment dynamics in transition countries.

First, from the point estimates of the breakpoints, a first broad clustering of the countries in two groups can be made in terms of the number of breaks. The first one consists of the small countries (Baltic countries and Malta) plus Poland, that display from 3 to 4 structural changes. The second group includes medium-size economies (Czech Republic, Hungary, Slovakia and Slovenia) that only experience one or two breaks. The reason behind these differences may be closely related to the degree of openness, larger in the case of the tiny economies and, then, more exposed to external shocks. In the case of Poland, the higher number of breaks is due to the special transition strategy followed from the very beginning of the nineties. A second feature derived from the analysis is that the estimated breaks are associated with political or institutional events of relevance in the transition process. Some of these events are common to all the countries (such as the Russian crisis), whereas other are idiosyncratic.

Third, and related to the persistence measures, the Baltic states (Estonia, Latvia and Lithuania) have low autoregressive coefficients and, thus, lower persistence and shorter half-lives, whereas a second group of countries, less homogeneous, consisting of the Czech Republic, Malta, Poland, Slovakia and Slovenia, show higher persistence. The differences found across countries may be explained, in this case, by the institutional framework in the labor market. Those countries that enjoy more social benefits and higher minimum wages are those which exhibit the largest half-lives. Also related to this last feature, although the Baltic countries turn out to have a relatively high NAIRU estimate, it has been decreasing in the last regimes, maybe due to the ease of adjustment after a shock. Finally, some of the previous conclusions are reinforced when analyzing the evolution of the modified Phillips curves, that have shifted downwards in all the countries, with special intensity in the Baltic countries.

Figure 1. Unemployment rates and NAIRU OLS-based estimates of the CEECs

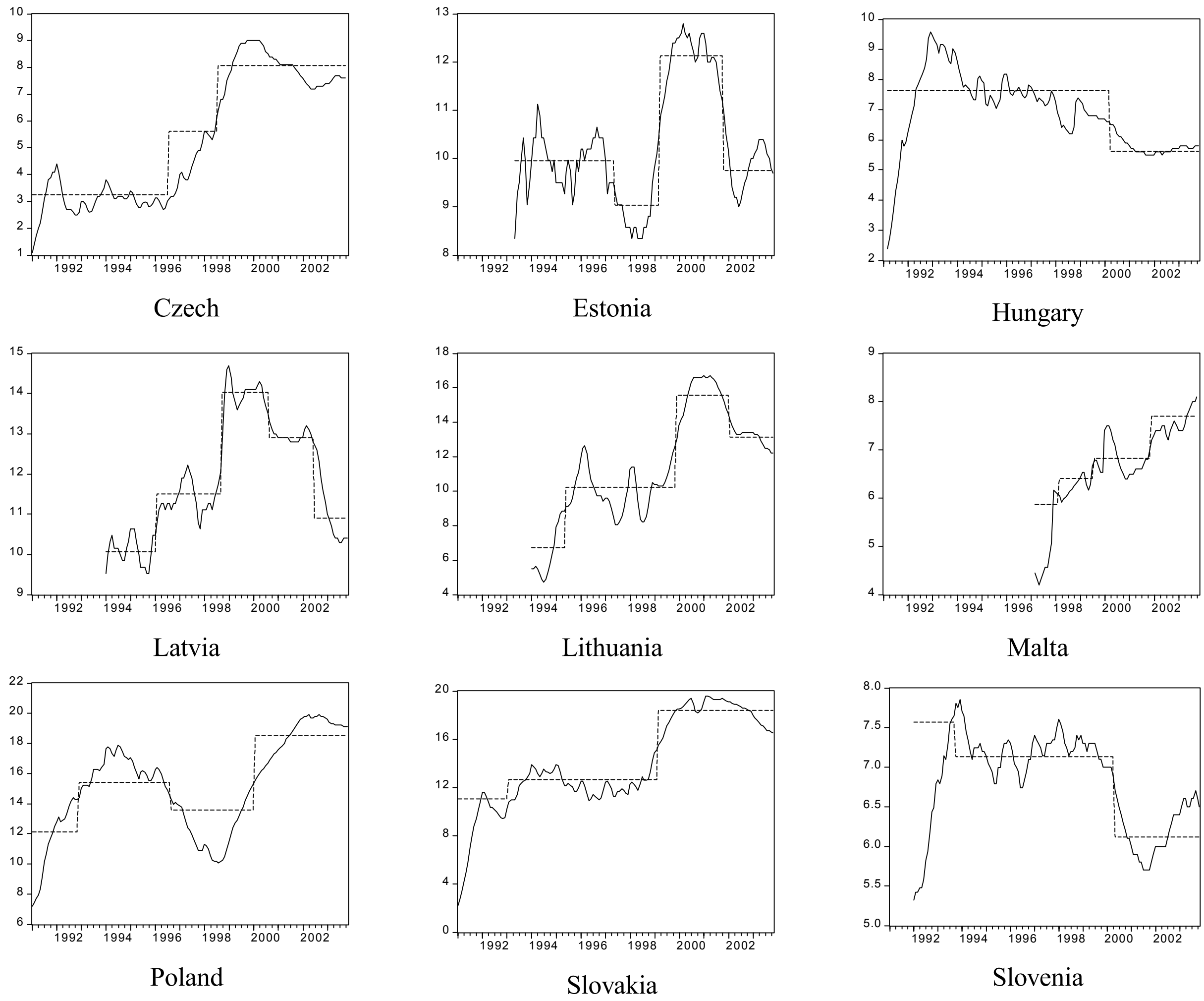
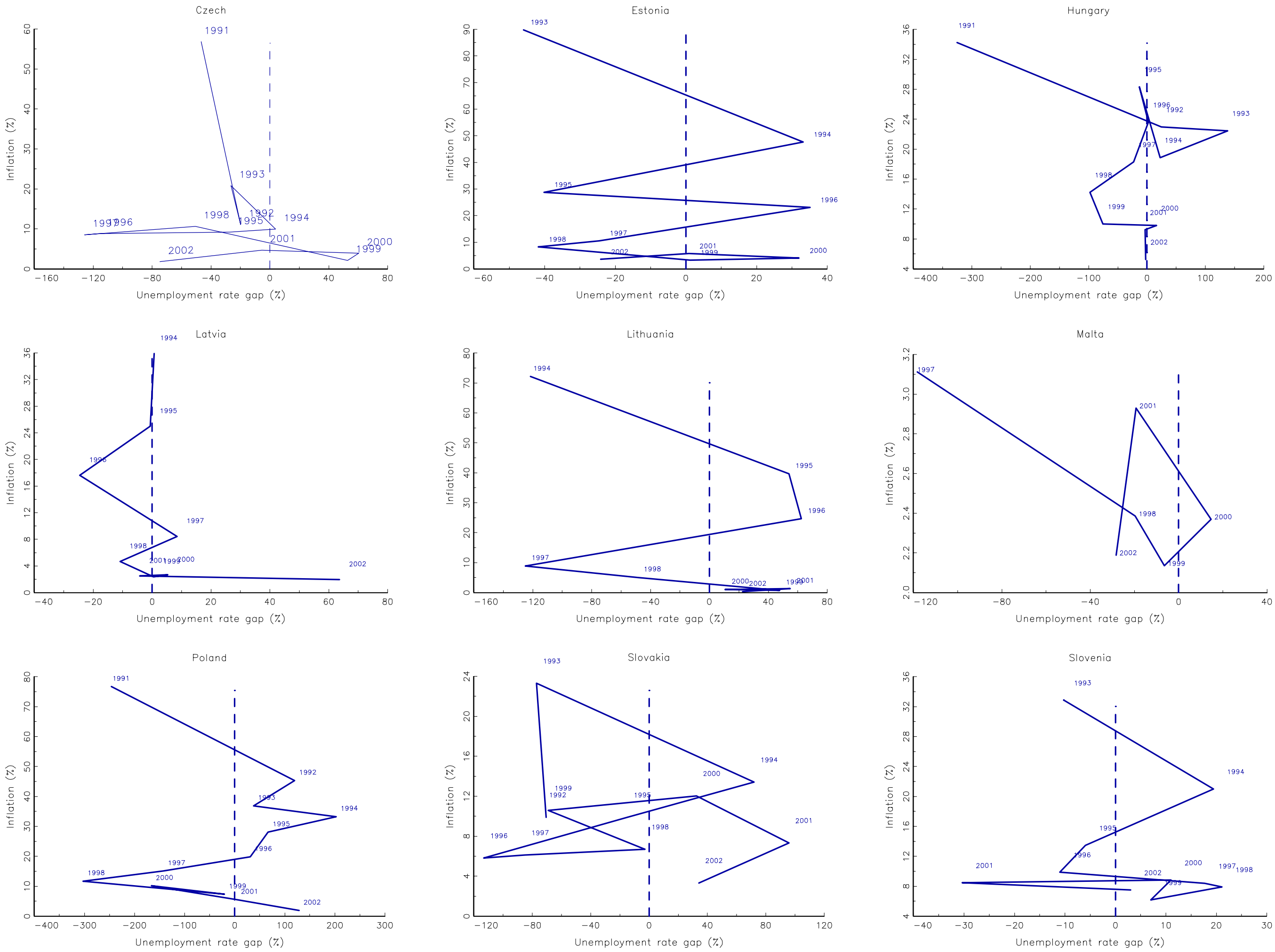


Figure 2. Unemployment rate gap and inflation



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