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Fiscal shocks and budget balance persistence in the EU countries from Central and Eastern Europe*

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Abstract: This paper analyses the time series properties of the fiscal balance in the 10 EU countries from Central and Eastern Europe. The persistence of shocks in the variable has been analysed by means of unit root tests that account for the possibility of non-linearities and structural changes. The results of the linear and non-linear unit root tests find only mild evidence in favour of the stationarity hypothesis, with asymmetric effects present in a few cases. After controlling for structural changes in the data generation process, the results point to stochastic stationarity of the series. Thus, in spite relatively steady headline figures, the public balance processes exhibit substantial instability in the EU countries from Central and Eastern Europe.

J.E.L. Classification: C32, E24
Key words: Unit roots, structural breaks, budget balance, EU.

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1. Introduction

This paper provides a systematic analysis of the time series properties of the fiscal balance in the 10 countries from Central and Eastern Europe that joined the European Union in 2004 or 2007. The focus is on the persistence of the fiscal balance to shocks, and in particular the conditions under which the fiscal balance exhibits stationarity. A range of unit root tests are applied, including tests that take non-linearities and structural breaks into account. The data are quarterly and cover the period from the first quarter of 1999 to the second quarter of 2010.

The background for the analysis is the prevalence of the shocks that affected the fiscal outcome in the EU countries from Central and Eastern Europe during the first decade of the 21st century. After communism and central planning had been shed, the 1990s was a decade of macroeconomic instability, administrative reforms and structural change. From the end of the 1990s the public finance institutions and structures were approaching those of the EU countries of Western Europe (Fabrizio and Mody 2010). Nonetheless, the countries still experienced a large number of shocks in the form of business cycle fluctuations, interest rate changes, discretionary policy changes, etc. The global financial crisis, manifesting itself in the third quarter of 2008, is a case in point. The crisis put a new spotlight on fiscal policies and debt management in Central and Eastern Europe as several of the countries faced financing difficulties and received IMF-led bailout programmes.

This paper examines to which extent shocks have had persistent effects on the budget balance in the 10 EU countries from Central and Eastern Europe. If shocks are persistent, short term changes in the fiscal balance will affect the fiscal balance in the longer term; a shock affecting the fiscal balance negatively might be especially troublesome if the fiscal balance is not a variable stationary. Clearly, the persistence of shocks is important for fiscal policy management.

All 10 countries in the sample joined the European Union during the sample period. Estonia, Latvia, Lithuania, Poland, the Czech Republic, Slovakia, Hungary and Slovenia joined in May 2004, while Romania and Bulgaria joined in January 2007. EU members are obliged to abide by the Stability and Growth Pact, which stipulates, inter alia, that the headline deficit cannot exceed 3 percent of gross domestic product (GDP) except in exceptional circumstances; the European Commission will initiate an Excessive Deficit Procedure (EDP) against a country if this criterion is deemed to have been violated. An EDP furthermore implies that the country is in violation of the fiscal criterion of the Maastricht Treaty, one of the convergence criteria which must be satisfied for a country to join the euro area. By January 2011 three of the countries – Slovakia, Slovenia and Estonia – had gained entry to the euro area. The institutional requirements on the fiscal balance imposed by EU membership add impetus to an analysis of the time series properties of this variable.

The empirical results may also possess some lessons for the long-term viability of the public finance stance. Stationarity of the budget balance is a *sufficient condition* for fiscal sustainability, when fiscal sustainability is taken to mean that continuation of the historical performance will not result in debt accumulation breaching the intertemporal budget constraint (the transversality condition) – given that interest payments are sufficiently independent of the accumulation of debt (Bohn 2007). It may, all the same, be argued that equating debt sustainability and the fulfilment of the transversality condition is a rather abstract exercise. [1]

It is important to emphasise, moreover, that stationarity of the budget balance is *not* a *necessary condition* for fiscal sustainability interpreted as the transversality condition being satisfied. In fact, the budget balance can be integrated of any arbitrary order and still satisfy the intertemporal budget constraint; the discount factor in the transversality condition dominates in the limit the accumulation of debt irrespective of the order of integration of the debt (Bohn 2007). In spite of this, stationarity of the budget balance is sometimes taken to epitomise a *strong form* of budget balance sustainability although the exact implication of this strong form is not clearly defined (Bohn 2007, Holmes et al. 2010).[2]

[1] The fulfilment of the intertemporal budget constraint or transversality condition is largely a rather academic conception of debt sustainability. First, the transversality condition relates to events infinitely far into the future, which makes it difficult to assess its applicability for debt management in a practical framework. Axel Leijonhufvud states this in the following explicit way: "[W]e had better remember that transversality conditions at an infinite horizon are not to be taken seriously" (Leijonhufvud 2011, p. 5). Second, the assumption of interest rates being "sufficiently independent" of the fiscal outcome is unlikely to be satisfied in practice as the interest rate may react strongly to both the budget balance and the debt stock. Finally, the implicit assumption of unchanged behaviour infinitely far into the future is also unsatisfactory. Guides to practical fiscal management will typically not include a reference to the transversality condition, but instead focus on the accumulation and financing of debt in the short or medium term (Balassone et al. 2010; Budina and van Wijnbergen 2007; EC 2010, sec. I.4).

[2] Fiscal long-term performance has also been studied using various multivariate methods (Afonso and Rault 2008, Prohl and Schneider 2006, Haug 1995). One approach entails testing for co-integration between the budget balance and the debt stock; another approach entails testing for a co-integrating (1,1) vector between expenditures and revenues. These approaches have many of the same limitations as univariate methods as regards their implications with respect to the transversality condition.

A large number of empirical studies consider the persistence of the budget balance and other fiscal variables, mainly in countries in North America and Western Europe. In a seminal paper Hamilton and Flavin (1986) consider the time series properties of the budget balance and the debt stock of the US federal government based on annual data from 1960 to 1984. One of their main findings is that a modified budget balance measure exclusive of interest payments and expressed in real terms is a stationary variable. This result – in combination with the finding that also the debt stock measure is stationary – is interpreted as an indication of an overall prudent fiscal policy stance in the sample period. Other studies of the US federal deficit have presented a more mixed picture. Wilcox (1989) examines the same data used in Hamilton and Flavin (1986) and finds that there are signs of parameter instability and structural breaks in the data.

Trehan and Walsh (1988) argue that the assessment of budget sustainability should be based on the time series properties of the real value of the *overall* budget balance, inclusive of interest payments and seigniorage revenue. Using Dickey-Fuller tests on annual US data for the very long period 1890-1986, they find that the budget balance measure is stationary. Using a shorter dataset from 1960 to 1984, Trehan and Walsh (1991) cannot reject the hypothesis that the overall budget balance exhibits a unit root, but they ascribe this to the low power of the test because of a dataset with only few observations.

Turning to analyses using European data, Vanhorebeek and Rompuy (1995) examine whether the overall budget balance as a percentage of GDP is stationary using data for eight West European countries from 1970 to 1994. The finding is that the budget balance is stationary for Denmark, France and Germany (although the results vary somewhat depending on whether or not a trend is included), but not for the other countries included in the sample. Caporale (1995) finds that the primary balance as a percentage of GDP is stationary for most of the West European countries examined. Greiner and Semmler (1999) test for unit roots in both the primary and the overall budget balances in Germany using annual data for 1955-1994. The authors fail to reject non-stationarity even when the period after German reunification is excluded. Holmes et al. (2010) test for stationarity of the budget balance in West European EU countries using annual data from 1971 to 2006. The estimations are based on panel data methodology, but take into account cross-sectional dependence and allow for a structural break. The budget balance is found to be stationary, whether or not the estimations include a structural break.

This paper contributes to the literature testing fiscal policy persistence in a number of ways. First, the paper is to our knowledge the first to provide a comparative study of the persistence of the budget balance to shocks in Central and Eastern Europe. Second, the paper exploits that quarterly budget data have recently become available for EU countries. The use of quarterly data ensures that a sufficient number of observations are available for unit root testing in each country individually, i.e. there is no need to use panel data methods which presume cross-sectional homogeneity. Third, the paper employs a wide range of testing methodologies allowing for *inter alia* non-linear adjustment and structural breaks.

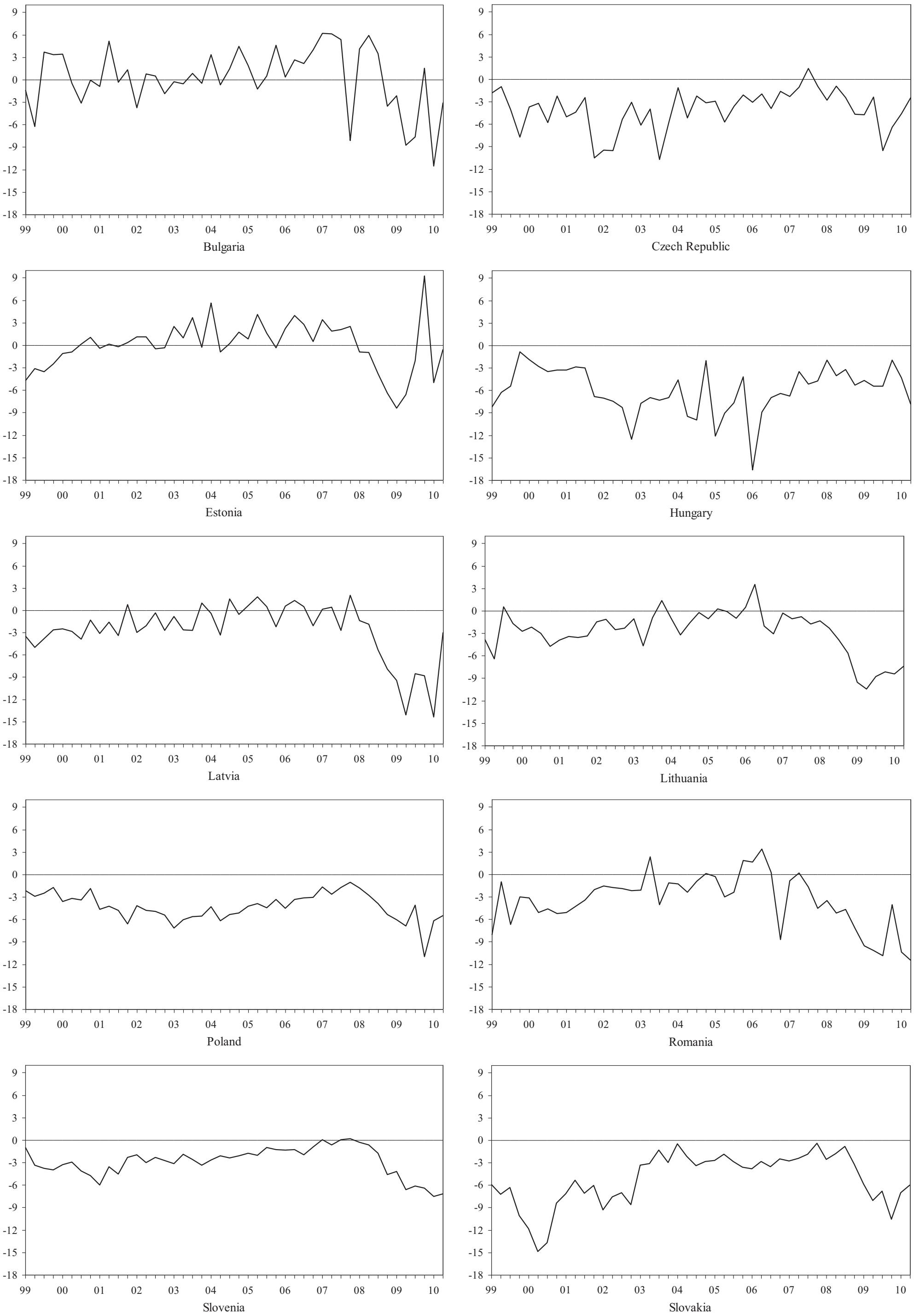
The rest of the paper is organised as follows. Section 2 presents the data used and provides a brief introduction to the fiscal policies in the 10 EU countries from Central and Eastern Europe. Section 3 sets out the econometric methodology used for testing for the order of integration. Section 4 provides the empirical results for the tests on the fiscal balance in the 10 countries. Finally, Section 5 summarises the paper.

2. A first look at the data

The paper uses quarterly data on the overall budget balance as a percentage of GDP for the general government in each of the 10 EU countries from Central and Eastern Europe. Data are downloaded from the electronic database of Eurostat.[3] The sample covers 1999:1 to 2010:2. The quarterly budget balance data exhibit substantial seasonal variation and the time series have therefore been seasonally adjusted using the X-12 filter (assuming additive seasonality, since multiplicative methods cannot be applied to data with negative observations). Figure 1 shows the seasonally adjusted budget balance for the 10 EU countries from Central and Eastern Europe.

[3] Variable code = gov_q_ggnfa (non-financial accounts for general government), downloaded 28 Nov. 2010.

Figure 1: Budget balance in percent of GDP, seasonally adjusted, quarterly



Source: Eurostat

Figure 1 reveals that there is substantial heterogeneity across the 10 countries. The Central European countries – the Czech Republic, Hungary, Poland, Slovenia and Slovakia – have generally exhibited moderate or, in some cases, substantial negative budget balances. Latvia, Lithuania and Romania have had broadly balanced budgets, while two countries – Bulgaria and Estonia – stand out for having had quarterly budget surpluses in most quarters within the sample period. The different fiscal outcomes may partly be explained by the constraints afforded by different exchange rate systems; countries with floating exchange rate systems have generally pursued laxer and more accommodating fiscal policies than countries with fixed exchange rate systems (Lewis 2007). The overall picture is, however, that the fiscal policy outcomes in the EU countries in Central and Eastern Europe have been prudent, in particular when compared to the outcomes in EU countries in Western Europe (Staehr 2008). This finding is partly the result of rapid GDP growth and a low debt stock in most EU countries from Central and Eastern Europe, but also of deliberate measures for budget consolidation at different stages (Afonso 2006, Staehr 2010).

Figure 1 also reveals that the budget balance exhibits substantial variation across individual countries in the sample period. Two main observations warrant attention. First, many countries experienced an improvement in the overall budget balance during the first part of the sample, which represents the run-up to EU accession. This pertains to the Czech Republic, the Baltic States, Romania, Slovenia and Slovakia. A number of studies have examined the extent to which the improvement could be attributed to governments "readying" themselves for EU membership and the obligation to fulfil the fiscal criteria of the Stability and Growth Pact and the Maastricht Treaty. Lewis (2007) answers in the affirmative, Annett (2006) and Collignon (2006) find that there was no discernible effect, while Berger et al. (2007) argue that the different developments across the countries joining the EU reflect their political bargaining power vis-à-vis the West European EU countries.

Second, the budget balance deteriorated in almost all the sample countries in the wake of the global financial crisis. In almost all the countries in the sample, the economic downturn and counter-cyclical policy measures led to substantial budget deficits from 2008. Judged on the overall budget balance, Estonia and Hungary stand out as exceptions. Estonia faced a deep downturn that already started in 2008, and initially experienced a substantial deterioration of the fiscal balance. Discretionary policy measures, however, reversed this picture in 2009 as the country sought to satisfy the deficit criterion of the Maastricht Treaty and gain access to the euro area. Hungary faced a liquidity squeeze in the fourth quarter of 2008 and had to seek an IMF-led bailout, which was followed by a substantial tightening of fiscal policy. Latvia and Romania also received bailouts in the fourth quarter of 2008 and the first quarter of 2009 respectively, but implemented less aggressive tightening of their fiscal policies (Staehr 2010, EC 2010).

3. Econometric methods

The econometric approach in this paper implies the application of a battery of unit root tests with the aim of determining the order of integration of the fiscal balance (in percent of GDP) in the 10 EU countries from Central and Eastern Europe. We employ different unit root tests assuming data generating processes that are linear, non-linear or contain structural breaks.

Linear unit root tests

These tests represent upgraded versions of previously existing unit root tests which are based upon linear equations. The tests applied in this paper are due to Ng and Peron (2001), who propose some modifications to the Phillips (1987) test (MZa), Phillips and Perron (1988) (MZt), Bhargava (1986) (MSB), and the Point Optimal Test by Elliot, Rothenberg and Stock (1996) (MPT). This is done by combining a Modified Information Criterion for the lag length and a Generalised Least Squares method for detrending the data.

Non-linear unit root tests

Although the Ng and Perron (2001) tests are improved versions of previous tests, they are based on a linear data generation process (DGP). According to the recent literature on non-linear modelling in time series econometrics, unit root tests based upon linear DGPs may be biased towards committing Type II errors, if the true DGP is in fact non-linear, see e.g. Kapetanios et al. (2003). This implies that the test may fail to reject the null if the process is non-linear but globally stationary. In the case of the fiscal balance as a percentage of the GDP, this could be the case. For this variable it may be sensible to consider the existence of a certain threshold of values where deviations may not trigger any type of policy action, as those deviations are too small to apply a costly fiscal policy. Statistically, this implies that the variable may behave as an I(1) process within the aforementioned threshold. This is known as the inner regime. But, if deviations from the equilibrium are significant and posse dangers to the fiscal balance, authorities may apply policies to reduce those deviations, and therefore the variable may behave as a mean reverting and stationary process in the outer regime. This state-dependent two regime process will imply that the further the fiscal balance deviates from the equilibrium, the faster will be the mean reversion. Hence, we will observe an autoregressive parameter which depends on the size of the deviations, relaxing the assumption of a constant parameter.

This type of non-linear DGP can be modelled as an exponential smooth transition autoregression (ESTAR) process, which allows us to test also for unit roots. This is the idea behind the KSS test. The authors propose a unit root test which prescribes a globally stationary ESTAR process under the alternative hypothesis. The unit root hypothesis can be tested against the alternative of a globally stationary ESTAR process using the regression

$$y_t = \varphi y_{t-1} + \phi y_{t-1} F(\theta; y_{t-1}) + \varepsilon_t , \tag{1}$$

where the error ε_t is $IID(0, \sigma^2)$ and $F(\theta; y_{t-1})$ is the transition function, which is assumed to be exponential (ESTAR) and takes the form

$$F(\theta; y_{t-1}) = 1 - exp\{-\theta y_{t-1}^2\} , \tag{2}$$

with $\theta > 0$. It is common to combine equations (4) and (5), and rewrite them as

$$\Delta y_t = \alpha y_{t-1} + \gamma y_{t-1}(1 - exp\{-\theta y_{t-1}^2\}) + \varepsilon_t . \tag{3}$$

As aforementioned, assuming that the process behaves as an I(1) process in the inner regime implies imposing $\alpha = 0$ in equation (3), although the process is globally stationary. The null hypothesis H_0: $\theta = 0$, cannot, however, be directly tested, since in practice the parameter γ cannot be identified under the null. KSS propose the use of a Taylor approximation of equation (6) taking the form

$$\Delta y_t = \beta y_{t-1}^3 + error . \tag{4}$$

Testing H_0: $\beta = 0$ against H_1: $\beta < 0$ is equivalent to testing for unit roots in the outer regime in equation (3). Equation (4) may incorporate lags in order to control for autocorrelation in the residuals with a lag length possibly chosen by means of an information criterion. The KSS test is applied to the demeaned data and no other deterministic component is considered in the DGP.

The non-linear function used by KSS in order to take non-linearities into account, assumes that shocks have symmetric effects upon the variable, i.e. the sign of the shocks does not matter, only the size for the absolute value of the effect. However, for many economic variables this assumption may be too restrictive. Hence, Sollis (2009) proposes a KSS-type test, which distinguishes between asymmetric or symmetric effects under the alternative hypothesis, i.e. the speed of mean reversion will be different depending on the sign of the shock and not only its size, unlike in the ESTAR models. This asymmetric ESTAR model (AESTAR) is defined as

$$\Delta y_t = G_t(\gamma_1, y_{t-1})\{S_t(\gamma_2, y_{t-1})\rho_1 + (1 - S_t(\gamma_2, y_{t-1}))\rho_2\}y_{t-1} + \varepsilon_t\ , \qquad (5)$$

where $G_t(\gamma_1, y_{t-1}) = 1 - \exp(-\gamma_1(y_{t-1}^2))$, with $\gamma_1 \geq 0$, and $S_t(\gamma_2, y_{t-1}) = \{1 + \exp(-\gamma_2 y_{t-1})\}^{-1}$, with $\gamma_2 \geq 0$. Equation (5) may incorporate lags of the dependent variable to control for autocorrelation.

The null hypothesis of unit root can be specified as H_0: $\gamma_1 = 0$. However, under the null hypothesis, the parameters γ_2, ρ_1 and ρ_2, cannot be identified. In order to get around this problem, Sollis (2009), by means of Taylor approximations, proposes to test for unit roots in this non-linear framework using the auxiliary equation

$$\Delta y_t = \beta_1 y_{t-1}^3 + \beta_2 y_{t-1}^4 + error. \qquad (6)$$

It follows that testing for unit roots in model (6) implies testing the nul hypothesis H_0: $\beta_1 = \beta_2 = 0$, by means of an F-test. Given that the order of integration of the residuals is unknown, Sollis (2009, p. 121) provides appropriate critical values for this F-type test. The null hypothesis of symmetric ESTAR versus the alternative of AESTAR, can be tested once H_0: $\beta_1 = \beta_2 = 0$ has been rejected. In this case, testing for the null hypothesis of symmetric ESTAR implies testing the hypothesis H_0: $\beta_2 = 0$ by means of a standard hypothesis test (t-test, F-test or LM-test), for which standard critical values apply.

Unit root tests allowing for structural breaks

The unit root tests discussed above may suffer from power problems when there are structural breaks in the DGP. In this case, these tests may incorrectly conclude that the variable contains a unit root, when in fact it is stationary around a broken or shifting drift, i.e. the test may over over-non-reject the null hypothesis (see e.g. Perron, 1989).

Since the seminal contribution in the field of structural breaks by Perron (1989), a number of authors have tried to provide with improved tests in terms of power. In a more recent contribution, LS develop a unit root test which takes into account the possibility of two structural changes. According to these authors, earlier unit root tests with structural changes, such as those by Zivot and Andrews (1992) and Lumsdaine and Papell (1997), may provide misleading conclusions when the unit root hypothesis is rejected. Rejecting the null hypothesis implies that the series presents structural changes, but the variable can be either stationary or a unit root process. This means that the rejection of the null hypothesis does not always imply that the series is trend-stationary, because the null hypothesis of those earlier unit root tests with structural breaks does not incorporate the possibility of structural changes. The LS test can be performed by estimating the following equation:

$$\Delta y_t = \delta' \Delta Z_t + \phi \bar{S}_{t-1} + u_t \,, \tag{7}$$

where Z_t is a vector of exogenous variables, $\bar{S}_t = y_t - \bar{\Psi}_1 - Z_t \bar{\delta}$ for t = 2, ... T, $\bar{\delta}$ is a vector of estimated values of δ from a regression of Δy_t upon ΔZ_t, and $\bar{\Psi}_1 = y_1 - Z_1 \bar{\delta}$, where y_1 and Z_1 are the initial values for y_t and Z_t. To define the null and alternative hypotheses, let us consider the DGP given by

$$y_t = \delta' Z_t + e_t \,, \tag{8}$$

where $e_t = \beta e_{t-1} + \varepsilon_t$ and $\varepsilon_t \sim NIID(0, \sigma^2)$. For a two-break model, we can define $Z_t = [1, D_{1t}, D_{2t}]'$, where $D_{jt} = 1$ for $t \geq T_{Bj+1}$, j = 1, 2, and 0 otherwise. T_{Bj} is the date of the breaking point. The null and alternative hypotheses can then be defined as H_0: $y_t = \alpha_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + \vartheta_{1t}$ and H_1: $y_t = \alpha_1 + d_1 D_{1t} + d_2 D_{2t} + y_{t-1} + \vartheta_{2t}$, where ϑ_{1t} and ϑ_{2t} are stationary error terms, B_{1t} and B_{2t} = 1 for $t = T_{B1} + 1$ and $t = T_{B2} + 1$, respectively, and 0 otherwise. Thus, the unit root hypothesis is H_0: $\phi = 0$ and the test statistics are given by $\bar{\rho} = T \bar{\phi}$ and τ, the latter being the t-statistic associated with ϕ. The two-break unit root test selects the time breaks endogenously by minimising the test statistic, i.e. the test selects the model which provides more evidence against the null hypothesis.

4. Empirical results

Table 1 shows the results of the Ng-Perron, KSS and Sollis tests, the upper panel for the time sample 1999:1-2010:2. Under the assumption of a linear DGP, the null of a unit root can be rejected in favour of stationarity for only Slovakia and Slovenia, cf. the columns labelled MZa, MZt, MSB and MPT.

Table 1: Ng-Perron, KSS and Sollis (2009) unit root test results

1999:1-2010:2						
	MZa	**MZt**	**MSB**	**MPT**	**KSS**	**Sollis**
Bulgaria	-2.914	-1.174	0.402	8.322	0.317	0.146
Czech Rep.	-4.468	-1.491	0.333	5.489	-0.092	0.014
Estonia	-3.333	-1.277	0.383	7.338	-3.304**	9.222**
Hungary	-3.210	-1.261	0.392	7.624	-0.183	3.000
Latvia	-2.681	-1.157	0.431	9.137	-2.803*	4.041*
Lithuania	-2.520	-0.974	0.386	9.015	-0.810	4.040*
Poland	-3.439	-1.222	0.355	7.087	-1.085	1.018
Romania	-0.397	-0.249	0.629	24.078	-0.352	0.173
Slovenia	-69.841**	-5.777**	0.082**	0.637**	-0.434	0.226
Slovakia	-6.621*	-1.818*	0.274*	3.703*	-2.090	3.425
1999:1-2008:2						
	MZa	**MZt**	**MSB**	**MPT**	**KSS**	**Sollis**
Bulgaria	1.128	1.547	1.371	127.961	-0.323	0.188
Czech Rep.	-1.747	0.501	13.144	0.093	0.007	-0.980
Estonia	-0.980	-0.695	0.709	24.751	-2.507	3.524
Hungary	-2.742	-1.103	0.402	8.697	-0.124	3.160
Latvia	-1.072	-0.728	0.679	22.681	-0.259	3.489
Lithuania	-1.199	-0.772	0.643	20.346	-3.013**	4.831*
Poland	-1.755	-0.929	0.529	13.845	-2.003	2.009
Romania	-1.049	-0.720	0.686	23.170	-1.027	0.643
Slovenia	-1.518	-0.810	0.533	14.875	-2.540	3.533
Slovakia	-4.383	-1.411	0.321	5.701	-1.842	2.275

Note: The order of lag to compute the tests has been chosen using the modified AIC (MAIC) suggested by Ng and Perron (2001). The Ng-Perron tests include an intercept, whereas the KSS test has been applied to the de-meaned data. The symbols * and ** mean rejection of the null hypothesis at the 10% and 5% significance levels respectively. The critical values for the Ng-Perron tests and F-test have been taken from Ng and Perron (2001) and Sollis (2009) respectively, whereas those for the KSS have been obtained by Monte Carlo simulations with 50,000 replications.

Critical Values

	MZa	**MZt**	**MSB**	**MPT**	**KSS**	**Sollis**
5%	-8.100	-1.980	0.233	3.170	-2.886	4.886
10%	-5.700	-1.620	0.275	4.450	-2.603	4.009

If we allow for the possibility of a non-linear model, we also find evidence against the null for Estonia and – at the 10% level of significance – for Latvia and Lithuania, cf. the columns labelled KSS and Sollis. Given stationarity in the non-linear model for the three Baltic States, we can also test whether the adjustment is symmetric or asymmetric. As discussed in Section 3, this is done by testing H_0: $\beta_2 = 0$ in equation (6). In the three cases, the null of a symmetric ESTAR can be rejected at the 10% significance level (results available upon request). In other words, for these countries the effect of a shock with the same magnitude, but different sign, will have a different effect on the speed of adjustment towards equilibrium.

The results exhibit an interesting pattern. Slovenia, Slovakia and the three Baltic States were found to have stationary budget balances when the DGP is assumed to be, respectively, linear or non-linear. In 2007 Slovenia became the first of the EU countries from Central and Eastern Europe to join the euro area, followed by Slovakia in 2009 and Estonia in 2011. The two other Baltic States also made preparations to join the euro zone immediately after gaining EU membership, among other things by tying their exchange rates to the euro within the ERM II. The other countries in the sample have not joined the ERM II and the countries, with the exception of Bulgaria, have generally pursued a relatively active fiscal policy as discussed in Section 2. The upshot is that the unit root tests detect a pattern of time series properties which mirrors the countries' "proximity" to the euro area. Whether this pattern is a coincidence or it pertains to causal factors is left for further research.

Table 1 also reports the results of the Ng-Perron, KSS and Sollis (2009) unit root tests for the period 1999:1-2008:2 where the last part of the sample comprising the global financial crisis has been omitted. The results provide more evidence in favour of the null hypothesis of a unit root. This may be caused by structural breaks changing the time series properties of the data across the sample, but it may also be the result of the unit root tests suffering from low power in short samples.

Finally, Table 2 displays the results of the Lee and Strazicich (2003) unit root tests with one and then two structural breaks in both the drift and the trend. It is important to test for the presence of structural breaks, given that, as already mentioned, misspecifications in the deterministic components of the auxiliary regressions may lead to a tendency for excessive non-rejection of the null hypothesis of a unit root. The discussion in Section 2 suggested that structural breaks may be likely, given the possible effects of the run-ups to EU and euro area membership, the global financial crisis and other structural changes in these fast-growing economies.

Table 2: Lee and Strazicich (2003) unit root tests results with one or two structural breaks

	One structural break		Two structural breaks		
	TB1	**Statistic**	**TB1**	**TB2**	**Statistic**
Bulgaria	2007:3	-3.643	2002:2	2007:3	-8.674**
Czech Rep.	2003:4	-5.535**	2003:4	2007:2	-6.330**
Estonia	2008:2	-6.305**	2008:1	2009:2	-8.485**
Hungary	2006:1	-3.581	2004:3	2008:1	-5.394**
Latvia	2007:3	-4.043	2007:4	2009:2	-8.486**
Lithuania	2008:2	-5.659**	2005:3	2008:3	-7.056**
Poland	2005:4	-4.681**	2002:4	2008:2	-8.639**
Romania	2005:3	-6.867**	2006:3	2007:3	-6.982**
Slovenia	2003:4	-3.422	2002:4	2006:2	-5.240**
Slovakia	2002:4	-4.348*	2002:4	2008:4	-5.780**

Note: TB1 and TB2 stand for the time period for the first and second break, respectively. The model considers a break in the drift and in the trend. The symbols * and ** mean rejection of the null hypothesis at the 10% and 5% significance levels respectively. The critical values have been obtained from Lee and Strazicich (2003, 2004):

Critical values

	TB1 test	**TB1 and TB2 test**
5%	-4.50	-5.28
10%	-4.20	-4.98

With one structural break (the left hand side of the table), the unit root hypothesis can be rejected at conventional levels of statistical significance for six countries, namely the Czech Republic, Estonia, Lithuania, Poland, Romania and Slovakia. For Estonia and Lithuania the break points appear in the second quarter of 2008, just prior to the global financial crisis. These two countries, along with Latvia, were severely affected by the global financial crisis with accumulated declines in GDP in 2008-09 of around 20 percent of GDP. For the Czech Republic, Poland, Romania and Slovakia the structural breaks fall in the middle of the sample period.

More interesting results emerge when a model with two structural breaks is considered (the right hand side of the table). The null hypothesis of a unit root is rejected in all cases. In most cases, the structural change identified coincides with one of the structural breaks identified in the model with two structural breaks.

With the possible exception of Slovenia, the *second structural breaks* correspond to the outbreak of the global financial crisis as signified by the bankruptcy of Lehman Brothers in the third quarter of 2008. The economic downturn and the worsened conditions in non-prime financial markets afforded substantial public finance challenges in Central and Eastern Europe. A range of policy measures were introduced in the form of stimulus programmes in some countries and discretionary tightening in other countries (EC 2010, Staehr 2010).

The *first structural break* appears to be more widely spread across the time sample than the second structural break. In many cases the break appears in 2002-03, a period in which the countries generally saw rapid economic growth partly driven by capital inflows and expanding trade in anticipation of the countries joining the EU. An interesting picture emerges for Estonia and Latvia for which the two structural breaks are only 5-6 quarters apart (the first break appears in 2007:4 and 2008:1, and the second in 2009:2). Both countries were already in recession in the beginning of 2008, i.e. prior to the bankruptcy of Lehman Brothers, and subsequently faced very deep recessions.

The findings in Table 2 suggest that it is important to take structural changes into account when assessing the time series properties of the fiscal balance. Although asymmetries seem to be present in the fiscal balance processes in some of these countries, it appears that the unit root found with the previous tests may be due to the existence of important structural changes in the series. The results suggest that shocks mainly had persistent effects on the fiscal balance through structural changes affecting the mean or the trend of the variable.

Structural breaks have also been detected in other studies of the fiscal balance, but the result is in contrast to the main finding in Holmes et al. (2010) who consider the fiscal balance in West European EU countries and find that the budget balance is stationary whether or not the estimations include a structural break. The continued structural changes in the post-communist countries including EU and euro area membership and the global financial crisis constitute important events in the recent history of these countries.

5. Conclusions

This paper has analysed the degree of persistence of the fiscal balance as a percentage of GPD for the Central and Eastern European EU countries. The aim has been to attain measures of the longer-term susceptibility of the fiscal stance to shocks in the sample period and in this way gain some insights into the appropriateness of their fiscal policies. The analysis of the countries' fiscal balances is important, given that negative shocks may have long run effects and therefore be difficult to offset if the variables contain a unit root.

These considerations point to the need for a thorough analysis of the time series properties of the fiscal balance.

Turning first to the results from Table 1, non-stationarity of the fiscal balance is found in most cases. There is evidence of stationarity in Slovakia and Slovenia and, if a non-linear data generating process is assumed, also in Estonia, Latvia and Lithuania. The evidence towards stochastic stationarity of the series in these five countries may be related to these countries joining or seeking to join the euro area during the sample period, which may have made policymakers implement measures that would avoid persistent effects on the budget balance from shocks. It is noticeable, however, that a shortening of the time sample points to non-stationarity in essentially all cases.

The results from Table 2 suggest that the fiscal balance exhibited structural breaks with changes in the constant and trend of the fiscal balance. According to the results of the unit root tests with structural changes, the hypothesis of a unit root is rejected in all cases in favour of the stationarity of the fiscal balances when two distinct structural breaks are allowed in the data generating process.

The prevalence of structural breaks is arguably associated with the continued transition towards market economic structures, the preparation for EU and euro area membership and the global financial crisis, all of which are events that may have caused structural changes in the long run path of the series. The results underscore the need to take into account the possibility of changes in economic and structural fundamentals that may affect the fiscal balances in these countries.

The different results of the tests with and without structural breaks make it difficult to draw firm conclusions regarding the fiscal policy stance in the given countries. On the one hand, it appears that the fiscal balances tend to return to the equilibrium path after a shock when structural breaks are incorporated. On the other hand, the cause of these structural breaks is unexplained in simple univariate models and it is possible that some of the breaks may originate from fiscal shocks. The overall picture is therefore one of substantial instability of the public balance processes in the EU countries from Central and Eastern Europe, a result which differs markedly from the results for the EU countries from Western Europe obtained by Holmes et al. (2010).

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