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Time series behaviour of the real interest rates in transition economies

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Stationarity properties of real interest rates are examined for 21 transition economies. Owing to transaction costs and other frictions, it is quite plausible that we are dealing with potential non-linearities in the real interest rate. Therefore we examine stationarity of the real interest rate allowing for non-linearities and asymmetric adjustment with smooth structural change in the data generating process. Our findings suggest that taking account of non-linearities in the data generating process results in a rejection of the unit root null hypothesis for some countries which seem to be non-stationary according to conventional unit root tests. This finding points to the importance of allowing for both structural breaks and asymmetric adjustment in the real interest rate series of transition countries.

Keywords: real interest rate; transition economies; structural break; nonlinearity; unit root

JEL classification: C22, E43.

1. Introduction

The real interest rate is a key variable in many financial and macroeconomic models. In fact, as noted by Phillips (2005), the real interest rate plays a significant role in formulation of a wide range of economic models, including 'individual agent decision-making regarding investment, savings and portfolio allocations, options pricing models in finance, and the modern theory of inflation targeting in macroeconomics'. Therefore, it is not surprising that the stochastic property of the real interest rate has attracted intensive interest of economists both on the theoretical and empirical grounds. However, in spite of the fact that huge amount of work has been accumulated to date, no consensus has been achieved about the stochastic properties of the real interest rate.

The real interest rate was first formalised and examined by Fisher (1896, 1930). Fisher expressed the real interest rate as the difference between the nominal interest rate and inflation rate. According to the Fisher hypothesis, the nominal interest rate moves one-for-one with the expected inflation rate. Such a relationship between nominal interest and inflation rates implies that monetary policy has no effect on the real rate of interest, which, in turn, suggests that the real interest rate follows a stationary process. A wide range of economic models, including modern macroeconomic growth and real business cycle models, are built on the assumption that the real interest rate is stationary

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(e.g. Rose, 1988). Finance theory also assumes that the real interest rate is a stationary process (e.g. Cox, Ingersoll, & Ross, 1985; Vasicek, 1977). On the other hand, the Mundell-Tobin proposition (Mundell, 1963; Tobin, 1965) suggests that the nominal interest rates move less than one-for-one with the inflation rate because, in response to increasing inflation, households would hold less in money balances and more in other assets, which would reduce real interest rates. In addition, the so-called limited participation models (e.g. Christiano, Eichenbaum, & Evans, 1997; Grossman & Weiss, 1983; Rotemberg, 1984) also imply that monetary shocks reduce real interest rates, which suggests that shocks to the real interest rate may be persistent.

The empirical evidence on stochastic properties of real interest rates is also mixed. Stochastic properties of the real interest rate were first investigated by Rose (1988). He examined orders of integration of inflation rates and nominal interest rates, and found that the nominal interest rate contains a unit root whereas the inflation rate is stationary. Rose concluded that any linear combination of stationary inflation rate and non-stationary nominal interest rate should be non-stationary. Following Rose (1988), many studies examined long-run properties of the real interest rates with mixed results. Mishkin (1992), Wallace and Warner (1993), Crowder and Hoffman (1996), and Lai (1997) found that the real interest rate follows a stationary process for different countries. On the other hand, King, Plosser, Stock, and Watson (1991), Gali (1992), Shapiro and Watson (1988), Mıshkın and Sımon (1995), and Rapach and Weber (2004), among others, found that real interest rates are not stationary.

Empirical studies generally tested stationarity of real interest rates within linear framework. That is, the null hypothesis of a unit-root was tested against a linear stationary alternative. However, it is generally agreed that policy regime changes, transaction costs, risk premia and market imperfections may give rise to nonlinear dynamics in interest rates. Existence of transaction costs prevents economic agents from adjusting their portfolios continuously. However, if deviation from the equilibrium exceeds a certain threshold level, then benefits of adjustment exceed the cost and therefore economic agents act so as to move the system back towards the equilibrium level (e.g. Balke & Fomby, 1997). In addition, asymmetry in the central bank's preferences regarding the weight assigned to deviations of inflation the output gap from target levels might give rise to a nonlinear interest rate reaction function. It is argued that central bankers react more aggressively to positive (negative) deviations of inflation (output gap) when compared to reactions to negative (positive) deviations of inflation (output gap) from targeted levels (e.g. Dolado, Maria-Dolores, & Ruge-Murcia, 2004; Hasanov & Omay, 2008; Taylor & Davradakis, 2006). These findings suggest that the dynamics of the real interest rate are inherently nonlinear.

Balke and Fomby (1997), Caner and Hansen (2001), Kapetanios, Shin, and Snell (2003) and Bahmani-Oskooee, Kutan, and Zhou (2007), among others, have shown that linear models have low power in distinguishing between the unit root model and a nonlinear but stationary alternative. Therefore, in recent years, researchers applied nonlinear econometric methods to deal with possible nonlinear dynamics in real interest rates. Kapetanios et al. (2003) and Maki (2005) have provided evidence that the real interest rate behaves as a stationary process with asymmetric mean reversion for some industrialised economies. Million (2004) also found asymmetric mean reversion in the *ex post* real interest rate for the US data. Lanne (2006) found a common nonlinear component in the nominal interest rate and inflation in the long-run for the US data. Koustas and Lamarche's (2010) results indicate that the *ex post* real interest rate can be modelled as a three-regime self-exciting autoregressive (SETAR) process in Canada, France and Italy. Chang and

Chang (2012) and Chang (2012) find mean reversion in real interest rates for the G10 countries and for some G20 countries, respectively. Holmes, Dutu, and Cui (2009) have provided evidence that Australian and New Zealand real interest rates switch between regimes in mean, variance and persistence. Christopoulos and León-Ledesma (2007) found evidence of nonlinear cointegration between nominal interest rate and inflation rate in the case of the US. In addition to possible nonlinearities, many economic variables, including the real interest rate, may have undergone structural breaks due to major shifts in policy regimes or in underlying economic environment. In fact, Garcia and Perron (1996), Caporale and Grier (2000), Bai and Perron (2003), Rapach and Wohar (2005), Iskenderoglu (2011), Haug (2014) found structural breaks in real interest rates.

All the above-mentioned studies have considered mainly developed countries. A few studies that examine Fisher hypothesis in transition countries include Herwartz and Reimers (2006), Kasman, Kasman, and Turgutlu (2006), and Berument, Ceylan, and Olgun (2007)[1]. Herwartz and Reimers (2006) examine cointegration between nominal interest rates and inflation as well as stationarity of real interest rates for more than one hundred countries, including some transition countries, using panel co-integration techniques. They find that key parameters of the model are dependent on time or state. Kasman et al. (2006) examine Fisher's hypothesis using fractional co-integration analysis for 33 countries, including three transition countries, namely, the Czech Republic, Hungary, and Poland. They find that conventional cointegration tests do not support the Fisher relationship. However, they find fractional cointegration between nominal interest rates and inflation rates, providing evidence in support of the Fisher hypothesis. Berument et al. (2007) examine the Fisher relationship between interest rates and inflation expectations with and without inflation risk for the G7 countries and 45 developing economies, including eight transition countries (Albania, Armenia, Bulgaria, Hungary, Kazakhstan, Lithuania, Romania, and Russia). They find a positive and statistically significant relationship between interest rate and inflation only for Hungary, Kazakhstan, Romania and Russia. According to the results of Berument et al. (2007), however, in all these countries inflation rate increases less than one-for-one with expected inflation rate, which contradicts to the Fisher hypothesis in its strongest form.

Although transition economies have attracted huge interest from economists in recent years, stochastic properties of the real interest rate series in these countries have not yet been fully examined. These countries have been implementing massive economic reforms aimed at restructuring their centrally planned economies into market-based economy. Under socialist regime, prices had been set by the central authority that differed strikingly from the world prices and played no role in resource allocation. Similarly, all investment decisions were taken by the central authority and hence interest rates had no effect on savings and investment decisions. Financial institutions were not developed and did not play the role of financial intermediaries. However, during the transition period these countries have succeeded to liberalise their economies, establish market institutions, and integrate to the world economy. In addition, it is argued that some special features of transition countries such as high growth rates and/or high inflation rates and variable risk premium affect the real interest rate (Greenspan, 2004; Svensson, 2003). Therefore, examination of the stochastic properties of interest rates in these countries may improve our understanding of dynamics of interest rates. Furthermore, market determined interest rate plays a crucial role for the market oriented economies (Feltenstein, 1994). For example, non-stationary real interest rate has crucial implications on the effectiveness of monetary policy. Our main aim in this article is to contribute to the empirical literature by examining the stochastic properties of real

interest rates in transition countries. Another contribution of the article is that, in addition to conventional linear tests, we also apply newly developed test procedures that allow for both nonlinear dynamics and structural breaks in the data generating process. The results of the tests suggest that real interest rates in most of the countries under investigation follow a stationary process.

This article is structured as follows. In the Section 2 we present the relevant theory. In Section 3 we describe the econometric methodology. In Section 4 we present the data and the results of unit root tests. The last section concludes the article.

2. Real interest rate

The real interest rate is usually computed using the Fisher (1930) equation. Assuming that both the nominal interest rate and expected inflation rate are low, one can write:

$$r_t \cong i_t - \pi_t^e \tag{1}$$

where i_t is the nominal interest rate, r_t the *ex ante* real interest rate, and π_t^e the agent's expectation of inflation. The problem with this measurement of the real interest rate is that the expected inflation rate, π_t^e, is not s not directly observed. Therefore, in the literature, several methods have been proposed to measure the real *ex ante* interest rate (e.g. Mishkin, 1988). One of the widely used approaches is to use the observed inflation rate. Assuming that the expected *ex ante* inflation rate differs from the actual inflation rate observed *ex post* by a forecast error ε_t:

$$r_t^p = i_t - \pi_t \equiv r_t - (\pi_t - \pi_t^e) = r_t + \varepsilon_t \tag{2}$$

where r_t^p is the *ex post* real interest rate, π_t is the actual inflation rate. This approach of measuring *ex ante* real interest rate makes use of *ex post* real interest rate together with the assumption of rational expectations. If it is assumed that expectations are formed rationally, then $\pi_t^e = E[\pi_t | \Omega_t]$ and therefore, $E[\varepsilon_t | \Omega_t] = E[(\pi_t^e - \pi_t) | \Omega_t] = 0$, where $E[\bullet | \Omega_t]$ is the mathematical expectations operator conditional on all information available at time t, Ω_t. This result indicates that the inflation forecast errors ε_t are unforeseeable given any information available at time when expectations are formed.

The assumption of rational expectations implies that inflation forecast errors ε_t have zero mean and are uncorrelated over time. In addition, if one further assumes that the forecast error term ε_t is a stationary process, implying that the error term has a finite unconditional variance, it can be shown that the real interest rate is constant (e.g. Rose, 1988). The constant real interest rate implies that the nominal interest rate moves one-for-one with expected inflation rate. If changes in inflation rate are reflected one-by-one in nominal interest rates, then the inflation rate will have no permanent effect on real interest rates. This inference corresponds to long-run neutrality of money.

3. Methodology

Stationarity of real interest rates can be examined by applying conventional augmented Dickey-Fuller ([ADF] Dickey & Fuller, 1979) test. Let r_t denote the real interest rate. The ADF test is based on the following equation:

$$\Delta r_t = \alpha r_{t-1} + x_t^{'}\delta + \sum_{i=1}^{p} \beta_i \Delta r_{t-i} + \varepsilon_t, \tag{3}$$

where Δ is difference operator, $x_t^{'}$ is a vector of optional exogenous regressors, which may consist of a constant, or a constant and trend, α, β_i and δ are parameters to be estimated, and the ε_t are assumed to be white noise. Then the null hypothesis of unit root (H_0: $\alpha = 0$) against alternative of a stationary process (H_1: $\alpha < 0$) can be tested using the conventional t-ratio for α as $t_\alpha = \hat{\alpha}/s.e.(\hat{\alpha})$, where $\hat{\alpha}$ is the estimate of α and $s.e.(\hat{\alpha})$ is the coefficient standard error.

One of the drawbacks of the ADF test is that it has low power when adjustment to the equilibrium is nonlinear. As briefly discussed above, transaction costs may lead to nonlinear dynamics in the real interest rate. If the gain from adjustment is not sufficient large to cover transaction costs then economic agents will not adjust their portfolio so as to bring the real rate of return to the equilibrium level. If the gain from adjustment is sufficiently large, however, then agents shall act so as to move the system back towards the equilibrium. This suggests that if the deviation from the equilibrium level is small, then the real interest rate may not revert to the equilibrium. However, if the deviation from the equilibrium is large, then real rate will revert to the equilibrium level, implying that the adjustment of the real interest rate to the desired level might be inherently non-linear.

Kapetanios et al. (2003) argued that conventional unit root tests may lack power if the dynamics of the series are nonlinear, and developed a new procedure that allow for nonlinear adjustment. The unit root test procedure proposed by Kapetanios et al. (2003) is based on the following exponential smooth transition (ESTAR) regression model:

$$\Delta r_t = \varphi r_{t-1} + \gamma r_{t-1}[1 - \exp(-\theta r_{t-1}^2)] + \varepsilon_t, \tag{4}$$

where r_t is the series under investigation. The transition function $F(\theta, r_{t-1}) = 1 - \exp(-\theta r_{t-1}^2)$, is continuous, U-shaped around zero and bounded from zero and one. The parameter θ measures the speed of transition between two regimes that correspond to extreme values of the transition function.

The ESTAR model has a nice property in that it allows modelling of different dynamics of series depending on the size of the deviations from the equilibrium level. As discussed briefly above, small perturbations in the real interest rate may have no tendency to revert to the equilibrium levels. If the deviations from the equilibrium are large enough, however, then the real interest rate shall revert to the equilibrium level. In the context of ESTAR model, this would imply that while $\phi \geq 0$ is possible, one must have $\gamma < 0$ and $\phi + \gamma < 0$ for the process to be globally stationary. Under these conditions, the process might display unit root for small values of r_{t-1}^2, but for larger values of r_{t-1}^2 it has stable dynamics, and as a result, is geometrically ergodic. As shown by Kapetanios et al. (2003), the ADF test may not be very powerful when the true process is nonlinear yet globally stationary.

Imposing $\phi = 0$ (which implies that r_t follows a unit root in the middle regime) the ESTAR model can be written as:

$$\Delta r_t = \gamma r_{t-1}\left[1 - \exp\left(-\theta r_{t-1}^2\right)\right] + \varepsilon_t \tag{5}$$

The null hypothesis for the global stationarity of the process r_t can be written as H_0: $\theta = 0$ against the alternative H_1: $\theta > 0$. However, testing the null hypothesis directly is not feasible since the parameter γ is not identified under the null. To circumvent this problem, Kapetanios et al. (2003) develop a t-type test statistic by replacing the transition function $F(\theta, r_{t-1}) = 1 - \exp\left(-\theta r_{t-1}^2\right)$ with its first-order Taylor approximation around $\theta = 0$, yielding the following regression model:

$$\Delta y_t = \delta r_{t-1}^3 + e_t \tag{6}$$

where e_t contains not only original error term ε_t but also the error term resulting from Taylor approximation. The null hypothesis of unit root can be tested using t-statistics from that testing $\delta = 0$.

In the more general case where errors in (5) are serially correlated, assuming that the serially correlated terms enter in a linear fashion, one may extend the auxiliary regression (5) to

$$\Delta r_t = \delta r_{t-1}^3 + \sum_{i=1}^{p} \beta_i \Delta r_{t-i} + e_t \tag{7}$$

Although the test procedure of Kapetanios et al. (2003) is convenient for testing the stationarity in the case of nonlinear adjustment, this test procedure does not take account of possible structural breaks in the data generating process. Transition countries have undergone major structural changes during the transition period (e.g. Fischer & Sahay, 2000; Fischer, Sahay, & Vegh, 1996; Foster & Stehrer, 2007). The transition countries had to implement a wide range of economic reforms, aiming at price and trade liberalisation, privatisation, enterprise restructuring, and establishment of market-based financial systems in order to restructure their centrally planned economies to market economies. Such changes may have caused equilibrium real interest rates in these transition countries to shift during the analysed period. In addition, these countries have integrated into the world financial markets. Such changes in the transition countries may have caused equilibrium real interest rates in these countries to shift during the analysed period.

The failure to take account of a possible structural break in data generating process may give rise to spurious results. In his seminal contribution, Perron (1989) argued that the 1973 oil shock was followed by a change in the slope of the trend for most aggregate economic variables. He has shown that power of unit root tests may decrease if there is a structural change, even if the data are indeed stationary. In order to account for a structural break, Perron (1989) suggested adding dummy variables corresponding to a pre-specified break date to the standard ADF regression. Subsequently, Zivot and Andrews (1992) proposed new procedure that selects the break date endogenously. Leybourne, Newbold, and Vougas (1998) argued that the assumption of instantaneous structural change may not be appropriate for economic time series, and suggested a new test procedure that allow for gradual structural change. Recently, Sollis (2004) extended the test procedure of Leybourne et al. (1998) to allow for both gradual structural change and asymmetric adjustment.

Sollis (2004) combines the smooth transition methodology of Leybourne et al. (1998) with the threshold autoregressive methodology of Enders and Granger (1998), and develops unit root tests that allow for a smooth transition between deterministic linear trends, around which asymmetric adjustment may occur. Following Leybourne et al. (1998), Sollis (2004) models the structural change in the series by the following smooth transition regression model:

$$r_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + v_t \gamma > 0 \tag{8}$$

where v_t is a zero-mean I(0) process and $S_t(\gamma, \tau)$ is the logistic smooth transition function, based on a sample size of T, $S_t(\gamma, \tau) = [1 + \exp\{-\gamma(t - \tau T)\}]^{-1}$. The transition function $S_t(\gamma, \tau)$ is a continuous function bounded between zero and one. The parameter τ determines the timing of the transition midpoint, and the parameter γ determines the speed of transition between regimes. If it is assumed that v_t is a zero-mean I(0) process, then regression (8) implies that the series under investigation r_t is stationary around a

nonlinear 'attractor', whose mean changes from initial value α_1 to $\alpha_1 + \alpha_2$ and slope changes from initial value β_1 to $\beta_1 + \beta_2$ simultaneously and with the same speed of adjustment (Leybourne et al., 1998).

Whether the series under investigation r_t converge to nonlinear attractor or not can be tested via a two-step procedure. In the first step, Equation (8) is estimated using a nonlinear least squares algorithm, and in the second step the following asymmetric ADF regression model is estimated for the residuals $\hat{v}_t$ obtained from the first step:

$$\Delta\hat{v}_t = I_t\alpha_1\hat{v}_{t-1} + (1 - I_t)\alpha_2\hat{v}_{t-1} + \sum_{i=1}^{p} \beta_i\Delta\hat{v}_{t-1} + \eta_t \qquad (9)$$

where $I_t = 1$ if $\hat{v}_{t-1} \geq 0$, $I_t = 0$ if $\hat{v}_{t-1} < 0$, and η_t is a zero-mean stationary process. If $\alpha_1 = \alpha_2 = 0$ in equation (t), then $\hat{v}_t$ and therefore y_t contains a unit root, while if $\alpha_1 = \alpha_2 < 0$, y_t is a stationary smooth-transition threshold autoregressive (ST-TAR) process with symmetric adjustment, and if $\alpha_1 < 0$, $\alpha_2 < 0$ and $\alpha_1 \neq \alpha_2$, then r_t is a stationary ST-TAR process with asymmetric adjustment. As shown by Sollis (2004), one may test the null hypothesis of unit root using the most significant of the t-statistics from those testing $\alpha_1 = 0$ and $\alpha_2 = 0$ or by using the F-statistic for testing $\alpha_1 = \alpha_2 = 0$ in (9).

4.   Data and test results

In this study, we consider the real interest rate series of 21 transition countries, whose monthly interest rate and inflation data were available in the International Monetary Fund's International Financial Statistics (IFS) database. Inflation was calculated as a percentage change of consumer price index expressed on a per annum basis. We use lending rates for all countries as the nominal interest rate. This was the only interest rate available for most countries over a longer period of time. The sample period is 1999:01 to 2014:06 except for five countries. The sample ending date for Croatia is 2014:03, for Latvia is 2014:01, for Lithuania is 2010:10, for Poland 2006:12 and for Slovenia 2009:10 because of data limitations.

The real interest rate data was generated using the following equation:

$$1 + r_{t+1} = \frac{1 + i_t}{1 + \pi_{t+1}} \qquad (10)$$

where r is the real interest rate, i is the nominal interest rate and π is the inflation rate[2].

4.1.   Unit root test results

We first test the stationarity of the real interest rate series disregarding any possible structural changes and non-linearities in the series. The results of the ADF test are provided in (Table 1). When no trend is included in the ADF regression, the null hypothesis of unit root is rejected in fifteen cases out of 21, namely for the Armenian, Belarusian, Bulgarian, Croatian, Georgian, Czech, Estonian, Hungarian, Latvian, Kazakh, Kyrgyz, Moldovan, Romanian, Slovenian, and Ukrainian real interest rate series. With a linear trend, the null hypothesis of unit root is rejected for 12 series, namely for Bulgarian, Croatian, Georgian, Estonian, Hungarian, Latvian, Kazakh, Kyrgyz, Moldovan, Romanian, Russian, Slovenian, and Ukrainian real interest rate series.

Before applying nonlinear unit root tests of Kapetanios et al. (2003) and Sollis (2004), we first tested for nonlinearity of the real interest rate series under investigation.

Table 1. Results of the ADF test.

Country	Constant	Constant and trend
Albania	–2.462	–2.623
Armenia	–2.817***	–2.435
Azerbaijan	–2.312	–1.888
Belarus	–2.999**	–2.804
Bulgaria	–3.303**	–3.991**
Croatia	–3.733*	–3.788**
Georgia	–3.303**	–3.991**
Czech Republic	–2.641***	–2.667
Estonia	–3.799*	–3.696**
Hungary	–3.236**	–3.209***
Latvia	–3.520*	–3.543**
Lithuania	–1.504	–1.632
Kazakhstan	–4.080*	–4.185*
Kyrgyz Republic	–3.424**	–3.628**
FYR of Macedonia	–1.207	–2.608
Moldova	–3.768*	–3.632**
Poland	–0.892	–3.097
Romania	–4.946*	–4.940*
Russia	–2.549	–4.142*
Slovenia	–2.856***	–3.870**
Ukraine	–3.225**	–2.999

Notes: *, **, and *** Denote rejection of the null hypothesis of unit root at 1, 5 and 10% significance levels.
Source: Authors' calculation.

For this purpose, we applied *F* versions of the LM-type linearity tests suggested by Terasvirta (1994). The results of the linearity tests are presented in (Table 2)[3]. As can be readily seen from Table 2, we find strong evidence of nonlinearity in real interest series of all countries except for the Azerbaijan, Czech Republic, Estonia, Hungary, and Kyrgyz Republic.

Next we consider the results of nonlinear unit root test procedure of Kapetanios et al. (2003) (the KSS test), which are presented in (Table 3). When we use de-meaned series, the null hypothesis of unit root is rejected in the case of 10 countries, namely in the cases of Armenia, Belarus, Croatia, Georgia, Latvia, Lithuania, Kazakhstan, Moldova, Romania, and Ukraine. When we further de-trend the data, we find the real interest rate series of eight countries, namely of Belarus, Georgia, Latvia, Kazakhstan, Moldova, Poland, Romania, and Ukraine are stationary. These results suggest that real interest rate series of these countries might be inherently nonlinear, implying that shocks to the real interest rate in these countries affect economic variables in a nonlinear fashion.

Now, we turn to the results of the Sollis (2004) smooth transition-threshold autoregressive (ST-TAR) unit root procedure. The results of the ST-TAR unit root test that allows for both gradual structural change and asymmetric adjustment are provided in (Table 4) below.

As can be seen from the table, the null hypothesis of unit root is rejected in the case of 10 countries, namely of Albania, Armenia, Bulgaria, Croatia, Georgia, Latvia, Kyrgyz Republic, Macedonia, Romania, and Russia once we allow for break in the mean of the series. When we further allow for break in the trend of the data as well, the real interest rate of eight countries, namely, for Albania, Armenia, Bulgaria, Georgia, Latvia, Romania, Russia, and Slovenia are found to be stationary.

Table 2.   Linearity test results.

	d	F statistics
Albania	5	3.415 (0.000)
Armenia	1	2.726 (0.000)
Azerbaijan	9	1.225 (0.209)
Belarus	12	3.413 (0.000)
Bulgaria	3	1.778 (0.017)
Croatia	4	2.541 (0.000)
Georgia	8	2.307 (0.001)
Czech Republic	13	1.002 (0.480)
Estonia	6	1.136 (0.297)
Hungary	2	1.789 (0.104)
Latvia	1	6.250 (0.000)
Lithuania	1	2.546 (0.022)
Kazakhstan	1	8.378 (0.000)
Kyrgyz Republic	1	0.596 (0.734)
FYR of Macedonia	1	2.717 (0.000)
Moldova	5	2.778 (0.000)
Poland	8	1.954 (0.025)
Romania	8	3.017 (0.000)
Russia	13	3.351 (0.000)
Slovenia	6	2.431 (0.002)
Ukraine	1	2.950 (0.001)

Notes: d is the lag order that minimises the probability associated with the linearity test. Probabilities of the F statistics are given in parenthesis (Terasvirta, 1994).
Source: Authors' calculation.

Table 3.   Results of the nonlinear KSS unit root test.

Country	De-meaned series	De-meaned and de-trended series
Albania	−1.026	−0.468
Armenia	−3.248**	−2.901
Azerbaijan	−1.446	−2.333
Belarus	−2.674***	−3.202***
Bulgaria	−2.598	−2.102
Croatia	−2.810***	−2.785
Georgia	−3.276**	−5.452*
Czech Republic	−2.016	−2.193
Estonia	−0.906	−0.850
Hungary	−2.652	−2.602
Latvia	−8.236*	−7.996*
Lithuania	−2.841***	−2.269
Kazakhstan	−3.600*	−3.450**
Kyrgyz Rep.	−2.653	−2.860
R.of Mac.	−1.264	−2.027
Moldova	−4.115*	−4.161*
Poland	−1.764	−3.512**
Romania	−5.131*	−3.634**
Russia	−2.346	−2.491
Slovenia	−0.789	−0.492
Ukraine	−4.705*	−3.714**

Notes: *, **, and *** Denote rejection of the null hypothesis at 1, 5 and 10% significance levels. Critical values of the test statistic at 1, 5 and 10% significance levels are −3.48, −2.93 and −2.66 for the de-meaned data and −3.93, −3.40, −3.13 for the de-trended data, respectively (Kapetanios et al., 2003, 364).
Source: Authors' calculation.

Table 4. Results of the ST-TAR unit root test.

Country	Break in the mean, no trend		Break both in mean and trend	
	t_{max}	F stat	t_{max}	F stat
Albania	–4.572*	11.643**	–3.753***	9.626
Armenia	–4.434*	11.640**	–4.395*	11.485***
Azerbaijan	–2.782	5.109	–3.242	7.279
Belarus	–2.314	3.318	–2.956	4.420
Bulgaria	–3.777**	11.142**	–3.890***	10.326
Croatia	–3.258***	8.315***	–3.363	9.067
Georgia	–4.250*	11.307**	–4.379**	12.178***
Czech Republic	–2.610	3.549	–2.558	3.372
Estonia	–3.662	7.204	–3.547	6.636
Hungary	–2.589	5.259	–2.580	5.236
Latvia	–3.638**	7.277	–3.768***	7.784
Lithuania	–1.601	1.416	–3.119	8.743
Kazakhstan	–3.082	5.133	–2.877	6.802
Kyrgyz Republic	–3.250***	8.222***	–2.670	6.613
FYR of Macedonia	–3.636**	6.621	–3.63568	6.62117
Moldova	–2.345	4.348	–3.363	7.406
Poland	–3.121	5.781	–3.058	5.341
Romania	–3.826**	12.428*	–4.238**	13.968**
Russia	–3.453**	6.586	–5.306*	14.257**
Slovenia	–3.010	4.551	–4.230**	9.356
Ukraine	–3.138	6.036	–3.239	6.798

Notes: *, **, and *** Denote rejection of the null hypothesis at 1%, 5% and 10% significance levels. For de-meaned series critical values of the t-max and F-statistics at 1%, 5% and 10% significance levels are –3.994, –3.417, –3.169 and 12.244, 9.191 and 7.844 for a sample size of 100 and –3.890, –3.385, –3.140 and 11.786, 9.029 and 7.759 for a sample size of 200. For de-meaned and de-trended series these significance levels are –4.586, –4.049, –3.803 and 16.834, 13.408, and 11.862 for a sample size of 100 and –4.393, –3.937, –3.704, and 15.892, 12.787 and 11.349, respectively for a sample size of 200 (Sollis, 2004, 413).
Source: Authors' calculation.

4.2. Discussion of the results

Before proceeding to discussion of the stationarity test results, we must remember some technical features regarding stationarity tests employed in this article. First we note that none of the test procedures used in this article has absolute power over the other procedures in all cases. In fact, the KSS test has relatively good size and power over the ADF test only if the true data generating process is nonlinear. Similarly, the ST-TAR has good size and power if the true data generating process exhibits relatively large and smooth breaks as well as asymmetric adjustment. It is natural that introduction of additional nonlinearities and structural changes will distort the power of the unit root tests if the true data generating process is almost linear or if the size of the change in the mean or slope of the trend is relatively small (e.g. Kapetanios et al., 2003; Leybourne et al., 1998; Sollis, 2004). Consequently, each of the tests have relatively good size and power properties for different data generating processes, and their results for individual series must be compared cautiously (see also Hasanov & Telatar, 2011; Öge Güney & Hasanov, 2014).

Second, the theory suggests that real interest rates revert to a constant mean, but not to a trend. However, it is a well-known fact that transition countries have experienced rapid productivity gains due to economic restructuring during the transition period (e.g. Halpern & Wyplosz, 1997; Kutan & Yigit, 2007). In addition, these countries have

integrated into the world financial markets and attracted huge capital inflows. Such changes may have caused to a trend fall in real interest rate series. Furthermore, domestic savings have also grown with rapid economic growth in these countries, which ultimately reduces interest rates. Taking these specific features of transition countries, therefore, we also add a trend to the test equations. In passing we note that many studies on the validity of the purchasing power parity (PPP) hypothesis for the transition countries found evidence in favour of PPP only after allowing for a time trend although 'pure' PPP theory suggests that real exchange rates revert to mean, but not to a trend (see for example, Telatar & Hasanov, 2009). Now, we turn to discussion of the results.

The ADF test suggested that the real interest rate series of 15 out of total 21 countries, namely, Armenia, Belarus, Bulgaria, Croatia, Georgia, Czech, Estonia, Hungary, Latvia, Kazakhstan, Kyrgyz Republic, Moldova, Romania, Slovenia, and Ukraine are reverting to constant mean. This finding provides evidence in favour of the 'pure' Fisher hypothesis for these countries, and suggests that shocks to nominal interest rates (e.g. monetary policy) will not affect real interest rates, and hence, real economic activity. The ADF test also suggested that the Russian real interest rates are trend stationary. Although Russia has experienced a severest financial and economic crisis in 1998, it recovered from the crisis very quickly thanks to increasing energy and commodity prices. Therefore, the trend stationarity of Russian real interest rates may be attributed to growing inflow of oil revenues. On the other hand, the real interest rate series of Albania, Azerbaijan, Lithuania, FYR Macedonia and Poland are not found to be stationary according to the ADF test.

The ADF test assumes that the adjustment towards equilibrium level is linear, i.e., adjustment does not depend on the sign or size of disequilibrium. However, as we briefly discussed in introduction, official interventions or transactions costs may cause to nonlinearities in the adjustment towards equilibrium. After allowing for possible nonlinearities in the adjustment, we found that the real exchange rates of Lithuania and Poland are also stationary. The fact that the Polish real interest rate is reverting to a trend but not to a constant mean may be result of productivity gains in the country during the accession period. Furthermore, this result suggests that only large deviations from the equilibrium were corrected in these two countries. Note also that the nonlinear KSS unit root test rejected the null hypothesis of unit root only for those countries for which we found nonlinearities. This result once more demonstrates that the KSS test has a good power if the series exhibit nonlinear dynamics.

After allowing for structural breaks in addition to possible nonlinearities in adjustment to equilibrium, we found that the real interest rate series of Albania and FYR Macedonia are also mean reverting whereas mean has changed gradually over time. This suggests that the real interest rate series, and hence, the equilibrium relationship between real economic variables of these two countries might have shifted during the analysed period. These two Balkan countries are less integrated to the European financial markets, and lagged behind other CEE countries in reforming their economies and financial institutions. Other CEE transition countries reformed their economies and integrated into the European Union during the accession period. Therefore, it is interesting to find that only the unit root test procedure that allows for structural breaks provided evidence of stationarity in the case of these two countries.

The only country for which we found no evidence of stationarity of real interest rates is Azerbaijan. In fact, none of the tests considered in this article provided evidence in favour of stationarity of real interest rates in Azerbaijan. In their analysis of RIRP hypothesis, Öge Güney and Hasanov (2014) also found that the RIRP holds for all

transition countries under consideration except for Azerbaijan[4]. The non-stationarity of real interest rates in Azerbaijan can be attributed to specific features of the Azerbaijani economy. Due to huge oil revenues, Azerbaijan is not dependent on external financial sources. Indeed, Azerbaijan in net lender in international financial markets and the ratio of external debt to GDP is the lowest amongst transition countries. Furthermore, share of foreign-owned banks is lower in Azerbaijan when compared to other transition countries. These facts, coupled with low level of financial development explain why Azerbaijani real interest rates are non-stationary.

All in all, we find that the real interest rate series of all transition countries but Azerbaijan are stationary. This finding suggests that there is a room for stabilisation policies only in Azerbaijan whereas monetary policy in other transition countries will have no *persistent* effects on the real economic variables in the *long run*. However, as we were able to reject the null hypothesis of unit root only after allowing for structural breaks and nonlinearities in the adjustment to equilibrium, there might be a limited room for monetary stabilisation policies for some of the transition countries as well. For example, real interest rates of Albania and FYR were found to be stationary *only* after allowing for *both* structural breaks and nonlinearities in the adjustment towards equilibrium. This finding points to the validity of the so-called 'Lucas Critique', which criticises reliance on policy conclusions based entirely on the relationships observed in the past as expectations of agents varies with economic policy and changes in the economic environment. Note also that we rejected the null of unit root in the real interest rate of Poland only after allowing for nonlinear adjustment. Therefore, changes in monetary policy will affect real economic variables in Albania, FYR Macedonia and Poland in a nonlinear fashion as are result suggest that adjustment to equilibrium might be nonlinear in these three countries. For all the remaining transition countries, monetary policy will have no persistent effects on real economic variables in the long run as our results suggest that adjustment is not dependent on the sign and size of deviations from the equilibrium. This finding, in practice means that real interest rates will correct both positive and negative small and large deviations leaving no room for monetary policy to have a persistent effect on real economic variables.

5. Conclusion

The real interest rate is one of the most important variables in macro economics and statistical characterisation of the real interest rate has important policy and theoretical implications. For example, if real interest rates contain a unit root, then monetary policy will have permanent effects on real economic variables. Although a considerable amount of literature has been devoted to testing the stationarity of the real interest rate for developed countries, only limited work is done for transition countries. Therefore, in order to fill a gap in the empirical literature, in this article we investigate the stochastic properties of real interest rates in the case of for transition countries. Taking account of the fact that these countries have undergone major reforms during the transition period, we apply newly developed unit root tests that allow for nonlinearity and structural breaks.

We find that most of the series under consideration are stationary, in accordance with the Fisher hypothesis. In addition, the test results suggest that real interest rates of majority of the transition countries exhibit some form of nonlinearity. This finding implies that monetary policy in these countries affects real economic variables in a nonlinear fashion. Our results point to importance of taking account of both nonlinearities and structural breaks when analysing interest rate series of transition countries.

Disclosure statement

No potential conflict of interest was reported by the authors.

Notes

1. Another related issue is the so-called 'real interest rate parity' (RIRP) hypothesis, which implies that the expected real interest rates between two countries are equal. If RIRP holds, then individual countries will be unable to affect their real interest rate, which will be set internationally. Cuestas and Harrison (2010) have tested the RIRP across 13 Central and Eastern European (CEE) countries, and find a stronger evidence of RIRP after allowing for asymmetries in the speed of mean reversion.
2. We considered both *ex ante* and *ex post* real interest rates in all test regressions. *Ex ante* inflation rates were estimated by the method of moving averages. The results of the tests using both *ex ante* and *ex post* real interest rates were quite similar to each other. Therefore, following earlier research, in this study we report only the results using *ex post* real interest rates to save space. The results with *ex ante* rates are available from authors upon request.
3. The details of linearity tests and estimates of auxiliary regressions are available upon request from the authors.
4. The RRIP rests on uncovered interest parity (UIP) and PPP hypotheses. According to RIRP, the real interest rate differentials between two countries follow a white-noise process, implying that domestic real interest rates cannot deviate from international interest rates permanently

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