THE STATIONARITY OF CONSUMPTION-INCOME RATIOS WITH NONLINEAR AND ASYMMETRIC UNIT ROOT TESTS: EVIDENCE FROM FOURTEEN TRANSITION ECONOMIES

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Abstract

This paper analyses the stationarity properties of the consumption-income ratios for a sample of 14 transition economies by taking account of nonlinearities and asymmetries together using the unit root tests based on the TAR models. The results provide evidence in favour of the stationary consumption-income ratios for all countries.

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I. Introduction

In recent years a number of studies have paid considerable attention in empirical literature to testing for the existence of a unit root in the consumption-income ratio (also known as the average propensity to consume (APC)). Investigating the stochastic properties of consumption-income ratio is inevitably crucial issue, because it provides important evidence about the validities of the basic assumptions relating the APC ratio of the consumption theories, so does the validity of the theory itself. A stationary APC means that APC converges towards a constant in the long run. Findings of stationary APC provide some support to the relative income hypothesis, the habit persistence model, the permanent income hypothesis and the life cycle hypothesis. On the other hand, an integrated behaviour means that policy shocks are likely to have a permanent effect on the APC and may be implied by the Keynesian absolute income Hypothesis, the involuntary savings theory and the Marxian undercompensating theory.

The objective of the study is to examine whether the APCs are stationary for 14 transition economies which includes 8 countries from Central and Eastern European Countries (CEE) (Bulgaria, Croatia, the Czech Republic, Hungary, Poland, Romania, Slovakia, and Slovenia), the 3 Baltic states (Estonia, Latvia, and Lithuania), and the 3 Commonwealth of Independent States (CIS) (Belarus, Kazakhstan, and Russia) for over a period following their transition stages which is defined as 1989 for Central Europe and 1990-91 for the CIS and the Baltic states. Since the macroeconomic impact of the transition from planning to markets has been played a role like a dramatic shock in the rates of growth and inflation, researching the stationarity of APC for these economics can provide a strong evidence whether a shock is persistence or not. If the shock is not transitory, APC should exhibit a unit
root process after the transition period; otherwise households in these countries try to disperse the effect of transition on their consumption period to time by keeping the APC almost stable.

Our study differs from the previous studies examining the stationarity of APC from two perspectives. Firstly, it considers possible mis-classification of the integrated nature of the consumption-income ratios due to a failure to allow for nonlinear symmetries and asymmetries together. To our knowledge, this is the first attempt to investigate the effect of nonlinearities and asymmetries on the stochastic behaviour of the APC in a univariate context. To investigate the stochastic properties of APC, in addition to the conventional linear ADF (Dickey and Fuller, 1979), Phillips and Perron (1988; PP) and Kwiatkowski et al. (1992; KPSS) tests, we employ the Kapetanios et al. (2003) ESTAR unit root test which allows for nonlinear (and linear) adjustment towards to equilibrium and also the asymmetric ESTAR nonlinearity test developed by Sollis (2009). The second difference of our study is its sample countries. We are not aware of any previous study that investigates the mean-reverting nature of APC for these fourteen transition countries together. Since the transition can be thought as a shock which causes to force the APC away from its equilibrium levels for the transition countries, the converges of APC to its long-run equilibrium would have a strong supportive evidence for the stationarity of APC. Further, the outcomes of the unit root tests applied here will provide an idea whether adjustments of consumption to changes is GDP happened as nonlinearly or linearly and asymmetrically or symmetrically.

Our results suggest that for all countries the nonstationarity of the APC could not be rejected, but the evidence in favour of the existence of a nonlinear and/or an asymmetric adjustment process is mixed. While for 6 countries, Belarus, Croatia,
Lithuania, Poland, Romania, and Slovakia the APC is linearly stationary, the APC for
the other 8 countries exhibit a nonlinear long-run path. Further for 5 economies,
Bulgaria, Czech Republic, Estonia, Latvia, and Russia, the APC is globally
asymmetric stationary.

The paper is organized as follows. In the next section, we provide the
theoretical insights and empirical findings on the stationarity of the APC. Then we
discuss the unit root testing methods. The data and the test results are given in the
following section. The final section contains a general discussion of our findings and
concludes the paper.

II. Theoretical Insights and Empirical Findings on the Stationarity of the APC

2.1. Theoretical Insights

The consumption function provides an excellent illustration of a typical
sequence in the development of knowledge in economics (Branson, 1989). The first
step of this is the modern consumption theory which begins with "fundamental
psychological law" of consumption proposed by Keynes (1936: 96) in his General
Theory: ¹

“The fundamental psychological law ,..., is that men are
disposed, as a rule and on the average, to increase their
consumption, as their income increases, but not by as much as
the increase in their income.”

Based on this ‘law’, according to the Keynesian consumption function, the
Absolute Income Hypothesis (AIH, hereafter), aggregate consumption is a stable, not
necessarily linear, function of disposable income,
where $C_t$ and $Y_t$ denote the (real values of) total personal consumption expenditure and total disposable income, respectively at time $t$. $\beta$, the marginal propensity to consume (MPC) is expected to be constant and close to one, and $\alpha$, the autonomous component of consumption, is assumed to be small but positive. By capturing the conjectures of the fundamental law, the AIH has two important features:

1. The average propensity to consume (APC) is greater than the MPC: $APC > MPC$, where $APC=C/Y$ and
2. As income rises, the proportion of it consumed falls: $\frac{dAPC}{dY} < 0$ so the income elasticity of consumption defined as $MPC/APC$ would be less than unity.

The second step in the sequence was providing econometric information about consumer behaviour. This early econometric history of the consumption function gave efforts to test the relationship between consumption and income proposed by the AIH with whatever kind of data was available, using whatever specification seemed reasonable (Bunting, 2001). However, aside from these studies, especially the seminal study made by Kuznets (1946) - a Nobel prize winner - was a turning point in the development of the consumption function literature, because the other studies were either cross-sectional or short-run time-series in detail whereas Kuznets’s study was of a long-run nature. Kuznets showed that except for the Depression years, the APC in the US over the period 1869–1938 fluctuated narrowly between 0.84 and 0.89. In other words, APC was approximately mean-reverted, even income increased a lot, consumption kept almost a stable fraction of income, so consumption was a proportion rather than a function of income. In addition, Kuznets concluded that the intercept of the consumption function seemed
to "ratchet" up over time; the income elasticity of consumption was less than unity for only short-run, but in the long run, it was unity. These empirical ‘anomalies’ -as Friedman (1957) termed- turned up a ‘consumption puzzle’, an interesting and seemingly contradictory fact with the assumptions made by the AIH.⁴

The third step in the sequence of research into the consumption function was to the attempt at resolving the consumption puzzle by replacing a forward-looking consumer unit instead of a Keynesian consumer who saw the current income as the only determinant of the current consumption. These rigorous and elaborate theories were the relative income hypothesis (RIH) of Duesenberry (1949), the permanent income hypothesis (PIH) of Friedman (1957), and the life cycle hypothesis (LCH) developed in a series of papers Ando, Brumberg, and Modigliani beginning in the early 1950s. The RIH has two claims. The first claim is that: The decisions on consumption are determined by ‘relative’ consumption patterns or the demonstration effect also known as the ‘keeping up with the Joneses’. Duesenberry (1948: 74) assumed the strength of any individual’s desire to increase his consumption expenditure as a function of the ratio of his expenditure to some weighted average of the expenditures of others with whom he wishes to keep up. The second claim is based a habit persistence nature of consumption and in determining current consumption the role of past peak levels of income have a particular importance. Duesenberry (1948: 70) is based his argument on a fundamental psychological postulate that

“...it is harder for a family to reduce its expenditure from a higher level than for a family to refrain from making high expenditures in the first place.”
The habit persistence nature of the RIH implies that consumption grows at a constant rate by causing a constant long-run APC. On the other hand, based in intertemporal consumption theory both the LCH and the PIH provide an explanation for consumption smoothing over time by replacing an expected lifetime or a permanent income as a long-run issues instead the observed current income. Further, in these models, any unexpected change in current income will not affect current consumption, as long as this change is thought as transitory. In addition, as average income increased over time, the intercepts of the short-run consumption functions will have a persistent upward shifting or ratcheting found in the cross sectional functions by keeping the APC as a constant and equals to the MPC, so income elasticity of consumption is equal to unity.

Assuming the APC is proportional to changes in income or not causes different policy implications. With a declining APC, the GDP growth will be more volatile, since the more variable components in aggregate demand such as investment and exports become dominant in cyclical fluctuations. In other words, consumption spending cannot serve as a built-in stabilizer for the economy unless the government takes measures to raise the share of such expenditures in output (Abeysinghe and Choy, 2004). As Hadjimatheou (1987) notes that the meaning of a country’s APC should be constant through time is that a nation’s saving rate is independent of its level of per-capita income and positively related to its long-run growth rate. Hence, unless it’s long-run income growth rate changes, the APC will be constant. Thus, the decreasing APC, as assumed in the AIH, challenge both the need and effect of interventionist fiscal policies. To stimulate the economy by increasing private consumption is definitely possible by cutting taxes which give a rise to the disposable income. There is also no a specific role for the interest rate,
money or the exchange rate. Thus, changes in the instruments we normally think a central bank controls cannot, on their own, affect consumption expenditure directly. Of course, if the instruments available to central banks are able to affect disposable income, then according to the AIH, they will also affect consumption indirectly (and probably with a lag thereby introducing lags of monetary policy on the economy). On the other hand, if consumers try to smooth their consumption over time as suggested by the RIH, the PIH, and the LCH, aggregate demand management should have little effect on the short-run spending; that is discretionary demand management is less effective (Brady, 2008). According to Friedman (1957: 238),

"current consumption is adapted to some measure of longer-run income status, ... a much larger part of current income is interpreted as autonomous and a much smaller part as dependent on current income ... . The result is a smaller investment multiplier, and an inherently cyclically more stable system."

Also, Modigliani and Brumberg (1954, p.431) believed in that

"our new understanding of the determinants of saving behaviour cast some doubts on the effectiveness of a policy of income redistribution for the purpose of (changing) the average propensity to save."

So for the PIH and the LCH, monetary policies play a greater role than fiscal policies to determine the aggregate consumption as opposite to the AIH. The main reason of this difference is that the PIH and the LCH treat consumption as determined by permanent income or wealth which is greatly affected by the changes in the monetary policies, because of the need of portfolio adjustments which occur across the whole range of financial assets and (durable) goods. Moreover, in the PIH
and the LCH, for the consumers who only care about the present and the future, so their expectations play a crucial role, the source of the shocks can be important: Consumers will react differently whether a shock is permanent or transitory. Consumption is especially responsible to permanent changes rather than transitory shocks. Thus, for consumers, in determining the aggregate consumption a transitory shock has an ignorable impact. A transitory change in taxes will no effect on consumption, so fiscal policy will be ineffective. As a result, the AIH’s assumption of a decreasing APC implies that consumers only care about the short run, so does the current income and in determining the consumption, fiscal policies can be adapted. On the other hand, the RIH, the LCH and the PIH assume that the aggregate APC will be constant over time, because the consumers are forward-looking and try to smooth their consumption so permanent changes made by monetary policies could be more effective to stimulate the aggregate consumption.

2.2. Empirical Findings on the Stationarity of the APC

There are a number of studies investigating the constancy of the APC by taking into the new developments in time series. Since whether the APC is mean reverting or not will affect empirical modelling of consumption functions and the implementation of economic policies as well as our understanding of savings behaviour and business cycles, the examination of the integrated nature of the APC has a long history in theoretical and applied economics. The starting point for these studies can be described as the seminal work of Nelson and Plosser (1982) which examines the presence of stochastic trends in many macroeconomic time series. The nonstationarity of APC contradicts with the belief of unit income elasticity of
consumption, so does the constancy of it. If the APC has a unit root, it can therefore be concluded that the APC is path dependent as its current value heavily depends on past levels. In this case, temporary shocks affect the variable permanently as the effect accumulates over time. So, each stochastic shock will have a permanent effect on the series, so that for a pure random walk, all fluctuations will represent permanent changes in the trend. Hence, shocks to a random walk will persist forever. In contrast, a stationary APC suggests the presence of a long-run equilibrium relationship between consumption and income, and hence APC converges towards a constant in the long run. For a stationary APC, shocks seem to be temporary, as suggested by the PIH and the LCH allowing consumers a desire to smooth their lifetime consumption paths.

The conflict among the results has generally been evaluated by depending on unit root tests applied. To have a clear-cut idea what the previous studies have reached and where the literature goes along the econometric techniques chosen, the authors of this study believe in that grouping the empirical investigations about the stationarity of the consumption-income ratios by depending upon their preferred econometric methodologies will be fruitful. The first group of empirical research consists of the studies which are based on the univariate Augmented Dickey-Fuller (ADF) type unit root and/or cointegration tests. The second group of investigations consists of the studies used the panel unit root test. The common belief of these studies is to increase the power of univariate unit root tests employed for the studies in the first group. The third group of studies takes into account for structural changes. These researches think that without giving a careful attention to describe the true data generating process, all types of unit root tests, univariate or panel, can cause a rejection of the stationarity of the APC, because if the APC includes a structural
break, the unit root tests will have low power. The last group includes the studies give
a special interest to unit root tests which are based on nonlinearities and/or
asymmetries. According to these researchers, taking into account for nonlinearities
and/or asymmetries, increase the power of univariate unit root tests.

One of the first groups of studies, Molana (1991) found a nonstationary APC
for the UK over period 1966Q4-1981Q4 with the ADF unit root test. Hall and
Patterson (1992) also analysed the UK case over period 1972Q2-1988Q3 using the
ADF unit root test and Johansen maximum likelihood (ML) extension of cointegration
and resulted with a nonstationary APC. However, Blinder et al (1985) reached a
stationary APC for the United States over period 1954Q1-1984Q4 by estimating a
consumption function with the OLS by taking into account anticipated and
unanticipated wealth and incomes separately. In addition, Campbell (1987) also
examined the US’s APC with the ADF and the Engle and Granger (EG) cointegration
test. They concluded that APC for the US is stationary when the ADF test is applied
to the residuals from the EG cointegration equation for consumption and income, but
the same residuals exhibit an I(0) process when the PP unit root test is used, which
implied that the APC for the US is not stationary. Slesnick (1998), like Campbell
(1987) analyzed the stationarity of the APC for the US with the EG cointegration test,
but for the period 1980Q2-1993Q4, by adjusting the personal consumption
expenditure data by using the information comes from Consumer Expenditure
Surveys and concluded that consumption and income are cointegrated. From the
“great ratios” literature, Serletis and Krichel (1995), for 10 OECD countries, including
the UK and the US, applied a Johansen vector autoregressive (VAR) method and
reached a stationary APC for each country.7 Gil-Alana and Robinson (2001)8 used
the HEGY seasonal unit root and Robinson (1994)’s seasonal factional tests for the
consumption and income of the UK and Japan over period 1955Q1-1984Q4 and they concluded that APC’s for the UK and Japan can be best characterized as nonstationary. Okubo (2002) for Japan (1955Q2-1997Q1) reached a conclusion that there is no deterministic cointegration between consumption and income because of a structural trend break by applying cointegration tests with a deterministic cointegration restriction proposed by Park (1990), Ogaki and Park (1998) and Hansen (1992). On the other hand, for Singapore Abeysinghe and Choy (2004) found a nonstationary APC by applying the ADF tests to the residuals from the OLS estimation of consumption equations regressed on income and wealth for the period over 1978Q1-2003Q4. In addition Cook (2003) applied more powerful versions of Dickey-Fuller (DF) tests for the UK over the period 1955(1)-2001(3). These tests are Park and Fuller (1995)’s the weighted symmetric DF test and Shin and So (2001)’s the recursively mean-adjusted DF test. Cook resulted that APC is nonstationary with both tests.

The second group of studies is begun with the work of Jin (1995) who applied Levin and Lin (1992) panel unit root test in addition to univariate EG and Phillips and Qualaris (1990) cointegration tests for 12 OECD countries\(^9\) using annual data for 1960-88 and found an evidence in favor of stationary APC reaching a cointegration between income and consumption.\(^10\) However, Sarantis and Stewart (1999), by direct investigation the stochastic behaviour of the APC applying Im et al. (IPS) (2003) and Taylor and Sarno (1998) panel unit root tests, found a completely inverse test results which fail to reject the joint nonstationarity null for the APC’s of 20 OECD countries\(^11\) over the period 1955-1994. The conflict between these studies has lead to apply newly developed unit root tests aimed to raise statistical power in three different ways: The first way is to apply the panel unit root tests which explicitly allow for error
cross sectional dependence as studied by Romero-Ávila (2008). In this study Romero-Ávila (2008) found a clear confirming evidence for the Sarantis and Stewart (1999) by reaching the existence of a unit-root in the APC’s of 23 OECD economics\textsuperscript{12} over the period 1960-2005 by employing the univariate Ng and Perron (2001), Elliot et al. (1996), Sargan and Bhargava (1983) and KPSS unit root tests as well as the panel unit root tests of Smith et al. (2004) and Pesaran (2003) and a bootstrap version of the panel stationarity test of Hadri (2000). The second way is to apply the panel unit root tests assumes cross sectional dependence with structural breaks as conducted by Romero-Ávila (2009). For the same countries and same period with Romero-Ávila (2008), Romero-Ávila (2009) obtained a very different evidence for a regime-wise stationarity in OECD APC’s with Carrion-i-Silvestre et al. (2005) controlled cross-sectional dependence panel unit root test allow for an unknown number of multiple breaks. The most interesting outcome of Romero-Ávila (2009) is the conflict raised from his findings from panel unit root tests which do not control for structural breaks appear in line with those from previous studies since they are clearly supported the unit root hypothesis by employing IPS, Breitung (2000) and the inverse Chi-square tests of Maddala and Wu (1999) and Choi (2001) assuming cross sectional independence as well as Chang (2002), Breitung and Das (2005), Moon and Perron (2004) and Bai and Ng (2004), which explicitly allow for cross-sectional dependence. The third way is to take account nonlinearities or asymmetries in data generating processes. Nonlinearities in a panel version is thought by Cerrato et al. (2008) by employing heterogeneous nonlinear Cerrato et al (2009)’s panel unit root test which is a direct extension of the time series ESTAR test proposed by Kwiatkowskki et al. (2003)’s and Pesaran (2007)’s linear panel unit root tests, both tests are allowed for cross sectional dependence, for 24 OECD and 33 non-OECD
countries’ APC ratios over the period 1951-2003. Cerrato et al. (2008) concluded that 61% of OECD countries’ tests indicate nonstationarity whereas 68% of non-OECD countries’ results are nonstationary. On the other hand, asymmetries are included in the study made by Tsionas and Christopoulos (2002) for 14 European Union countries for over the period 1960-1999. They thought both in a univariate and in a panel context. In the univariate analysis, they employed Caner and Hansen (2000) unit root test in the context of threshold autoregressive (TAR) models, which account for asymmetries as well as ADF unit root tests. They concluded that while ADF test fails to reject the unit root null of the APC ratios for each country, the unit root test based on TAR models indicate stationary APC’s in at least one regime. In the panel context, Tsionas and Christopoulos (2002) applied both non-asymmetric IPS, Maddala and Wu (1999) and Harriz-Tzavalis (1999) panel unit root tests and a Maddala and Wu (1999) panel version of Caner and Hansen (2000)’s TAR model. They reached the similar evidence with the univariate tests: with non-asymmetrical panel unit root tests the APC ratios exhibit a unit-root processes, when asymmetries are taken account, the European Union’s APC’s are found as a stationary process in at least one regime.

The third group of studies takes into account for structural breaks. One of these studies, Cook (2005), for the same sample of Sarantis and Steawart (1999), investigated the possibility of the failure to reject the unit root hypothesis with structural change in consumption–income ratios for OECD economies and by employing the minimum LM unit root test of Lee and Strazicich (1999 and 2003) which incorporates endogenously determined structural breaks under the null. Cook (2005) concluded that with one or two breaks, the APC ratios for all countries exhibit stationarity as a very stark constrast with the evidence provided by Sarantis and Steawart (1999). Another study is made by Gomes and Franchini (2009) for 10 South American countries over ranging from 1951 to 2003 by using the ADF test as a
benchmark and its panel versions from Maddala and Wu (1999) and Choi (2001). Furthermore, the possibility of structural breaks is taken into account by means of the Minimum LM unit root test with one and two structural break(s) as in Lee and Strazicich (1999) and Lee and Strazicich (2003), respectively. The main findings of the paper, while the individual ADF test and its panel versions found evidence suggesting an integrated APC, after controlling for structural breaks the evidence indicate just the opposite except only Uruguay.

In the last group of empirical studies, a special interest is given to take account the nonlinearities and/or asymmetries in contrast to research employed by panel unit root tests and/or unit root tests with structural breaks. Cook (2002) for instance, consider the asymmetric mean reversion nature of APC with a first order momentum threshold autoregressive (MTAR (1)) unit root analysis for the data of Sarantis and Stewart (1999). It is found that among all the countries, only the APC ratios for Australia, Spain and Finland are mean-reverting and also the APC’s for Australia and Spain exhibit clear asymmetric mean reversion. The studies of Hall et al. (1997) and Paap and van Dijk (2003) thought the effect of nonlinearities in terms of Hamilton (1998)’s Markow Regime Switching method. While Hall et al. (1997) reached an evidence in favor of time-varying cointegration between consumption and income which implies a strong rejection of a linear, temporally stable stationary APC for Japan (1961:1-1987:4), Paap and van Dick (2003) found a linear cointegration between consumption and income which gives a long-run stationary APC, but after having detrended the nonlinear cycles of the series with a Markow trend for the US (1959:1-1999:4).

The next section describes the methodology employed here for testing the stationarity of the APC ratios for 14 transition economies.
III. Methodology

If economic and financial variables exhibit nonlinear behaviour, the standard unit root tests that are based on a linear AR process will have low power (Kapetanios et al., 2003 and Sollis et al. 2002). Further, the standard ADF, PP and KPSS tests do not allow for nonlinearities and/or asymmetries in data generating processes. However, if nonlinear or asymmetric features come from the data generating process, the white noise assumption for residuals is violated. By addressing this issue Kapetanios et al. (2003) (hereafter KSS) provide a test of the hypothesis of a unit root null against an alternative of a nonlinear mean reversion yet globally stationary with an exponential smooth transition autoregressive (ESTAR) process by employing the following ESTAR specification:

$$\Delta y_t = \gamma y_{t-1} 1 - \exp(\theta y_{t-1}^2) + \epsilon_t$$

where $G(\theta, y_{t-1}) = 1 - \exp(\theta y_{t-1}^2), \theta > 0$ is the exponential transition function which is $0 \leq G(\theta, y_{t-1}) \leq 1$ and it has symmetrical U-shaped around zero. $\epsilon_t$ is an identically and independently distributed (iid) sequence with constant variance and zero mean. The parameter $\theta$ gives a measure for the speed of transition between two regimes that correspond to extreme values of the transition function. A direct testing the null hypothesis $H_0: \theta = 0$ against the alternative $H_1: \theta > 0$ with the conventional methods is not feasible, because under this null, the parameter $\theta$ is unidentified. To overcome this identification problem, Kapetanios et al. (2003) computed a first-order Taylor series approximation around $\theta = 0$ by replacing the exponential transition function $G(\theta, y_{t-1})$. The yielding KSS regression is

$$\Delta y_t = \delta y_{t-1}^3 + \epsilon_t$$

(3)
where $e_t = \epsilon_t + R_t$, with the remainder from the Taylor expansion $R_t$ and $\delta = \gamma \theta$. To allow for higher order-dynamics the relevant auxiliary KSS regression can be rewritten as follows:

$$\Delta y_t = \delta y_{t-1}^3 + \sum_{i=1}^{p} \rho_i \Delta y_{t-i} + e_t$$

(4)

To test the unit root null hypothesis, $H_0: \delta = 0$ in Equation (3), Kapetanios et al. (2003) proposed to use the t-statistic: $t_{NL} = \hat{\delta} / \text{se}(\hat{\delta})$ where $\hat{\delta}$ and $\text{se}(\hat{\delta})$ are the OLS estimation and the standard error of $\hat{\delta}$, respectively.

Kapetanios et al. (2003) modify their model given by Equation (4) to deal with a nonzero mean and/or a linear deterministic trend by replacing the raw data for $y_t$, prior to the Equation (3) being estimated, with the de-meaned data $y_t^* = y_t - \bar{y}$ and with the de-meaned and de-trended data $y_t^* = y_t - \hat{\mu} - \hat{\phi} t$, where $\bar{y}$ is the sample mean and $\hat{\mu}$ and $\hat{\phi}$ are the parameters obtained by OLS regressions.

Although KSS test procedure is convenient for testing the null hypothesis of unit root in the case of nonlinear adjustments, Sollis (2009) extends this ESTAR test adding an additional term to capture asymmetry. Sollis (2009) propose a simple unit root null test by allowing symmetric or asymmetric nonlinear mean reversion under the alternative and when the unit root hypothesis is rejected, Sollis (2009) asymmetric ESTAR (AESTAR, hereafter) specification is also used to distinguish between symmetric and asymmetric nonlinearity with a standard F-test. By assuming $y_{t-1}$ as the transition variable, the proposed AESTAR model by Sollis (2009) is

$$\Delta y_t = (1 - \exp(-\gamma_1(y_{t-1}^2)) + \exp(-\gamma_2 y_{t-1}) \rho_1 + (1 - 1 + \exp(-\gamma_2 y_{t-1}) \rho_2^{-1} y_{t-1} + e_t$$

(5)
where \( \varepsilon_t \sim \text{iid} (0, \sigma^2) \). \( G_t(\gamma_1, y_{t-1}) = (1 - \exp(-\gamma_1(y_{t-1}^2)), \gamma_1 \geq 0 \) is the exponential transition function with \( \gamma_1 \) representing the speed of transition and carries the same properties with the KSS transition function in Equation (1). \( S_t(\gamma_2, y_{t-1}) = 1 + \exp(-\gamma_2 y_{t-1}) \), \( \gamma_2 \geq 0 \) is the logistic function which embodies the effects of asymmetries. When \( \rho_1 = \rho_2 = \rho \), the AESTAR model in Equation (5) will be equivalent to the KSS ESTAR model given by Equation (2), so the AESTAR model nests the symmetric KSS ESTAR model.

To clarify the asymmetric structure of the AESTAR model, Sollis (2009) assumed that \( \gamma_1 > 0 \) and \( \gamma_2 \to \infty \). Under this assumption, there are two possibilities for \( y_{t-1} \) depending on the speed of transition parameter \( \gamma_1 \). The first is that \( y_{t-1} \) occurs when \( y_{t-1} \) moves from zero to \(-\infty\) which causes the logistic function \( S_t(\gamma_2, y_{t-1}) \to 0 \). Thus an ESTAR transition occurs between the central regime model \( \Delta y_t = \varepsilon_t \) and the outer-regime model \( \Delta y_t = \rho_2 y_{t-1} + \varepsilon_t \). The second arises if \( y_{t-1} \) moves from zero to \( \infty \), then the logistic function \( S_t(\gamma_2, y_{t-1}) \to 1 \). In this situation, an ESTAR transition occurs between the central regime model \( \Delta y_t = \varepsilon_t \) and the outer-regime model \( \Delta y_t = \rho_1 y_{t-1} + \varepsilon_t \). An explicit implication here is that if \( \rho_1 \neq \rho_2 \), in either side of the attractor which is zero in this case, the autoregressive adjustment will be asymmetric. (As noted above, if \( \rho_1 = \rho_2 = \rho \), the AESTAR turns out to be the symmetric KSS ESTAR, by losing its asymmetric property.) On the other hand, while assuming \( \rho_1 \neq \rho_2 \), but \( \gamma_2 \) takes small and moderate values instead of converging \( \infty \), asymmetry can still be occur and generate a gradual transition of \( S_t(\gamma_2, y_{t-1}) \) between its limiting values 0 and 1. If \( \gamma_2 = 0 \), \( S_t(\gamma_2, y_{t-1}) \to 0.5 \) for every observation period \( t \).
and the composite function in Equation (5),
\[ G_t(\gamma_1, y_{t-1}) S_t(\gamma_2, y_{t-1}) \rho_1 + (1 - S_t(\gamma_2, y_{t-1})) \rho_2, \]
becomes symmetric without depending on the values of \( \rho_1 \) and \( \rho_2 \). So the parameter \( \gamma_2 \) precisely controls the degree of asymmetry for a particular value of \( \rho_2 - \rho_1 \) which provides the feature of the AESTAR model to test the symmetric ESTAR nonlinearity against the asymmetric ESTAR nonlinearity.

A direct testing of a zero mean the unit root null hypothesis against the alternative hypothesis of globally stationary symmetric or asymmetric ESTAR nonlinearity with a unit root central regime, the null for AESTAR model is \( H_0: \gamma_1 = 0 \). However, this null hypothesis leaves the parameters \( \gamma_2, \rho_1 \) and \( \rho_2 \) as unidentified which prevents the use of the conventional methods for testing. To solve this identification problem, Sollis (2009) propose the replacing \( G_t(\gamma_1, y_{t-1}) \) by a a first-order Taylor expansion around \( \gamma_1 = 0 \) in Equation (5) which yields

\[ \Delta y_t = \rho_1 \gamma_1 y_{t-1}^3 S_t(\gamma_2, y_{t-1}) + \rho_2 \gamma_1 y_{t-1}^3 (1 - S_t(\gamma_2, y_{t-1})) + \eta_t \]

(6)

where \( \eta_t = \varepsilon_t + R_t \), with \( R_t \), the remainder from the Taylor expansion. Since for testing the null \( H_0: \gamma_1 = 0 \) in Equation (6), the parameters \( \gamma_2, \rho_1 \) and \( \rho_2 \) are still unidentified, Sollis (2009) proposed a solution for this problem: Firstly defining a new function exhibits the same pattern of nonlinearity \( S_t^*(\gamma_2, y_{t-1}) \) for \( S_t(\gamma_2, y_{t-1}) \) where

\[ S_t^*(\gamma_2, y_{t-1}) = S_t(\gamma_2, y_{t-1}) - 0.5, \]

so that \( S_t^*(0, y_{t-1}) = 0 \) and then taking a Taylor expansion of \( S_t^*(\gamma_2, y_{t-1}) \) around \( \gamma_2 = 0 \) in Equation (5). The yielding AESTAR regression is

\[ \Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \eta_t \]

(7)

The augmented version of Equation (7) is rewritten as following:
$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \sum_{i=1}^{p} \rho_i \Delta y_{t-i} + \eta_t$$ (8)

The unit root null hypothesis becomes $H_0: \phi_1 = \phi_2 = 0$ in Equation (8). For the case of a nonzero mean and/or a linear deterministic trend, it can be applied to replace $y_t$ by the de-meaned data $y_t^* = y_t - \bar{y}$ and with the de-meaned and de-trended data $y_t^* = y_t - \mu - \hat{\phi}t$ prior to OLS estimation of Equation (8), as just proposed for testing the KSS ESTAR. To test the unit root null hypothesis Sollis (2009) proposes a simple $F$-test and denotes the calculated test statistics with $F_{AE, \mu}$ and $F_{AE, t}$ for de-meaned and de-meaned and de-trended data, respectively. Further, if the unit root hypothesis is rejected against the alternative of stationary symmetric or asymmetric ESTAR nonlinearity, using the model given by Equation (8) the null of symmetric nonlinear ESTAR nonlinearity, $H_0: \phi_2 = 0$ is tested against the alternative of asymmetric ESTAR nonlinearity, $H_1: \phi_2 \neq 0$ with a standard $F$-test. Sollis (2009) denotes the calculated test statistics for testing the null of symmetric mean reversion against the alternative of asymmetric mean reversion as $F_{as, \mu}$ and $F_{as, t}$ for de-meaned and de-meaned and de-trended data, respectively.

The next section provides the results of the above unit root tests for the APC ratios for 14 transition economies.

**IV. Data and Test Results**

This paper uses quarterly data to examine the (non)stationarity of the consumption-income ratio which are downloaded from the website of the database of the International Financial Statistics (IFS) of the International Monetary Fund. The

Figure 1 displays the evolution of (the natural logarithm of) APC for each country. Perhaps the most striking feature of the evolution of APC is the clear-cut effect of demising the command economy with the transition and applied inflation stabilization programs. Following the transition years, the APC for each country exhibits a dramatic change, however the direction of this changes differ from economies to economies. Belarus, Bulgaria, Kazakhstan, Romania, and Slovenia have experienced a declining trend in their consumption-income ratios. Estonia has faced with a sharp increase and then a rapid decrease in its APC following the early stages of transition years, but after having a modest increase around 1995, the Estonian APC have continued to slightly fall. Croatia also has faced with a change from a fast decrease to a fast increase around 1998 and then its APC has gradually fallen. On the other hand, while the consumers in Czech Republic, Latvia, Lithuania, and Russia seem trying to keep their consumption-income ratios after transition years, the APC for Poland has modestly increased until 2002 and then turned to be declined by reaching almost the same ratio, which can be interpreted as the APC stayed almost constant. Further, the APC for Hungary after a rapid decrease and the APC for Slovakia following a fast increase and then a sudden fall after the transition seem to be followed a continuously rising paths.
Figure 1. The Evolution of the Consumption-Income Ratios for the Transition Countries

The descriptive statistics for the APC for each country is presented in Table 1. The mean APC ranges from 0.51 to 0.71, while the consumers live in Russia or in Czech Republic consume approximately half of their income, Romanian and Bulgarian consumers needs more than 70% of their income to consume. While the maximum APC is 0.89 from Kazakhstan in 1994Q2, Latvia has the minimum APC with a 0.32 in 1992Q1, which both date exactly, corresponds to the beginning of their inflation stabilization programs. The countries with larger standard deviations and coefficients of variation are Kazakhstan and Latvia which means that the APC for these two countries are more dispersed than the APCs for other countries. On the
other hand, skewness measures the asymmetry in the APC series or of their distributions. The skewness for normal distributions assumed as symmetric is zero. In our sample, except Kazakhstan, Poland, Romania, and Slovenia, the other countries do not seem to have symmetric distributions. Negative skewness, which is seen for Bulgaria, Czech Republic, Croatia, Latvia, and Slovakia indicates that the left tail of the distribution is longer and fatter than the right tail. This can be a result of the pattern where the falls in the APC are more frequent than the rises in the APC. On the other hand, Belarus, Estonia, Hungary, Lithuania, and Russia have positive skewness values for their APC, hence this can be an evidence for a consumption patterns for the consumers of these countries that the right tails of the distributions are fatter, hence the rises in the APCs are more than the falls in the APCs. Further the normal value for kurtosis should be three, but especially for Czech Republic, Estonia, Latvia, Lithuania, and Russia, the kurtosis values of the APCs are very high which means that their distributions for the APC have sharper peaks and fatter tails.

<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Standard Deviation</th>
<th>Coefficient of Variation</th>
<th>Skewness</th>
<th>Kurtosis</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belarus</td>
<td>0.57</td>
<td>0.65</td>
<td>0.51</td>
<td>0.04</td>
<td>0.07</td>
<td>0.13</td>
<td>2.27</td>
</tr>
<tr>
<td>Bulgaria</td>
<td>0.71</td>
<td>0.76</td>
<td>0.65</td>
<td>0.02</td>
<td>0.03</td>
<td>-0.30</td>
<td>3.12</td>
</tr>
<tr>
<td>Czech Republic</td>
<td>0.51</td>
<td>0.62</td>
<td>0.40</td>
<td>0.03</td>
<td>0.06</td>
<td>-0.85</td>
<td>8.05</td>
</tr>
<tr>
<td>Croatia</td>
<td>0.61</td>
<td>0.64</td>
<td>0.57</td>
<td>0.02</td>
<td>0.03</td>
<td>-0.49</td>
<td>2.14</td>
</tr>
<tr>
<td>Estonia</td>
<td>0.55</td>
<td>0.64</td>
<td>0.52</td>
<td>0.02</td>
<td>0.04</td>
<td>1.39</td>
<td>7.80</td>
</tr>
<tr>
<td>Hungary</td>
<td>0.65</td>
<td>0.71</td>
<td>0.61</td>
<td>0.02</td>
<td>0.03</td>
<td>0.32</td>
<td>2.60</td>
</tr>
<tr>
<td>Kazakhstan</td>
<td>0.60</td>
<td>0.89</td>
<td>0.39</td>
<td>0.12</td>
<td>0.20</td>
<td>-0.02</td>
<td>2.17</td>
</tr>
<tr>
<td>Latvia</td>
<td>0.60</td>
<td>0.68</td>
<td>0.32</td>
<td>0.08</td>
<td>0.13</td>
<td>-2.59</td>
<td>9.44</td>
</tr>
<tr>
<td>Lithuania</td>
<td>0.65</td>
<td>0.76</td>
<td>0.60</td>
<td>0.03</td>
<td>0.05</td>
<td>1.47</td>
<td>7.16</td>
</tr>
<tr>
<td>Poland</td>
<td>0.63</td>
<td>0.67</td>
<td>0.59</td>
<td>0.02</td>
<td>0.03</td>
<td>0.08</td>
<td>2.53</td>
</tr>
<tr>
<td>Romania</td>
<td>0.70</td>
<td>0.77</td>
<td>0.62</td>
<td>0.04</td>
<td>0.06</td>
<td>0.08</td>
<td>2.53</td>
</tr>
<tr>
<td>Russia</td>
<td>0.51</td>
<td>0.64</td>
<td>0.45</td>
<td>0.03</td>
<td>0.06</td>
<td>0.68</td>
<td>4.47</td>
</tr>
<tr>
<td>Slovakia</td>
<td>0.56</td>
<td>0.59</td>
<td>0.50</td>
<td>0.02</td>
<td>0.03</td>
<td>-0.95</td>
<td>3.07</td>
</tr>
<tr>
<td>Slovenia</td>
<td>0.56</td>
<td>0.61</td>
<td>0.52</td>
<td>0.02</td>
<td>0.03</td>
<td>0.05</td>
<td>2.38</td>
</tr>
</tbody>
</table>

(1) The statistics given in the first five columns are belong to the APCs without taking their natural logarithms. The last two columns have the statistics of the natural logarithm of the APCs.
Following what have been reached above, it seems likely that linear models that assume normally distributed characteristics cannot be fully employed for the APCs for our sample countries. Nonlinear symmetric and asymmetric models could be the alternative for investigating stochastic properties of the consumption-income ratios.

Now we turn to examination of the stationarity properties of the APC ratios. We first test the stationarity of the APC ratios ignoring possible non-linearities and asymmetries in the series. We employ, for this purpose, the conventional ADF and PP unit root tests which assume that the series under investigations have a unit root under the null hypothesis. We also apply the KPSS unit root test which accepts the series are stationary under the null hypothesis. Table 2 represents the results of these three tests. The results of the ADF and PP do not provide a clear evidence for the stationarity of the APC ratios. Also the rejection of a unit root in the APC ratios is more frequent in PP than in ADF. However, perhaps surprisingly, the KPSS test concludes that for all countries the APC ratios are linearly stationary irrespective a linear time trend is included or not. The possible reason of these conflicting results obtained from linear unit root tests can be ignoring non-linearities and/or asymmetries.
Table 2. Linear Unit-Root Tests Results (1)

<table>
<thead>
<tr>
<th>Country</th>
<th>ADF</th>
<th>PP</th>
<th>KPSS</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>With intercept</td>
<td>With intercept and trend</td>
<td>With intercept</td>
</tr>
<tr>
<td>Belarus</td>
<td>-2.409(0)</td>
<td>-2.806(0)</td>
<td>-2.392(3)</td>
</tr>
<tr>
<td>Bulgaria</td>
<td>-3.002(0)**</td>
<td>-4.071(0)**</td>
<td>-2.780(6)**</td>
</tr>
<tr>
<td>Czech Rep.</td>
<td>-4.504(1)*</td>
<td>-3.125(4)</td>
<td>-5.023(5)**</td>
</tr>
<tr>
<td>Croatia</td>
<td>-1.469(2)</td>
<td>-1.599(2)</td>
<td>-1.417(4)</td>
</tr>
<tr>
<td>Estonia</td>
<td>-2.035(4)</td>
<td>-5.300(1)*</td>
<td>-3.440(3)**</td>
</tr>
<tr>
<td>Hungary</td>
<td>-3.004(0)**</td>
<td>-3.860(0)**</td>
<td>-3.209(4)**</td>
</tr>
<tr>
<td>Kazakhstan</td>
<td>-1.532(1)</td>
<td>-2.679(2)</td>
<td>-1.401(3)</td>
</tr>
<tr>
<td>Latvia</td>
<td>-6.861(1)*</td>
<td>-6.175(1)*</td>
<td>-7.889(3)**</td>
</tr>
<tr>
<td>Lithuania</td>
<td>-2.774(4)**</td>
<td>-3.929(4)**</td>
<td>-4.206(5)*</td>
</tr>
<tr>
<td>Poland</td>
<td>-1.109(3)</td>
<td>-1.314(4)</td>
<td>-2.211(5)</td>
</tr>
<tr>
<td>Romania</td>
<td>-1.437(0)</td>
<td>-2.643(0)</td>
<td>-1.300(4)</td>
</tr>
<tr>
<td>Russia</td>
<td>-2.560(0)</td>
<td>-2.563(0)</td>
<td>-2.728(2)**</td>
</tr>
<tr>
<td>Slovakia</td>
<td>-1.523(1)</td>
<td>-2.146(4)</td>
<td>-3.162(3)**</td>
</tr>
<tr>
<td>Slovenia</td>
<td>-1.335(0)</td>
<td>-3.031(9)</td>
<td>-1.395(3)</td>
</tr>
</tbody>
</table>

(1) ***, **, * denote the rejection of the null hypothesis at 10 %, 5 % and 1 % significance levels, respectively. Number of augmentation terms in the ADF test was selected using Schwarz Information Criteria (SIC) and is indicated in parenthesis.

Next we turn our interest to the KSS ESTAR (Kapetanios et al., 2003) unit root test allow a smooth transition-type non-linear data generating process. Table 3 contains the results of the KSS ESTAR unit root test. Prior to applying KSS ESTAR test, de-meaned and de-meaned and de-trended data were obtained by saving the residuals from the OLS regressions of the APC series of each country on a constant and on both a constant and a time trend, respectively. As reveal from Table 2, while ADF does not reject the existence of a unit root in the APCs for Czech Republic, Estonia, Kazakhstan, and Russia irrespective whether a linear time trend is included or not, the KSS ESTAR test indicate a nonlinear global stationarity for these countries. However even if the ADF, PP and KPSS could not reject the linear stationarity of the APC ratios for Hungary and Latvia, the KSS ESTAR test concluded that the APC ratios for both countries follows a nonlinear globally stationary long-run path. Hence, among all countries, the KSS ESTAR test provides an evidence for nonlinear stationary APC for Czech Republic, Estonia, Hungary, Kazakhstan, Latvia, and Russia.
Table 3. KSS ESTAR Unit Root Test Results (1)

<table>
<thead>
<tr>
<th>Country</th>
<th>Demeaned Data</th>
<th>Demeaned and Detrended Data</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belarus</td>
<td>-2.555(0)</td>
<td>-2.440(0)</td>
</tr>
<tr>
<td>Bulgaria</td>
<td>-1.928(1)</td>
<td>-1.611(1)</td>
</tr>
<tr>
<td>Czech Rep.</td>
<td>-7.229(5)*</td>
<td>-7.214(5)*</td>
</tr>
<tr>
<td>Croatia</td>
<td>-1.738(4)</td>
<td>-1.703(0)</td>
</tr>
<tr>
<td>Estonia</td>
<td>-5.588(4)*</td>
<td>-4.166(9)*</td>
</tr>
<tr>
<td>Hungary</td>
<td>-3.955(0)*</td>
<td>-4.219(0)*</td>
</tr>
<tr>
<td>Kazakhstan</td>
<td>-2.858(1)**</td>
<td>-7.336(0)*</td>
</tr>
<tr>
<td>Latvia</td>
<td>-2.901(3)***</td>
<td>-3.163(3)***</td>
</tr>
<tr>
<td>Lithuania</td>
<td>-1.421(5)</td>
<td>-1.393(5)</td>
</tr>
<tr>
<td>Poland</td>
<td>-2.087(3)</td>
<td>-2.042(3)</td>
</tr>
<tr>
<td>Romania</td>
<td>-0.904(0)</td>
<td>-2.562(0)</td>
</tr>
<tr>
<td>Russia</td>
<td>-3.262(1)**</td>
<td>-3.732(1)**</td>
</tr>
<tr>
<td>Slovakia</td>
<td>-2.131(1)</td>
<td>-0.986(8)</td>
</tr>
<tr>
<td>Slovenia</td>
<td>-1.369(0)</td>
<td>-0.271(4)</td>
</tr>
</tbody>
</table>

Asymptotical Critical Values of t<sub>NL</sub> statistic: (2)

<table>
<thead>
<tr>
<th></th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>-3.48</td>
<td>-2.93</td>
<td>-2.66</td>
</tr>
<tr>
<td></td>
<td>3.93</td>
<td>3.40</td>
<td>3.13</td>
</tr>
</tbody>
</table>

(1) ***, *, * denote rejection of the null hypothesis at 10 %, 5 % and 1 % significance levels, respectively. The lag lengths were chosen by SIC and is indicated in parenthesis.

(2) Critical values of t<sub>NL</sub> statistic are taken from Table 1, Kapetanios et al. (2003: 364).

We now turn to the results of Sollis (2009) asymmetrical ESTAR (AESTAR) unit root test, which is presented in Table 4. The AESTAR test is applied for the same de-meaned and de-meaned and de-trended series obtained for the KSS ESTAR test. In Table 4, the test statistics F<sub>AE,μt</sub> and F<sub>AE,t</sub> corresponds to test the unit root null hypothesis for de-meaned and de-meaned and de-trended data, respectively. Further, as standard F-tests, F<sub>as,μ</sub> and F<sub>as,t</sub> statistics are calculated only when the unit root hypothesis is rejected against the alternative of stationary symmetric or asymmetric ESTAR nonlinearity, since these statistics are for testing the null of symmetric nonlinear ESTAR nonlinearity against the alternative of asymmetric ESTAR nonlinearity where the subscript m is used for de-meaned and t is for de-meaned and de-trended data. The rejection of unit root null hypothesis with AESTAR test is more frequent than the KSS ESTAR test, since the AESTAR test results suggest that Bulgaria and Slovenia also have nonlinear stationary APC ratios, additionally other 6 countries, namely Czech Republic, Estonia, Hungary,
Kazakhstan, Latvia, and Russia. The AESTAR test further provides evidence about asymmetric nature of the APC ratios for five countries, Bulgaria, Czech Republic, Estonia, Latvia, and Russia out of 14 economies in the sample.¹⁸

Table 4. AESTAR Unit Root Test Results

<table>
<thead>
<tr>
<th>Country</th>
<th>Fₐₑₘₜ</th>
<th>Fₐₛₘₜ</th>
<th>Fₐₑₜ</th>
<th>Fₐₛₜ</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belarus</td>
<td>3.358(0)</td>
<td>NA</td>
<td>4.377(0)</td>
<td>NA</td>
</tr>
<tr>
<td>Bulgaria</td>
<td>3.432(2)</td>
<td>NA</td>
<td>8.861(0)*</td>
<td>13.402*</td>
</tr>
<tr>
<td>Czech Rep.</td>
<td>27.413(5)*</td>
<td>1.862</td>
<td>30.921(5)*</td>
<td>5.856**</td>
</tr>
<tr>
<td>Croatia</td>
<td>1.475(4)</td>
<td>NA</td>
<td>1.451(0)</td>
<td>NA</td>
</tr>
<tr>
<td>Estonia</td>
<td>1.215(11)</td>
<td>NA</td>
<td>9.091(8)*</td>
<td>0.723</td>
</tr>
<tr>
<td>Hungary</td>
<td>8.691(0)*</td>
<td>1.584</td>
<td>9.221(0)</td>
<td>17.190*</td>
</tr>
<tr>
<td>Kazakhstan</td>
<td>4.680(1)***</td>
<td>1.168</td>
<td>27.586(0)*</td>
<td>1.189</td>
</tr>
<tr>
<td>Latvia</td>
<td>8.474(2)*</td>
<td>9.052*</td>
<td>4.924(3)</td>
<td>NA</td>
</tr>
<tr>
<td>Lithuania</td>
<td>0.751(6)</td>
<td>NA</td>
<td>0.757(6)</td>
<td>NA</td>
</tr>
<tr>
<td>Poland</td>
<td>2.658(3)</td>
<td>NA</td>
<td>2.824(3)</td>
<td>NA</td>
</tr>
<tr>
<td>Romania</td>
<td>0.570(0)</td>
<td>NA</td>
<td>3.946(0)</td>
<td>NA</td>
</tr>
<tr>
<td>Russia</td>
<td>8.224(1)*</td>
<td>5.112**</td>
<td>9.224(1)*</td>
<td>3.859***</td>
</tr>
<tr>
<td>Slovakia</td>
<td>2.704(1)</td>
<td>NA</td>
<td>0.524(8)</td>
<td>NA</td>
</tr>
<tr>
<td>Slovenia</td>
<td>5.531(3)**</td>
<td>1.916</td>
<td>0.371(4)</td>
<td>NA</td>
</tr>
</tbody>
</table>

Asymptotical Critical Values for T=100 (³)

<table>
<thead>
<tr>
<th></th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>1%</td>
<td>6.883</td>
<td>8.531</td>
<td>5.460</td>
</tr>
<tr>
<td>5%</td>
<td>4.954</td>
<td>6.463</td>
<td></td>
</tr>
<tr>
<td>10%</td>
<td>4.157</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

(1) ***, **, * denote rejection of the null hypothesis at 10 %, 5 % and 1 % significance levels, respectively. The lag lengths were chosen by SIC and is indicated in parenthesis.
(2) NA because the unit root hypothesis is not rejected.
(3) Critical values are taken from Table 1, Sollis (2009: 121).

V. Discussion and Conclusion

This paper provides a contribution to the ongoing debates concerning whether the consumption-income ratio is stationarity or not by giving a special attention to the possibilities of nonlinearities and/or asymmetries in data generating process for fourteen transition economies over a period after collapsing the command economy. For this purpose, in addition to conventional linear and symmetrical ADF, PP, and KPSS unit root tests, the Kapetanios et al. (2003) KSS ESTAR unit root test, which allows a nonlinear symmetric global stationarity in data generating process and the Sollis (2009) AESTAR test, which also adds asymmetric behaviour to the KSS ESTAR test are applied. Without taking account for
nonlinearities, while the conventional ADF and PP tests could not fail to reject the unit root hypothesis for Belarus, Czech Republic, Croatia, Estonia, Kazakhstan, Poland, Romania, Russia, Slovakia, and Slovenia, only for four countries namely Bulgaria, Hungary, Latvia and, Lithuania these two tests found an evidence of stationary APC unanimously. However, interestingly, the KPSS test concluded that the APC ratios for all countries are linearly stationary irrespective whether a linear time trend is included or not. However, when the KSS ESTAR test applied, the results showed that the APCs for Czech Republic, Estonia, Hungary, Kazakhstan, Latvia and Russia are globally nonlinearly nonstationary. Additionally the Sollis (2009) AESTAR test concludes that the APCs for Bulgaria and Slovenia also exhibit a nonlinear stationarity. Hence, the acceptance of stationarity hypothesis is more frequent in nonlinear unit root tests than the ADF and PP tests. As a conclusion, while Bulgaria, Croatia, Lithuania, Poland, Romania, and Slovakia have linearly stationary consumption-income ratios in accordance with the KPSS test, the APC for Bulgaria, Czech Republic, Estonia, Hungary, Kazakhstan, Latvia, Russia and, Slovenia exhibit a nonlinear stationarity. Further, the Sollis (2009) AESTAR test provide an evidence of an asymmetric behaviour for the APC ratios for five countries namely Bulgaria, Czech Republic, Estonia, Latvia, and Russia.

Since the consumption-income ratios for all countries are found as stationary, ‘the big bangs’ or the shocks happened by moving from command to market economies after the transition periods are seem to be absorbed by distributing the effects of these in time by the consumers of these economies, hence it can be concluded that the existence of mean reversion implies that policy shocks are likely to have temporary effects on the average propensity to consume in these fourteen transition countries. On the other hand, the test results imply that the inclusion of the durable component in examining the consumption-income ratio can affect the outcomes of the linear conventional unit root tests, since the existence of durable
consumption can be addressed as a direct source of nonlinear and asymmetric adjustment for consumption to changes in income. Indeed, when the consumption on durable goods is taken under investigation, especially, in the presence of imperfect capital markets (liquidity constraints), uncertainties and risks about the asset returns affecting the wealth and any change in the expectations about future, an immediate responds of consumers can be thought as altering expenditures on durable goods for keeping their standard of living, i.e. their consumption-income ratio, same. Therefore, while researching the stationarity of the consumption-income ratio, the use of standard unit root tests should be cautiously handled about possibility of a presence of a nonlinear and an asymmetric adjustment. However, of course, the reasons of nonlinearities and –especially asymmetries in five countries- in consumption are a subject matter for further researches.

The results obtained here appear to be congruent with the Relative Income Hypothesis, the Permanent Income Hypothesis, and the Life-Cycle Hypothesis which all assume a forward-looking consumer. It can be implied that households of these economies do not exhibit myopic consumer behaviour, their time horizon is much longer than the Absolute Income hypothesis assumed, and hence they try to smooth their consumption pattern in their life time. Additionally, since the average propensities to consume of the consumers in these countries are stationary, even if it has little effect, monetary policy is likely be more efficient in affecting the current consumption than fiscal policy as assumed by the Permanent Income Hypothesis, and the Life-Cycle Hypothesis. Thus, the consumers is likely be more respond to changes in monetary policies in transition countries, which is another subject for further research.
Notes

1. “Prior to Keynes, consumption had been viewed as a passive residual, the amount of income remaining after saving. In this view, since the decision to save was determined by the payment for the utility lost from consuming, consumption was depended on the interest rate.” (Bunting, 2001:149) and "(t)here are not many people who will alter their way of living because the rate of interest has fallen from 5 to 4 percent” (Keynes, 1936: 94). Consumption was both more important and more complicated for Keynes, since it fundamentally affects the level of economic activity (Skidelsky, 1997: 311-12).

2. We can use a simplified linear graph to show what the consumption function (1) of the AIH looks like and how its two features relate consumption to income.

From the graph it should be clear that firstly the MPC (the slope of the consumption function) is less than the APC (the slope of a line from the origin to a point on the consumption function) for all levels of income, and secondly the APC varies as income changes, so APC is not mean-reverting and also consumption will not rise proportionally to income. The Keynesian model, \( C_t = a + bY_t \), when transformed into an average consumption model, \( C_t/Y_t = a/Y_t + b \), predicts a long run declining spending rate and rising saving rate. That is, increases in \( Y_t \) cause \( a/Y_t \) to decline and the average propensity to consume should fall.

3. For the detailed explanations, see Thomas (1989).

4. Following Kuznets (1946), for instance Ackley (1961), and Bunting (1989) reached the supporting views of Kuznets.

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5. Palley (2010) makes a contribution to synthesize the AIH, the RIH and the PIH in a consumption theory which is called the ‘relative permanent income’ theory. The key feature of this theory is that the share of permanent income devoted to consumption is a negative function of household relative permanent income. This model also assumes a constant APC.

6. In explanation of the ratchet effect, the LCH assumes that in the short-run the APC will be a decreasing function of current income and the increases in wealth causes an upward shift in the consumption functions (Modigliani and Brumberg, 1954). However, in the long-run as the income of all people increases by causing the total value of the assets they hold tends to rise, we would expect the APC to be roughly constant. This implies that there will be no significant change in the APC over time. The PIH on the other hand, explains the consumption puzzle by suggesting the difference between permanent and transitory components of income. It emphasizes that people experience temporary changes in their income from year to year so they try smoothing their consumption. To the PIH, if much of year-to-year variation in income is transitory, the APC becomes a decreasing function of income. Friedman (1957) assumed that households with high income will have lower than average APC, since with a given level of permanent income, those people’s transitory income levels are high and these temporary increases in income will not give a rise to their current consumption. On the other side, low income households will have higher than average APC, because low or negative levels of transitory incomes will be directed to consumption. However, in the long-run time series data, variations in income will be mainly permanent so the positive and negative transitory income shocks can be thought as cancelling each other out; therefore, consumers do not save any increases in income, but consume them instead. This means that changes in
consumption will be proportional to the changes in income and so the APC will be almost constant in the long-run.


8. This study is belong to the Great-ratios literature.

9. The countries included in Jin (1995) are Australia, Austria, Belgium, Canada, Finland, Germany, Greece, Japan, Norway, Switzerland, the UK, and the US.

10. It should be noted that Jin (1995) found a nonstationary APC for Austria, Greece, and the US with univariate Engle and Granger cointegration test.

11. Sarantis and Stewart (1999) widened Jin (1995)’s sample by adding the countries of Denmark, France, Iceland, Ireland, Italy, the Netherlands, Spain and Sweden.


13. Cerrato et al. (2008) excluded Ireland and included Mexico and Turkey among OECD countries to the sample by taken Romero-Ávila (2008, 2009) and also they widened their sample by adding the non-OECD countries: Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, Egypt, El Salvador, Ethiopia, Guatemala, Honduras, India, Israel, Kenya, Mauritius, Morocco, Nicaragua, Nigeria, Pakistan, Panama, Paraguay, Peru, Philippines, South Africa, Sri Lanka, Taiwan, Thailand, Trinidad and Tobago, Uganda, Uruguay, Venezuela.
The countries included by Tsionas and Christopoulos (2002) are Austria, Belgium, Denmark, Finland, France, Greece, Iceland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden and the UK, which are all placed in the samples of Romero-Ávila (2008, 2009) and Cerrato et al. (2008).

These countries are Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Paraguay, Peru, Uruguay and Venezuela which all included in the sample of Cerrato et al. (2008).

For a detailed explanation of the effects of the inflation stabilization programs on macroeconomics of the transition economics see Fischer et al. (1996) and Fischer and Sahay (2000).

For comparative purposes, the mean APC is 0.63 for OECD-Total, 0.59 for OECD-Europe and 0.64 for Major Seven Countries for period 1991:1-2009:4 and 0.58 for European Union during the period from 1995:1 to 2009:4 with the data which are taken from the website of National Accounts statistics of OECD. Further, for individual OECD countries the mean APCs are 0.59 for Australia and Germany, 0.56 for Austria, 0.58 for Canada, 0.50 for Denmark and the Netherlands, 0.53 for Finland, 0.57 for France, 0.60 for Italy, Japan, New Zealand and, Switzerland, 0.47 for Norway, 0.54 for Korea, 0.65 for the UK and 0.69 for the US over the period 1991:1-2009:4. The mean APC is also 0.49 for Sweden for 1993:1-2009:4, 0.53 for Belgium, 0.39 for Luxembourg, 0.65 for Portugal and 0.58 for Spain for 1995:1-2009:4, 0.70 for Turkey over 1998:1-2009:4 and 0.72 for Greece over 2000:1-2009:4.

We also tried to apply Leybourne et al. (1998; LNV) unit root test which allows for smooth structural changes in the data generating process. However, whenever we tried to estimate the OLS regressions of the LNV model, we faced with a message of near singularity. Leybourne et al. (1998: 87) suggested that near
singularity means there is no transition in the series, hence applying the ADF test is more suitable.

**References**


