Current Account Sustainability in G7 and BRICS: Evidence from a Long Memory Model with Structural Breaks[#,**]

Christophe André[♣], Mehmet Balcilar[♥], Tsangyao Chang[♦], Luis A. Gil-Alana[♠] and Rangan Gupta[*]

Abstract
In this paper, we extend the existing literature on the sustainability of current account deficits by examining the relevance of long memory and structural breaks in modelling the dynamics of current account to GDP ratios. Unlike standard unit root tests, which can only indicate whether a series is stationary or not by looking at 0 or 1 for the orders of integration and which have low power, especially in cases where the series is characterized by a fractional process, the long memory approach provides an exact measure of the degree of persistence. However, long memory models are known to overestimate the degree of persistence of the series in the presence of structural breaks, which are very likely in quarterly macroeconomic data covering a long period. Indeed, we show that regime changes do exist in both the mean and trend of the current account to GDP ratios. Thus, we test persistence allowing for both smooth and sharp breaks. Our methodology also allows us to include any number of sharp breaks, whereas standard unit root tests only permit either one or two breaks. Hence, our approach is more general and more robust to misspecifications caused by the omission of breaks than standard methods. To the best of our knowledge, this is the first paper testing for the sustainability of current account balances in the seven major advanced economies (G7) and the BRICS countries using long-memory models incorporating both smooth and sharp breaks. We show that current accounts are sustainable in both groups of countries, with the G7 and South Africa displaying long-memory behavior.

JEL Codes: C22, F32

Keywords: Current account, sustainability, long-memory, smooth and sharp breaks, G7, BRICS

[#]The views expressed in this paper are those of the authors and do not necessarily reflect those of the Organisation for Economic Co-operation and Development (OECD) or the governments of its member countries.

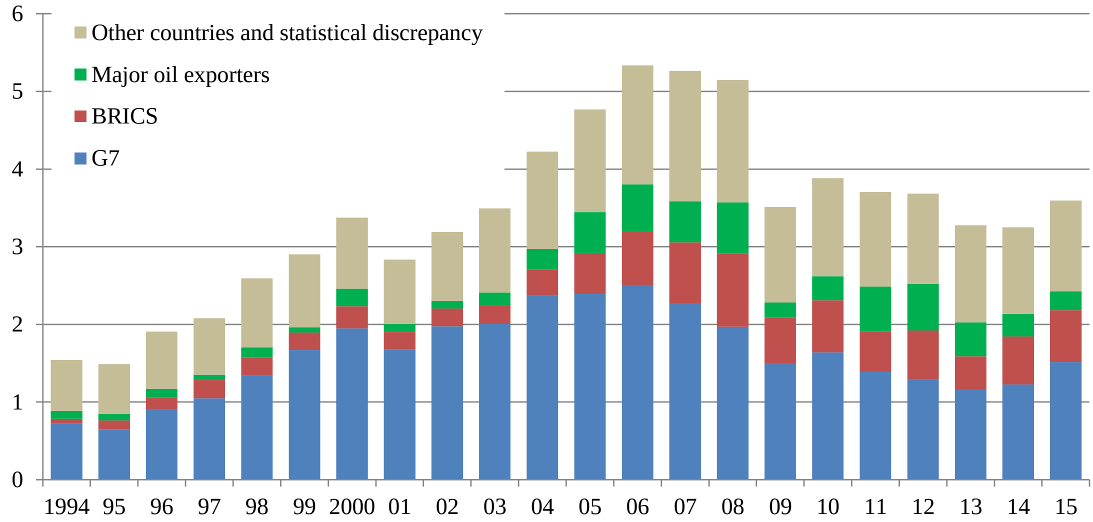
[**] We would like to thank two anonymous referees for many helpful comments. However, any remaining errors are solely ours.

[♣]Economics Department, Organisation for Economic Co-operation and Development (OECD), 75775 Paris, Cedex 16, France. Email: christophe.andre@oecd.org.
[♥] Department of Economics, Eastern Mediterranean University, Famagusta, via Mersin 10, Northern Cyprus, Turkey; Montpellier Business School, Montpellier, France; and Department of Economics, University of Pretoria, Pretoria, 0002, South Africa. Email: mehmet@mbalcilar.net.
[♦]Department of Finance, College of Finance, Feng Chia University, Taichung, Taiwan. Email: l: tychang@fcu.edu.tw.
[♠]Corresponding author. University of Navarra, Faculty of Economics and ICS (NCID), Edificio Amigos, E-31080 Pamplona, Spain. Email: alana@unav.es.
[*]Department of Economics, University of Pretoria, Pretoria, 0002, South Africa. Email: rangan.gupta@up.ac.za.

1. Introduction

Global current account imbalances have become a major issue in the international economic policy debate, especially within the Group of Twenty (IMF, 2015). In an increasingly integrated global economy, current account imbalances are to be expected, as capital can move internationally to finance the most promising investments. There is no reason to assume that in an open economy, domestic investment, in the short-term, should equal domestic saving, which is equivalent to having a balanced current account. However, countries face a long-term budget constraint (Trehan and Walsh, 1991; Taylor, 2002) and there is widespread concern that a disorderly unwinding of imbalances could lead to disruptions in the global financial system and economy (Eichengreen, 2011; De Mello et al., 2012, IMF, 2014). These concerns reflect both considerations related to the stability of the capital flows which are financing the current account deficits and more structural issues, for example that external imbalances may reflect unsustainable developments in domestic private or public debt. Three groups of countries account for the bulk of global current account imbalances: the major advanced economies (G7), the leading emerging economies (BRICS) and the major oil exporters (Figure 1). In this paper, we focus on the sustainability of current account imbalances in the two first groups, since current account dynamics in the third group are mainly driven by oil prices.

Figure 1: Global current account imbalances by group of countries



Note: Global imbalances are calculated as the sum of the absolute values of current account balances of countries for which data are available in the World Bank's World Development Indicators database. The database covers most of the world economy, but data for some (generally small) countries are missing, implying that global imbalances may be slightly underestimated. Major oil exporters include Algeria, Angola, Azerbaijan, Bahrain, Brunei, Chad, Republic of Congo, Ecuador, Equatorial Guinea, Gabon, Iran, Iraq, Kazakhstan, Kuwait, Libya, Nigeria, Oman, Qatar, Saudi Arabia, Sudan, Timor-Leste, Trinidad and Tobago, Venezuela and Yemen.

Stationarity of the current account would ensure that the long run national budget constraint is met. Standard unit root tests have been widely used in the literature to test for the stationarity of the current account. However, current account dynamics are characterized by high persistence and structural breaks. The former relates to structural factors affecting the saving-investment balance, for example financial market development, social protection and demographics, while the latter reflects the impact of different types of events, for example oil price shocks, sovereign debt crises or financial deregulation. The literature shows that standard unit root tests, which can only indicate whether a series is stationary or not by looking at orders of integration of 0 or 1, have low power in the case of a highly persistent, or long-memory time series, which are characterized by a fractional process (Diebold and Rudebusch, 1991; Hassler and Wolters, 1995; Lee and Schmidt, 1996; and more recently, Ben Nasr et al., 2014). Hence, we adopt the long memory approach, which enables an exact measurement of the degree of persistence and of the time span that it would take for a shock to the current account balance to die off, if at all. However, long memory models are known to overestimate the degree of persistence of the series in the presence of structural breaks (Cheung and Lai, 1993; Diebold and Inoue, 2001; Ben Nasr et al., 2014). Therefore, we account for structural breaks in testing for persistence. We also allow for two types of structural breaks. Indeed, while some regime changes may be instantly reflected in the current account (sharp break), some adjustments may be more protracted, especially when using a relatively high data frequency, such as quarterly (smooth break). Our methodology also allows us to model any number of sharp breaks, unlike standard unit root tests which only permit either one (Zivot and Andrews, 1992) or two breaks (Lumsdaine and Papell, 1997; Lee and Strazicich, 2003). Altogether, our approach is more general and robust to misspecifications than standard stationarity tests, as it is able to accommodate long memory processes, as well as to capture an ex ante unspecified number of both smooth and sharp breaks. To the best of our knowledge, this is the first paper to test for the sustainability of current account balances in the G7 and BRICS countries using long-memory models incorporating both smooth and sharp breaks.

We find three important results. First, the order of integration is below one for all the countries of the sample, which rules out explosive dynamics. Hence, current accounts are sustainable in all BRICS and G7 countries. Second, the order of integration varies significantly across countries. Current account to GDP ratios in the G7 countries and South Africa exhibit long-memory behavior, i.e. while mean-reversion does take place, the disequilibrium takes a while to be corrected. Conversely, current account to GDP ratios in Brazil, China, India and Russia are stationary. Hence, imbalances tend to be corrected more rapidly in these countries. Third, the degree of persistence of current account imbalances does not appear to be related to the sign of the imbalance, whereas financing constraints would suggest that countries with current account deficits are facing more pressure to adjust than countries running surpluses.

2. Review of the literature

Several studies show that current account positions are the outcome of inter-temporal choices of households, firms and governments, with the national economy facing an inter-temporal budget constraint, as foreign lenders will be unwilling to finance external deficits indefinitely (Buiter, 1981; Sachs, 1981; Obstfeld and Rogoff, 1995). More specifically, a number of articles demonstrate that the stationarity of the current account to GDP ratio is a sufficient condition for the inter-temporal national long-run budget constraint to hold (Trehan and Walsh, 1991; Wickens and Uctum, 1993; Coakley et al., 1996; Taylor, 2002). This result has paved the way for an abundant empirical literature using unit root tests to assess the sustainability of current account positions. The results are mixed and vary across country and time samples, as well as with the methodology employed (for a survey of the recent literature, see Chen, 2011a,b). Traditional unit root tests often fail to reject the presence of a unit root in the current account to GDP ratio (Nason and Rogers, 2006; Cuñado et al., 2010). Hence, researchers have turned to more sophisticated techniques to investigate current account dynamics. In particular, standard unit root tests are known to have low power in cases where the series is characterized by structural breaks (Perron, 1989, 1994, 1997; Zivot and Andrews, 1992; Lumsdaine and Papell, 1997; Lee and Strazicich, 2003, Enders and Lee, 2012) or a fractional process (Diebold and Rudebusch, 1991; Hassler and Wolters, 1995; Lee and Schmidt, 1996; Ben Nasr et al., 2014). A number of papers have addressed these issues.[1] Starting with studies allowing for regime shifts, Apergis et al. (2000) use a variety of unit root and cointegration tests allowing for structural breaks and find evidence of sustainability of the Greek current account deficit over the period 1960-1994. Baharumshah et al. (2003), using the Gregory and Hansen (1996) procedure, which allows for regime shifts, find that external imbalances in the period preceding the 1997 Asian crisis were unsustainable in

[1] Another approach has been the use of panel unit root tests (see for example Wu, 2000; Wu et al., 2001; Lau et al., 2006; Gnimassoun and Coulibaly, 2014).

Indonesia, the Philippines, and Thailand, but sustainable in Malaysia. Herzer and Nowak-Lehmann (2006), applying the same procedure, find that Chile's current account was sustainable over the period 1975-2004, despite the balance of payments crisis of the early 1980s. Clarida et al. (2006) find evidence of threshold effects and identify three different adjustement regimes in current accounts (surplus adjustment, deficit adjustment and inertia) in G7 countries over the period 1979-2003. Christopoulos and Leon-Ledesma (2010) find stationarity in the US current account over the period 1960-2004, using an exponential smooth transition (ESTAR) model. Onel and Utkulu (2006) find that with or without considering any structural break using the Zivot-Andrews (Zivot and Andrews, 1992) and Gregory-Hansen (Gregory and Hansen, 1996) tests, Turkish external debt was weakly sustainable over the period 1970-2002. Similarly, Topalli and Dogan (2016), using a Markov-switching model, find weak sustainability in the Turkish current account balance over the period 1990-2014. Conversely, Cecen and Xiao (2014), after establishing the nonlinearity of the Turkish current account to GDP ratio over the period 1987-2011, find that the ratio is nonstationary, using the threshold unit root test of Caner and Hansen (2001). Raybaudi et al. (2004) propose a procedure for identifying periods under which the current account accumulates at a non-stationary rate, based on imposing identifying restrictions in Markov switching type models. Applying their procedure on samples which vary across countries, but all end in the early 2000s, they find evidence that the long run national budget constraint is met in Brazil, Japan and the United Kingdom, but may not be in Argentina and the United States. Clower and Ito (2012) find that allowing current account series to follow a Markov switching process leads to rejecting the unit root hypothesis for a much larger part of the 70 countries in their sample, than when using standard linear unit root tests. Chen (2011a), using Markov switching models, finds that the long term national budget constraint is unlikely to hold for the major seven advanced economies (G7), except Germany and Japan, over samples ranging from the early 1970s to the late 2000s, except for Germany where the period covered starts after reunification. Chen (2011b) applies the same methodology to a set of eight smaller OECD countries and finds evidence of non-stationarity in most cases. Chen (2014) tests three types of nonlinearity, namely structural breaks, sign nonlinearity (asymmetric adjustments) and size nonlinearity (threshold effects) in ten mostly European countries over different samples ranging from the 1970s to 2012. He finds strong evidence of structural breaks and size nonlinearity, while sign nonlinearity is absent. Taking nonlinearities into account, he finds stationary in seven of the ten countries of the sample. Chen and Xie (2015) expand on this analysis by proposing a test for a unit root against the alternative that encompasses smooth breaks, size nonlinearity and sign asymmetry simultaneously and find broadly similar results. De Mello and Padoan (2010), using the Lee and Strazicich (2003) procedure, find that the current account to GDP ratios of the United States, China, Japan, Germany and the major oil-exporting countries are stationary around structural breaks, which are associated with shifts in the fiscal stance, exchange rate parities or potential output growth. De Mello et al. (2012) use the same methodology to tests for unit roots in the current account to GDP ratios of a large set of mature and emerging-market economies. They find that the external positions of most countries are stationary around structural breaks.

Turning to studies using fractional integration techniques, Dulger and Ozdemir (2005) study the sustainability of current accounts in G7 countries over the period 1974-2001 and find that the current account is covariance nonstationary in all countries. It is, however, mean-reverting in France, Italy and Canada, but not in other countries. Cuñado et al. (2010) find little evidence of mean-reversion in a sample of 16 European countries over the period 1960-2005. Conversely, Cuestas (2013) finds that current account to GDP ratios are mostly stationary in European transition economies over the period 1999-2011. Since long memory models are known to overestimate the degree of persistence of the series in the presence of structural breaks (Cheung and Lai, 1993; Diebold and Inoue, 2001; Ben Nasr et al., 2014), some studies have accounted for structural breaks in fractional integration models. Cuñado et al. (2008) show that when a break in 1983 is introduced, the US current account is sustainable over the period 1960-2006, although it displays long-memory behaviour. Kiran (2012) finds that, even when allowing for structural breaks, the Turkish current account is not stationary over the period 1970-2010. Chen (2013) finds that using the Markov switching autoregressive fractionally integrated moving average model (MS-ARFIMA) of Tsay and Härdle (2009) allows dealing with the sensitivity to the sample period found in Markov switching unit root regressions on the US current account to GDP ratio over the period 1970-2012. The MS-ARFIMA results suggest that the current account deficits observed were unsustainable.

3. Methodology

We use techniques based on the concepts of long range dependence or long memory processes, and more specifically fractional integration. The idea behind this concept is that the number of differences required to render a series stationary I(0) may not necessarily be an integer number (usually 1), but rather a fractional value.

For the purpose of this work, we say that a process $\{u_t, t = 0, \pm 1, \ldots\}$ is integrated of order 0 (and denoted as I(0)) if the infinite sum of the autocovariances is finite. Alternatively, it can be defined in the frequency domain by saying that it is a process with a spectral density function that is positive and finite at the zero frequency. There are many processes that may be included within this class, such as the white noise or stationary ARMA-type processes. Then, we say that a process $\{x_t, t = 0, \pm 1, \ldots\}$ is integrated of order d (and denoted as I(d)) if after d-differences the series becomes I(0). In other words, x_t is I(d) if

$$(1 - L)^d x_t = u_t, \qquad t = 0, \pm 1, \ldots, \tag{1}$$

and u_t is I(0), and where L denotes the lag-operator ($Lx_t = x_{t-1}$). Note that if u_t in (1) is an ARMA(p, q) process, we say then that x_t is a fractionally integrated ARMA or ARFIMA(p, d, q) process.

The fractional differencing parameter d as given in equation (1) plays a crucial role from different perspectives. From a statistical (and semantical) viewpoint, if d = 0, x_t is said to be short memory, as opposed to the long memory case when d > 0. Also, if d < 0.5, x_t is still covariance stationary, while it is nonstationary for d ≥ 0.5. Finally, from an economic perspective, if d < 1, the series is mean reverting with shocks having transitory effects and disappearing in the long run. On the contrary, if d ≥ 1 mean reversion does not occur and shocks will have a permanent effect on the series.

In the empirical application carried out in the following section we consider a model of the form:

$$y_t \;=\; \beta_0 \;+\; \beta_1 t \;+\; x_t, \qquad t \;=\; 1, 2, ..., \tag{2}$$

where y_t is the observed time series, β_0 and β_1 are the coefficients corresponding to the intercept and the linear time trend, and x_t is an I(d) process defined as in (1). We estimate d along with the other parameters of the model by means of the Whittle function in the frequency domain (Dahlhaus, 1989) along with a testing procedure developed by Robinson (1994) based on the LM principle and which has several advantages in comparisons with other methods. In particular, it tests the null hypothesis:

$$H_o : \; d \;=\; d_o \tag{3}$$

in (1) and (2) for any real value d_o. Since this method is parametric we need to specify any functional form for the I(0) error term u_t in (1). Unlike other stationary and nonstationary (unit roots) tests (Schmidt and Phillips, 1992) the limit distribution of this procedure is standard normal and does not change with features of the regressors included in equation (2).

Additionally, it is the most efficient method in the Pitman sense against local departures from the null. A semiparametric local "Whittle" method (Robinson, 1995) will also be implemented in the following section.

It is well-established that long memory models tend to overestimate the degree of persistence of the series in the presence of structural breaks (Cheung and Lai, 1993; Diebold and Inoue, 2001; Ben Nasr et al., 2014 and others). In other words, long-memory can be spurious due to regime changes. Hence, following Bahmani-Oskoee et al. (2014, 2015), we model the dynamics of the current account to GDP ratio by including both sharp shifts and smooth breaks in the estimation of a level and trend equation as follows:

$$y_t = \alpha + \beta t + \sum_{l=1}^{m+1} \theta_l DU_{l,t} + \sum_{l=1}^{m+1} \rho_l DT_{l,t} + \sum_{k=1}^{n} \gamma_{1,k} \sin(\frac{2\pi kt}{T}) + \sum_{k=1}^{n} \gamma_{2,k} \cos(\frac{2\pi kt}{T}) + \varepsilon_t \qquad (4)$$

In equation (4), *t*, *T*, and *m* are time trend, sample size and the optimum number of breaks, respectively. The other regressors are defined as:

$$DU_{k,t} = \begin{cases} 1 & if \quad TB_{k-1} < t < TB_k \\ 0 & otherwise \end{cases} \qquad (5)$$

$$DT_{k,t} = \begin{cases} t - TB_{k-1} & if \quad TB_{k-1} < t < TB_k \\ 0 & otherwise \end{cases} \qquad (6)$$

The terms *DU* and *DT* are entered in the model to capture, respectively, the sharp and smooth shifts.[2] Following the work of Gallant (1981), in order to obtain a global approximation from the smooth transition, we use the Fourier approximation and enter both terms of $\sum_{k=1}^{n} \gamma_{1k} \sin(\frac{2\pi kt}{T})$ and $\sum_{k=1}^{n} \gamma_{2k} \cos(\frac{2\pi kt}{T})$ into the model; where *n* and *k* represent the number of frequencies contained in the approximation ($\mathrm{n} \leq \frac{\mathrm{T}}{2}$) and particular frequency, respectively.

[2]Equation (4) is not only an extension of Enders and Holt (2012) but also a combination of Carrion-i-Silvestre et al. (2006) and Becker et al. (2006) tests.

The estimation of equation (4) involves three issues: the choice of m, the choice of n, and the choice of *k*. As noted by Becker et al. (2004), it is reasonable that we restrict *n* = 1, because if $\gamma_{1,k} = \gamma_{2,k} = 0$ can be rejected for one frequency, then the null hypothesis of time invariance is also rejected. Also Enders and Lee (2012) noted that imposing the restriction *n* = 1 is useful in order to save degrees of freedom and prevent an over-fitting problem. Hence we re-specify the equation (4) as follows:

$$y_t = \alpha + \beta t + \sum_{l=1}^{m+1} \theta_l DU_{l,t} + \sum_{i=1}^{m+1} \rho_i DT_{i,t} + \gamma_1 \sin(\frac{2\pi kt}{T}) + \gamma_2 \cos(\frac{2\pi kt}{T}) + \varepsilon_t \qquad (7)$$

It is important to note that we can remove the impact of possible structural breaks on the current account to GDP ratios based on the information of break dates. We follow the procedure adopted by Tsong and Lee (2011) to reconstruct time series of current account to GDP by taking into account both sharp shifts and smooth breaks as follows:

$$y_t = CA_t - \hat{\alpha} - \hat{\beta} t - \sum_{l=1}^{m+1} \hat{\theta}_l DU_{l,t} - \sum_{i=1}^{m+1} \hat{\rho}_i DT_{i,t} - \hat{\gamma}_1 \sin(\frac{2\pi kt}{T}) - \hat{\gamma}_2 \cos(\frac{2\pi kt}{T}) \qquad (8)$$

where y_t^* is the current account to GDP ratio adjusted by the effect of possible structural breaks (both sharp and smooth), CA_t is the current account to GDP ratio, DU_t and DT_t are the same as in Equation (7). For details about how to estimate Equation (7), interested readers can refer to Bahmani-Oskoee et al. (2014, 2015). Once we have filtered the data for sharp and smooth breaks, we apply our long memory tests to the adjusted series of current account to GDP ratios.[3]

[3]Note that we also considered the case in equation (4) where we only have breaks in the mean. However, the fit of our model was better when we accounted for breaks in both the mean and trend. Complete details on the goodness-of-fit analyses is available upon request from the authors.

Table 1: Basic statistics for the current account to GDP ratio

Country	Sample	Mean	Std. dev.	Min	Max
Brazil	1979Q1-2016Q1	-1.7	2.3	-6.8	5.0
Canada	1961Q1-2016Q1	-1.7	2.0	-5.0	3.6
China	1982Q1-2016Q1	2.5	2.9	-3.3	10.6
Germany	1971Q1-2016Q2	2.2	3.1	-2.6	9.5
France	1973Q1-2016Q1	0.1	1.3	-2.5	4.4
United Kingdom	1955Q1-2016Q1	-1.1	1.8	-7.2	3.8
Italy	1970Q1-2016Q1	-0.3	2.1	-6.1	4.6
Japan	1968Q1-2016Q2	2.1	1.5	-2.1	5.0
India	1980Q1-2016Q1	-1.6	1.6	-6.2	3.1
Russia	1995Q1-2015Q4	5.9	4.4	-3.3	17.9
United States	1960Q1-2016Q1	-1.6	1.9	-6.3	1.2
South Africa	1960Q1-2016Q1	-1.2	3.7	-11.0	13.9

4. Data

The quarterly current account and GDP data are from the OECD Economic Outlook database, except for South Africa, where they are from the South African Reserve Bank. The period covered starts between the mid-1950s and the early 1970s for G7 countries and around 1980 for the BRICS, except South Africa (1960) and Russia (1995). The precise samples are displayed in Table 1, which also provides some basic statistics. The mean current account to GDP ratio is highest in Russia, where the sample covers a relatively short period of generally high oil prices, and lowest in Canada and Brazil. Standard deviations vary from 1.3 in France to 4.4 in Russia. Relatively high standard deviations in Russia and South Africa presumably reflect the importance of commodities, with volatile prices, in the exports of these countries. South Africa experienced the biggest deficit and Russia had the highest surplus. China and South Africa also had high peaks in their current account surpluses. Among G7 countries, the biggest deficits have been seen in the United Kingdom, the United States and Italy, while the highest surplus was recorded in Germany.

5. Empirical results

Table 2 displays the estimates of d in the model given by (1) and (2) under the two cases of uncorrelated and autocorrelated errors (Table 2). In the latter case, we use the exponential model of Bloomfield (1973), which is a non-parametric approach of modelling I(0) errors and produces autocorrelations decaying exponentially, as in the ARMA case. It is non-parametric in the sense that the model is only implicitly defined by its spectral density function and does not present an explicit functional form for u_t in (1). For the two cases we present the estimates for the three standard models with i) no deterministic terms (i.e., $\beta_0 = \beta_1 = 0$ a priori in (2)), ii) an intercept (β_0 unknown and $\beta_1 = 0$ a priori), and iii) with an intercept and a linear time trend (both β_0 and β_1 unknown) and along with the estimates of d we display the 95% confidence bands corresponding to the non-rejection values of d using Robinson's (1994) approach. We marked in bold in the table the most appropriate specification according to the statistical significance of these deterministic terms and we see that the time trend is not required in any single case.[4]

[4] Note that under H_o (3), the model becomes $y_t = \beta_0 + \beta_1 t + x_t$, $(1 - L)^{do} x_t = u_t$, which can be rewritten as $\tilde{y}_t = \beta_0 \tilde{1}_t + \beta_1 \tilde{t}_t + \grave{u}_t$, where $\tilde{y}_t = (1 - L)^{do} y_t$; $\tilde{1}_t = (1 - L)^{do} 1_t$; and $\tilde{t}_t = (1 - L)^{do} t_t$; and noting that u_t is I(0) by construction, the t-tests apply on the estimated coefficients on β_0 and β_1.

Table 2: Estimates of d based on a parametric method with uncorrelated errors

i) Uncorrelated errors			
Country	No regressors	An intercept	A linear time trend
Brazil	0.71 (0.62, 0.82)	**0.69 (0.60, 0.81)**	0.69 (0.61, 0.81)
Canada	0.90 (0.80, 1.03)	**0.89 (0.79, 1.02)**	0.89 (0.79, 1.02)
China	0.89 (0.77, 1.04)	**0.90 (0.79, 1.05)**	0.90 (0.79, 1.05)
Germany	**0.92 (0.84, 1.02)**	0.92 (0.84, 1.02)	0.92 (0.84, 1.02)
France	0.85 (0.73, 1.01)	**1.01 (0.86, 1.18)**	1.01 (0.87, 1.18)
United Kingdom	**0.67 (0.59, 0.78)**	0.67 (0.59, 0.78)	0.67 (0.58, 0.78)
Italy	**0.92 (0.82, 1.05)**	0.93 (0.82, 1.06)	0.93 (0.82, 1.06)
Japan	**1.11 (0.98, 1.26)**	1.11 (0.98, 1.26)	1.11 (0.98, 1.26)
India	0.56 (0.46, 0.70)	**0.55 (0.45, 0.70)**	0.55 (0.45, 0.70)
Russia	1.08 (0.88, 1.34)	**1.14 (0.95, 1.42)**	1.14 (0.95, 1.42)
United States	**1.02 (0.95, 1.12)**	1.02 (0.94, 1.12)	1.02 (0.94, 1.12)
South Africa	**0.73 (0.64, 0.84)**	0.73 (0.64, 0.84)	0.73 (0.64, 0.85)
ii) Autocorrelated (Bloomfield) errors			
Country	No regressors	An intercept	A linear time trend
Brazil	0.76 (0.59, 0.98)	**0.74 (0.59, 0.98)**	0.74 (0.58, 0.98)
Canada	0.78 (0.64, 0.96)	**0.74 (0.60, 0.96)**	0.74 (0.60, 0.96)
China	0.77 (0.57, 1.10)	**0.80 (0.60, 1.12)**	0.81 (0.58, 1.12)
Germany	**1.01 (0.86, 1.24)**	1.01 (0.85, 1.23)	1.01 (0.85, 1.23)
France	0.72 (0.54, 1.02)	**0.95 (0.59, 1.51)**	0.95 (0.61, 1.52)
United Kingdom	**0.66 (0.50, 0.85)**	0.64 (0.51, 0.84)	0.64 (0.47, 0.84)
Italy	**0.89 (0.67, 1.22)**	0.90 (0.68, 1.22)	0.90 (0.67, 1.22)
Japan	**0.90 (0.67, 1.22)**	0.89 (0.63, 1.21)	0.89 (0.64, 1.21)
India	0.49 (0.33, 0.69)	**0.47 (0.32, 0.69)**	0.47 (0.32, 0.69)
Russia	0.72 (0.45, 1.11)	**0.74 (0.44, 1.24)**	0.75 (0.44, 1.24)
United States	**1.01 (0.88, 1.21)**	1.00 (0.87, 1.20)	1.00 (0.87, 1.20)
South Africa	**0.76 (0.53, 1.04)**	0.74 (0.53, 1.06)	0.73 (0.54, 1.06)

Note: Bold entries correspond to the most appropriate specifications according to the deterministic terms.

Starting with the results based on white noise errors (in Table 2i), we observe that mean reversion (i.e., $d < 1$) only takes place for the cases of India ($d = 0.55$), the United Kingdom (0.67), Brazil (0.69) and South Africa (0.73). In all the other countries, though the estimates of d are smaller than 1 in some cases, we cannot reject the null hypothesis of unit roots (i.e., d = 1). Allowing for autocorrelation, the estimated values of d are smaller and mean reversion is now found in the cases of India (0.47), the United Kingdom (0.66), Canada and Brazil (0.74). Thus, for India, the United Kingdom and Brazil, mean reversion (d < 1) is found irrespective of the specification of the error term. For South Africa, it only takes place under no autocorrelation and for Canada only under autocorrelation.

Table 3: Semiparametric estimates of d (Robinson, 1995)

Country	11	12	13	14	15	16	17	18	19	20
Brazil	0.900	0.860	0.926	0.993	1.035	0.894	**0.790**	0.814	0.836	0.865
Canada	0.861	0.823	0.842	0.861	0.808	0.806	0.820	**0.782**	**0.709**	**0.729**
China	**0.745**	0.812	**0.756**	**0.773**	**0.786**	0.837	0.788	0.806	0.809	0.816
Germany	1.031	0.885	0.911	0.949	0.967	0.940	0.967	0.932	0.927	0.959
France	0.841	0.842	0.803	**0.756**	**0.684**	**0.658**	**0.689**	**0.713**	**0.700**	**0.722**
United Kingdom	**0.369**	**0.466**	**0.425**	**0.466**	**0.500**	**0.525**	**0.562**	**0.607**	**0.602**	**0.630**
Italy	**0.643**	**0.531**	**0.528**	**0.600**	**0.631**	**0.637**	**0.665**	**0.699**	**0.759**	**0.761**
Japan	**0.429**	**0.431**	**0.532**	**0.592**	**0.552**	**0.596**	**0.586**	**0.575**	**0.610**	**0.639**
India	**0.697**	0.773	0.810	0.803	**0.562**	**0.596**	**0.631**	**0.575**	**0.599**	**0.555**
Russia	**0.719**	0.789	0.851	0.938	0.940	0.991	0.945	0.973	0.968	1.015
United States	0.885	0.984	1.023	1.117	1.165	1.206	1.288	1.130	1.140	1.150
South Africa	**0.154**	**0.191**	**0.257**	**0.327**	**0.360**	**0.364**	**0.377**	**0.429**	**0.475**	**0.501**
Lower 5%	0.752	0.762	0.771	0.780	0.787	0.794	0.800	0.806	0.811	0.816
Upper 5%	1.247	1.237	1.228	1.219	1.212	1.205	1.199	1.193	1.188	1.183

Note: Bold entries correspond to the cases where evidence of mean reversion (d < 1) at the 5% level of significance is observed.

Table 3 displays the estimates of d using a semiparametric method (Robinson, 1995) that also uses the Whittle function in the frequency domain. It is a "local" method in the sense that it only uses a band of frequencies degenerating to zero. We report the results of d only for a selected group of bandwidth numbers, from m = 11 to 20. This choice clearly reflects a trade-off between bias and variance: the asymptotic variance is decreasing with m, while the bias is growing with m.[5]

We observe in this table that more cases of mean reversion occur using the semiparametric Whittle method. Only for Germany and the United States, is there no single case of mean reversion for the selected bandwidth numbers. Evidence of mean reversion for all bandwidths is obtained in the cases of the United Kingdom, Italy, Japan and South Africa, and the lowest degrees of integration seem to take place for the United Kingdom, Japan and South Africa.

In Table 4 we focus on the filtered data, and report the estimates of d and their corresponding confidence intervals in the same way as in Table 2, that is, using the parametric approach of Robinson (1994) for the two cases of uncorrelated and autocorrelated (Bloomfield) errors.

The first thing we observe in this table is that no deterministic terms are required in this context. Also, the nonstationary I(1) hypothesis (i.e. d = 1) is rejected in all series for the two models presented (uncorrelated and autocorrelated errors). The results are very similar for the two cases of uncorrelated and autocorrelated errors. Diagnostic tests of the residuals suggest that the uncorrelated model is appropriate in all cases.[6] Thus we choose the results in Table 4i). According to these results, we obtain:

a) Evidence of stationary I(0) behaviour: Russia (-0.41), India (0.02), Brazil (0.08) and China (0.11).

[5] The estimates here were obtained based on the first differenced data, then adding 1 to the estimated values of d.

[6] Based on the suggestion of an anonymous referee, we also conducted the analysis by filtering the data only with the smooth breaks. In general our results were relatively weaker (i.e. there was no evidence of stationarity), with mean reversion observed only with long-memory in the cases of Brazil, Canada, United Kingdom, India, Italy and South Africa under the case of non-autocorrelated errors, and for Brazil, Canada, United Kingdom and India under the case of autocorrelated errors. The rest of the cases displayed unit root behaviour. Overall, our results indicate that if we fail to account for sharp breaks, all the series show long-memory – a finding consistent with the literature on structural breaks and spurious long-memory. Complete details of these results are available from the authors upon request.

Table 4: Estimates of d for the sharp and smooth breaks filtered data

i) No autocorrelation			
Countries	No regressors	An intercept	A linear time trend
Brazil	**0.08 (-0.15, 0.36)**	0.08 (-0.16, 0.36)	0.08 (-0.16, 0.36)
Canada	**0.46 (0.27, 0.68)**	0.46 (0.27, 0.68)	0.46 (0.26, 0.69)
China	**0.11 (-0.09, 0.37)**	0.11 (-0.09, 0.37)	0.11 (-0.09, 0.37)
France	**0.31 (0.03, 0.65)**	0.32 (0.03, 0.95)	0.49 (0.03, 0.96)
United Kingdom	**0.26 (0.13, 0.43)**	0.26 (0.12, 0.43)	0.26 (0.12, 0.43)
Germany	**0.17 (0.00, 0.37)**	0.17 (0.00, 0.37)	0.17 (0.00, 0.37)
India	**0.02 (-0.20, 0.31)**	0.02 (-0.20, 0.31)	0.02 (-0.20, 0.32)
Italy	**0.36 (0.21, 0.55)**	0.37 (0.21, 0.58)	0.39 (0.22, 0.59)
Japan	**0.55 (0.34, 0.80)**	0.56 (0.35, 0.81)	0.56 (0.35, 0.81)
Russia	**-0.41 (-0.76, 0.06)**	-0.41 (-0.77, 0.06)	-0.41 (-0.77, 0.06)
South Africa	**0.37 (0.24, 0.54)**	0.37 (0.24, 0.54)	0.37 (0.24, 0.54)
United States	**0.48 (0.32, 0.67)**	0.48 (0.32, 0.68)	0.48 (0.32, 0.68)
i) Autocorrelated errors			
Countries	No regressors	An intercept	A linear time trend
Brazil	**0.08 (-0.07, 0.23)**	0.08 (-0.07, 0.23)	0.08 (-0.07, 0.23)
Canada	**0.46 (0.34, 0.59)**	0.46 (0.34, 0.59)	0.46 (0.34, 0.59)
China	**0.11 (0.00, 0.24)**	0.11 (0.00, 0.24)	0.11 (0.00, 0.24)
France	**0.31 (0.14, 0.49)**	0.32 (0.14, 0.80)	0.46 (0.14, 0.81)
United Kingdom	**0.26 (0.17, 0.35)**	0.26 (0.17, 0.35)	0.26 (0.17, 0.35)
Germany	**0.33 (0.24, 0.42)**	0.33 (0.24, 0.42)	0.33 (0.24, 0.42)
India	**0.02 (-0.11, 0.18)**	0.02 (-0.11, 0.18)	0.02 (-0.11, 0.18)
Italy	**0.37 (0.28, 0.47)**	0.37 (0.28, 0.47)	0.37 (0.28, 0.47)
Japan	**0.55 (0.43, 0.69)**	0.56 (0.43, 0.70)	0.56 (0.44, 0.70)
Russia	**-0.41 (-0.61, 0.18)**	-0.41 (-0.61, 0.18)	-0.41 (-0.62, 0.18)
South Africa	**0.37 (0.29, 0.46)**	0.37 (0.29, 0.46)	0.37 (0.29, 0.46)
United States	**0.48 (0.38, 0.58)**	0.48 (0.38, 0.59)	0.48 (0.38, 0.59)

Note: Bold entries correspond to the most appropriate specifications according to the deterministic terms.

b) Evidence of stationary long memory behaviour ($0 < d < 0.5$): Germany (0.17) and the United Kingdom (0.26).

c) For the remaining cases, though the values of d are in the stationary region (-0.5, 0.5) in all cases except Japan, the intervals include values which are above 0.5, though still statistically significantly smaller than 1. Point estimates for d are: France (0.31), Italy (0.36), South Africa (0.37), Canada (0.46), USA (0.48) and finally Japan (0.55).[7]

To sum up, when smooth and sharp breaks are filtered out, current account to GDP ratios are stationary in all BRICS countries except South Africa, which depicts long-memory, but with mean-reversion. In the G7 economies, imbalances tend to exhibit more persistence. While no evidence of explosive dynamics (unit roots) is found, long memory behaviour seems to be prevalent, even though non-stationarity can be ruled out in all cases. Hence, G7 countries seem to be able to sustain current account imbalances for longer than BRICS countries, except South Africa. This could reflect a looser financial constraint for G7 countries and South Africa, due to deeper financial markets and stronger trust from external creditors. However, somewhat surprisingly, the persistence of current account imbalances is not clearly related to whether countries experience mostly deficits or surpluses. The divide in current account dynamics is mainly between BRICS and G7 countries, whereas within both groups there are both countries mainly running deficits and countries more often in surplus over the estimation period. This suggests that other factors than the external financial constraint are at play, for example exchange rate adjustments or domestic determinants of saving and investment.

We will now provide further evidence of the relevance of our modeling choices. Recall that we have used the Fourier approximation to 'mimic' the time-varying parameter and hence nonlinearity in the series. In Tables 5a and 5b, we present the optimum breaks and frequency from the mean reverting function in equation (7) alongside the estimated *F*-statistic

[7] Based on the suggestion of an anonymous referee, we also estimated a MS-ARFIMA model of Tsay and Härdle (2009). We observed that in 9 out of the 12 countries, non-stationarity could not be rejected in at least one-regime. Italy's current account to GDP ratio was found to be stationary, while that of the United Kingdom and South Africa was found to show mean-reversion with long-memory behavior. Complete details of these results are available from the authors upon request.

that enables us to test for the absence of the nonlinear component in equation (7). In other words the *F*-statistic is computed by comparing the sum of squared residuals (*SSR*) from equation (7) with the nonlinear component (unrestricted model) with the *SSR* from equation (7) without the nonlinear component (restricted model). However, the critical values for the *F*-test are non-standard due to nuisance parameters (Becker et al. 2004). Hence, we follow Bahmani-Oskooee et al. (2014, 2015) and use Monte Carlo simulations to compute the critical values based on 10,000 replications. We fixed k at a maximum of 10 and m at a maximum of 7. From Panel A of Table 5a and 5b, we observe that the optimum frequency varies from one country to the other, with a minimum of 1 and maximum of 10. The computed *F*-statistics are in all cases greater than the critical values, even at the one percent level, except for Canada (greater at the 10 percent level). Hence, the mean reverting function with the nonlinear component is accepted. Turning to the results from panel B of Table 5a and 5b, we observe that there are 3 to 5 breaks in the ratios, thus vindicating the decision to model sharp breaks besides the smooth ones. Some of the sharp breaks relate to major economic events and policy decisions, which translated rapidly in large changes in the current account to GDP ratio, including for Brazil the introduction of a new currency in July 1994 and foreign exchange turbulences preceding the presidential election of October 2002, the German reunification in October 1990, the widening of fluctuation margins within the European exchange rate mechanism in July 1992 which is reflected in a break for France, the decision to let sterling float in June 1972 and the Russian moratorium on foreign debt repayment in August 1998, financial liberalization in India and South Africa in 1991 and 1985 respectively. Other sharp breaks are found in times of high volatility, for instance in different phases of the recent financial crisis, but there are also some cases which are more difficult to relate exactly to specific events.

Table 5a: Estimation results for the Mean Reverting function in Equation (8) for the BRICS Countries

Panel A: The results for optimum frequency and the F-statistic and its critical values

Current Account	Optimum frequency	F stat	90%	95%	97.5%	99%
Brazil	1	17.0377	2.638	3.281	4.047	4.684
Russia	9	10.066	2.488	3.305	4.071	5.111
India	4	31.262	2.421	3.200	3.979	5.424
China	10	47.499	2.241	3.139	3.771	4.977
South Africa	7	15.255	2.439	3.124	3.527	4.593

Panel B: The results for sharp drift dates in Equation (8)

Countries	Break dates				
Brazil	1984:Q3	1994:Q4	2002:Q3	2008:Q1	
Russia	1998Q4	2000:Q4	2006:Q4		
India	1991:Q4	2000:Q4	2008:Q3		
China	1990:Q1	2004:Q3	2009:Q3		
South Africa	1969:Q1	1977:Q2	1985Q3	1994:Q2	2004:Q2

Table 5b: Estimation results for the Mean Reverting function in Equation (8) for the G7 Countries

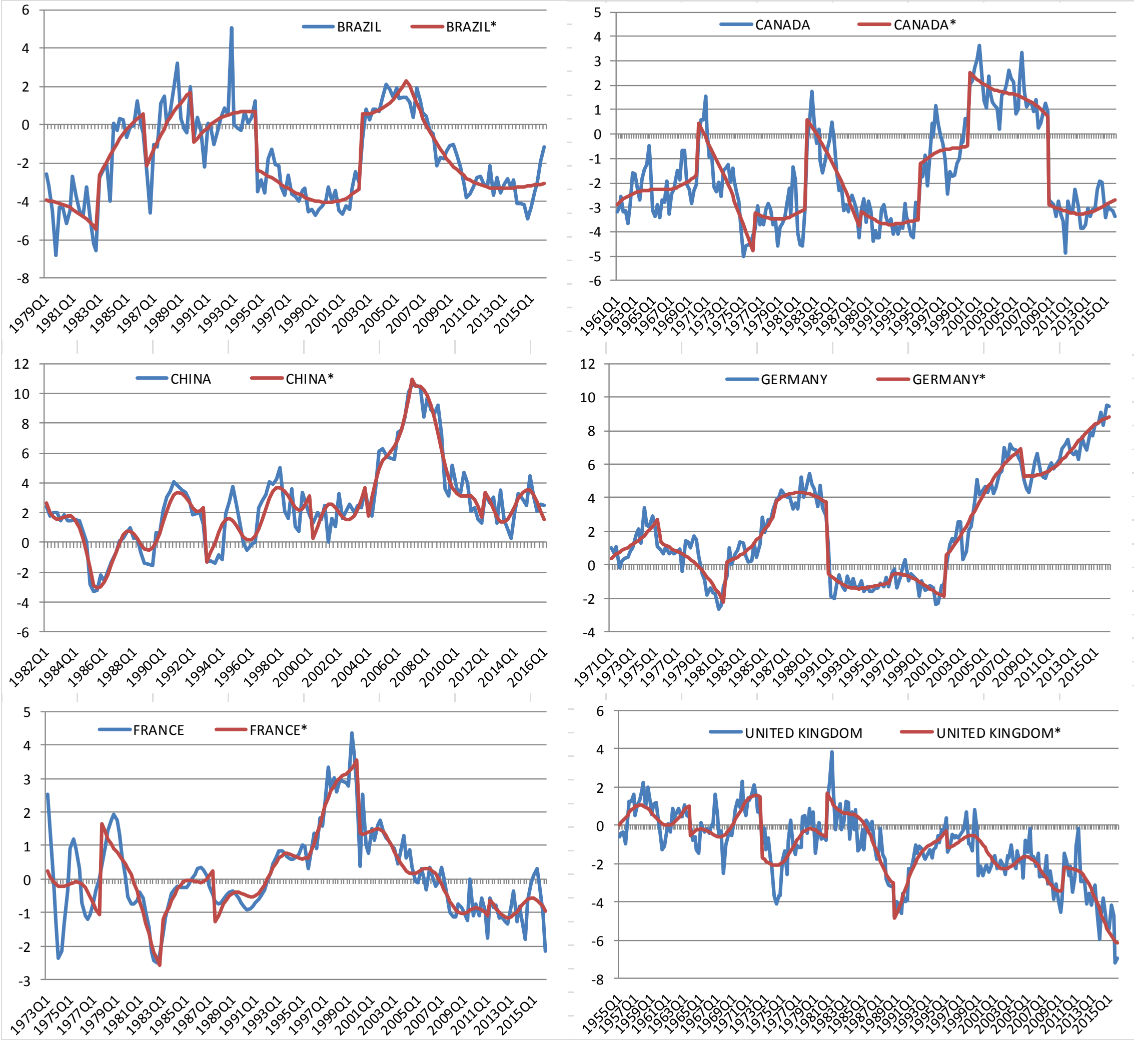
Panel A: The results for optimum frequency and the F-statistic and its critical values

Current Account	Optimum frequency	F stat	90%	95%	97.5%	99%
Canada	5	2.226	2.177	2.730	3.582	4.983
France	10	11.165	2.274	2.883	3.637	4.587
Germany	5	7.163	2.405	3.104	3.708	4.806
Italy	2	45.134	2.453	3.046	3.843	5.531
Japan	2	35.254	2.226	2.857	3.520	4.981
United Kingdom	9	28.487	2.251	3.040	3.593	4.563
United States	10	45.162	2.249	3.154	3.813	4.467

Panel B: The results for sharp drift dates in Equation (8)

Countries	Break dates			
Canada	1986:Q1	1995:Q1	2008:Q1	
France	1980:Q1	1992:Q3	2000:Q3	2007:Q3
Germany	1984:Q2	1991:Q1	2004:Q1	
Italy	1980:Q1	1993:Q2	2000:Q1	
Japan	1983:Q2	2001:Q2	2008:Q3	
United Kingdom	1972:Q3	1987:Q2	2006:Q3	
United States	1968:Q2	1983.Q3	1992:Q2	2008:Q1

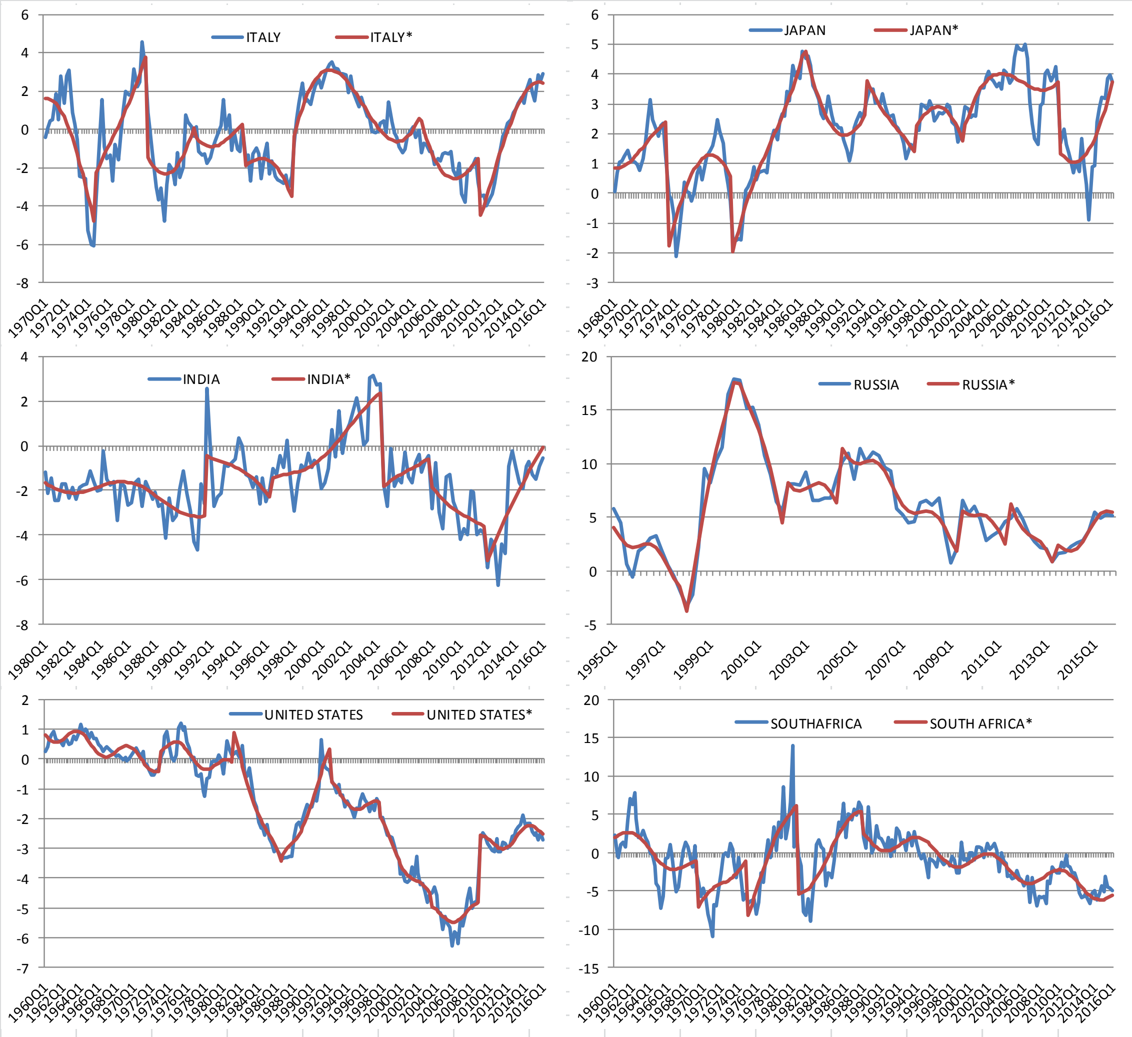
Figure 2: Current account balance (% of GDP), actual and fitted values



Note: Stars indicate fitted series.

Source: OECD and authors' calculations.

Figure 2: Current account balance (% of GDP), actual and fitted values (cont.)



Note: Stars indicate fitted series.

Source: OECD, South African Reserve Bank and authors' calculations.

Finally, we present the time paths of the current account to GDP ratios in Figure 2. The sub-figures show that there are structural shifts in the ratios, and hence points to the need to allow for both sharp shifts and smooth breaks in testing for a unit root and/or stationarity. We superimpose the predicted time paths from our model on the actual time paths, and we observe that the predicted series tracks the dynamic behaviour of the actual current account to GDP series well, suggesting that the decision to include the dummy variables and Fourier approximations is quite reasonable since the data generating processes are indeed nonlinear.

6. Concluding remarks

In this paper, we have shown the relevance of long memory and structural breaks in modeling the dynamics of current account to GDP ratios in BRICS and G7 countries. Specifically, we found that when smooth and sharp breaks are filtered out, current account to GDP ratios are stationary in all BRICS countries, except for South Africa which shows mean-reversion with long-memory, just like the G7 economies. Long memory behavior in South Africa and the G7 suggests that these economies are able to sustain current account imbalances for longer than Brazil, Russia, India and China. This could reflect a looser financial constraint for G7 countries and South Africa, due to deeper financial markets and stronger trust from external creditors. Nevertheless, the persistence of current account imbalances is not clearly related to the deficit or surplus position of the current account, which suggests that other factors than the external financial constraint are at play. The finding that current account to GDP ratios are mean-reverting in BRICS and G7 countries after accounting for structural breaks suggests that current account positions are sustainable. However, the occurrence of large breaks shows that current accounts tend to be characterized by sudden shifts rather than gradual adjustments. As sudden shifts can be disruptive, policymakers should try to address their underlying causes through structural reforms and policy design, as suggested by a number of authors (e.g. Milesi-Ferretti and Razin, 1996a,b; Mann, 2002; De Mello and Padoan, 2010; Chinn et al., 2014).

While we conectrate on in-sample fit of the data in this paper, as part of future research, it would be interesting to compare the forecasting performance of the current account to GDP ratio based on the model developed with smooth and sharp structural breaks relative to the one without it, and also possibly other models of current account sustainability which incorporates its fundamental drivers.

References

Apergis, N., Katrakilidis, K.P. and Tabakis, N.M. (2000), Current account deficit sustainability: The case of Greece, Applied Economics Letters, 7:9, 599-603.

Baharumshah, A.Z., Lau, E. and Fountas, S. (2003), On the sustainability of current account deficits: evidence from four ASEAN countries, Journal of Asian Economics 14 (2003) 465-487.

Bahmani-Oskooee, M., Chang, T. and Wu, T-P. (2014), Revisiting purchasing power parity in African countries: panel stationary test with sharp and smooth breaks, Applied Financial Economics, 24, 1429-1438.

Bahmani-Oskoee, M., Chang, T. and Wu, T-P. (2015), Purchasing Power Parity in Transition Countries: Panel Stationary Test with Smooth and Sharp Breaks, International Journal of Financial Studies, May, 3, 153-161.

Becker, R., Enders, W. and Lee, J. (2004), A general test for time dependence in parameters, Journal of Applied Econometrics, 19, 899-906.

Becker, R., Enders, W. and Lee, J. (2006), A stationarity test in the presence of an unknown number of smooth breaks, Journal of Time Series Analysis, 27, 381-409.

Ben Nasr, A., Ajmi, A.N. and Gupta, R. (2014), Modeling the volatility of the Dow Jones Islamic market world index using a fractionally integrated time varying GARCH (FITVGARCH) model, Applied Financial Economics, 24, 993-1004.

Bloomfield, P. (1973), An exponential model in the spectrum of a scalar time series. Biometrika, 60, 217-226.

Buiter, W.H. (1981), Time preference and international lending and borrowing in an overlapping-generations model, Journal of Political Economy 89, 769-797.

Caner, M., Hansen, B.E. (2001), Threshold autoregression with a unit root, Econometrica 69(6), 1555-1596.

Carrion-I-Silvestre, J. I., Del Barrio, T. and López-Bazo, E. (2004) Evidence on the purchasing power parity in a panel of cities, Applied Economics, 36, 961-6.

Cecen, A., and Xiao, L. (2014), Capital flows and current account dynamics in Turkey: A nonlinear time series analysis. Economic Modelling, 39, 240-246.

Chen, S-W. (2011a), Are current account deficits really sustainable in the G-7 countries? Japan and the World Economy 23, 190-201.

Chen, S-W. (2011b), Current account deficits and sustainability: Evidence from the OECD countries, Economic Modelling 28, 1455-1464.

Chen, S.-W. (2013), Long memory and regime switching properties of current account deficits in the US, Economic Modelling, 35, 78-87.

Chen, S.-W. (2014), Smooth transition, non-linearity and current account sustainability: Evidence from the European countries, Economic Modelling, 38, 541-554.

Chen, S.-W. and Xie, Z. (2015), Testing for current account sustainability under assumptions of smooth break and nonlinearity, International Review of Economics and Finance, 38, 142-156.

Cheung, Y-W and Lai, K.S. (1993), Do gold market returns have long memory? Financial Review, 28 (2), 181-202.

Chinn, M.D., Eichengreen, B. and Ito, H. (2014), A forensic analysis of global imbalances, Oxford Economic Papers 66, 465-490.

Christopoulos, D.K. and León-Ledesma, M. (2010), Current-account sustainability in the US: what do we really know about it? Journal of International Money and Finance 29, 442-459.

Clarida, R.H., Goretti, M. and Taylor, M.P. (2006), Are there thresholds of current account adjustment in the G7? Chapter 5 in: NBER, Current Account Imbalances: Sustainability and Adjustment, 169-203.

Clower, E. and Ito, H. (2012), The Persistence of Current Account Balances and its Determinants: The Implications for Global Rebalancing, ADBI Working Paper, No. 400, Asian Development Bank Institute, Tokyo.

Coakley, J., Kulasi, F. and Smith, R. (1996), Current Account Solvency and the Feldstein–Horioka Puzzle", The Economic Journal 106 (436): 620-627.

Cuestas, J.C. (2013), The Current Account Sustainability of European Transition Economies, Journal of Common Market Studies 51 (2), 232-245.

Cuñado, J., Gil-Alana, L.A. and Pérez de Gracia, F. (2008), New Evidence on US Current Account Sustainability, International Journal of Business and Economics 7(1), 1-21.

Cuñado, J., Gil-Alana, L.A. and Pérez de Gracia, F. (2010), European Current Account Sustainability: New Evidence Based on Unit Roots and Fractional Integration, Eastern Economic Journal 36 (2), 177-187.

Dahlhaus, R. (1989). Efficient parameter estimation for self-similar process. Annals of Statistics 7, 4, 1749-1766.

De Mello, L. and Padoan, P.C. (2010), Are Global Imbalances Sustainable? Post Crisis Scenarios, Review of Economics and Institutions, 2(1), Article 3.

De Mello, L., Padoan, P.C. and Rousová, L. (2012), Are Global Imbalances Sustainable? Shedding Further Light on the Causes of Current Account Reversals, Review of International Economics, 20(3), 489-516.

Diebold, F.X. and Inoue, A. (2001), Long memory and regime switching, Journal of Econometrics, 105, 131-159.

Diebold, F.S. and Rudebusch, G. (1991), On the power of Dickey-Fuller tests against fractional alternatives, Economic Letters, 35, 155-160.

Dulger, F. and Ozdemir, Z.A. (2005), Current account sustainability in seven developed countries. Journal of Economic and Social Research 7, 47-80.

Eichengreen, B. (2011), The G20 and Global Imbalances, Economic Policy Program, German Marshall Fund of the United States (GMF), Washington, DC.

Enders,W. and Holt, M. T. (2012), Sharp breaks or smooth shifts? An investigation of the evolution of primary commodity prices, American Journal of Agricultural Economics, 94, 659-73.

Enders, W. and Lee, J. (2012), A unit root test using a Fourier series to approximate smooth breaks, Oxford Bulletin of Economics and Statistics, 74(4), 574-599.

Gallant, R., (1981), On the basis in flexible functional form and an essentially unbiased form: The flexible Fourier form, Journal of Econometrics, 15, 211-353.

Gnimassoun, B. and Coulibaly, I. (2014), Current account sustainability in Sub-Saharan Africa: Does the exchange rate regime matter? Economic Modelling, 40, 208-226.

Gregory, A. and Hansen, B. (1996) Residual-based tests for cointegration in models with regime shifts, Journal of Econometrics, 70, 99-126.

Hassler, U. and Wolters, J. (1995), Long memory in inflation rates. International evidence, Journal of Business and Economic Statistics 13, 37-45.

Herzer, D. and Nowak-Lehmann, F. (2006), Is there a long-run relationship between exports and imports in Chile? Applied Economics Letters 13, 981-986.

IMF (2014), Are global imbalances at a turning point?, World Economic Outlook, October, International Monetary Fund, Washington, DC.

IMF (2015), Imbalances and Growth, Update of Staff Sustainability Assessments for G-20 Mutual Assessment Process, International Monetary Fund, Washington, DC.

Kasuga, H. (2004), Saving–Investment Correlations in Developing Countries, Economics Letters 83, 371-376.

Kiran, B. (2012), The Sustainability of Turkish External Debt: Evidence from Fractionally Integrated Approach Under Structural Breaks, Economic Research-Ekonomska Istraživanja, 25:1, 18-29.

Lau, E., Baharumshah, A.Z. and Haw, C.T. (2006), Current account: mean-reverting or random walk behavior? Japan and the World Economy 18, 90-107.

Lee, D. and Schmidt, P. (1996), On the power of the KPSS test of stationarity against fractionally integrated alternatives, Journal of Econometrics, 73, 285-302.

Lee, J. and Strazicich, M.C. (2003), Minimum Lagrange Multiplier Unity Root Test with Two Structural Breaks, Review of Economics and Statistics, 85(4), 1082-189.

Lee, J., List, J.A. and Strazicich, M.C. (2006), Non-renewable resource prices: Deterministic or stochastic trends? Journal of Environmental Economics and Management, 51, 354-370.

Lumsdaine, R. and Papell, D. (1997), Multiple trend breaks and the unit root hypothesis, Review of Economics and Statistics, 79, 212-218.

Mann, C.L. (2002), Perspectives on the U.S. Current Account Deficit and Sustainability, Journal of Economic Perspectives, 16(3), 131-152.

Milesi-Ferretti, G.M. and Razin, A. (1996a), Current Account Sustainability. Princeton Studies in International Finance No. 81.

Milesi-Ferretti, G.M. and Razin, A. (1996b), Current account sustainability: selected East Asian and Latin American experiences, NBER Working Paper No. 5791.

Nason, J.M. and Rogers, J.H. (2006), The present-value model of the current account hasbeen rejected: Round up the usual suspects, Journal of International Economics 68, 159-187.

Obstfeld, M. and Rogoff, K. (1995), The intertemporal approach to the current account, Chapter 34 in Handbook of International Economics, 3, 1731-1799, Elsevier.

Onel, G. and Utkulu, U. (2006), Modeling the long-run sustainability of Turkish external debt with structural changes, Economic Modelling 23, 669-682.

Perron, P. (1989).The great crash, the oil price shock, and the unit root hypothesis. Econometrica, 57, 1361-1401.

Perron, P. (1994), Trend, Unit Root Hypothesis and Structural Change in Macroeconomic Time Series, in Rao, B. Bhasakara, ed., Cointegration for Applied Economists, St. Martin's Press.

Perron, P. (1997), Further Evidence on Breaking Trend Functions in Macroeconomic Variables. Journal of Econometrics, 80(2), 355-385.

Raybaudi, M., Sola, M., Spagnolo, F. (2004), Red signal: current account deficits and sustainability. Economics Letters 84, 217-223.

Robinson, P.M. (1994), Efficient tests of nonstationary hypotheses, Journal of the American Statistical Association 89,1420-1437.

Robinson, P.M. (1995), Gaussian semi-parametric estimation of long range dependence, Annals of Statistics 23, 1630-1661.

Sachs, J.D. (1981), The current account and macroeconomic adjustment in the 1970s, Brooking Papers on Economic Activity 1, 201-268.

Schmidt, P. and Phillips, P.C. (1992), LM tests for a unit root in the presence of deterministic trends, Oxford Bulletin of Economics and Statistics 54, 3, 257-287.

Taylor, A.M. (2002), A century of current account dynamics, Journal of International Money and Finance 21, 725-748.

Topalli, N. and Dogan, I. (2016), The structure and sustainability of current account deficit: Turkish evidence from regime switching, The Journal of International Trade
and Economic Development, 25:4, 570-589.

Trehan, B. and Walsh, C. (1991), Testing intertemporal budget constraints: theory and application to US Federal Budget deficits and current account deficits. Journal of Money, Credit and Banking 23, 206-223.

Tsay, W.-J. and Härdle, W.K. (2009), A generalized ARFIMA process with Markov-switching fractional differencing parameter, Journal of Statistical Computation and Simulation 79, 731-745.

Tsong, C.C. and Lee, C.F. (2011), Asymmetric inflation dynamics evidence from quantile regression analysis. Journal of Macroeconomics, 33, 668-680.

Wickens, M.R. and Uctum, M. (1993), The sustainability of current account deficits: A test of the US intertemporal budget constraint. Journal of Economic Dynamics and Control, 17, 423-441.

Wu, J.-L. (2000), Mean reversion of the current account: evidence from the panel data unit-root test, Economics Letters 66, 215-222.

Wu, J.-L., Chen, S.-L. and Lee, H.-Y. (2001), Are current account deficits sustainable? Evidence from panel cointegration, Economics Letters 72, 219-224.

Zivot, E. and Andrews, D.W.K. (1992), Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis, Journal of Business and Economic Statistics, 10(10), 251-270.