CURRENCY UNION IN THE EAST AFRICAN COMMUNITY: A FRACTIONAL INTEGRATION APROACH

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ABSTRACT

This study examines the inflation rates from the five countries that belong to the East African Community and which recently signed a protocol outlining their plans for launching a monetary union within ten years. The aim is to examine the persistence in the inflation levels. As it is argued by the literature regarding monetary unions, countries hoping to form a union should present similar inflation patterns. Our study shows that that the countries present non-mean reversion, confirming that shocks will not recover in the long run. Moreover, fractional cointegration relationships are found between all countries with the exception of Tanzania.

JEL Classification: C22

Keywords: East Africa Union; monetary union; fractional integration

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1. Introduction

This study examines the time series behavior of inflation rate levels in the five countries (Burundi, Kenya, Rwanda, Tanzania and Uganda) that belong to the East African Community (EAC) and which recently signed a protocol outlining their plans for launching a monetary union within ten years. We conduct the study using a long memory modeling framework, based on fractional integration, with the aim of analyzing the level of persistence in the inflation levels of these eastern African countries. Alternative measures of persistence based on autoregressive models will also be employed. As is argued by most of the literature regarding monetary unions, countries hoping to join in such a union should present, among other features, similar inflation patterns. Our aim is to determine if this is indeed the case in the East African Community.

Our study shows that the five countries examined present non-mean reverting behavior in their inflation rates, meaning that shocks affecting these countries could be very negative, as they will not return to their original long term projections unless policy measures are implemented. It could therefore be suggested that these countries could embark into a monetary union with the aim of diminishing this risk. Additionally, evidence of fractional cointegration is found in the bilateral relationships between all the countries that form the EAC with the exception of Tanzania, giving thus further support to monetary union in the area.

The rest of the paper is organized as follows: Section 2 briefly describes the history of the East African countries. Section 3 deals with the literature review. Section 4 presents the data and the methodology used in the paper. Section 5 is devoted to the empirical results, while Section 6 concludes the paper.

2. A bit of history

The East African Community (EAC) is an intergovernmental organization comprising five countries in the African Great Lakes region in eastern Africa: Burundi, Kenya, Rwanda, Tanzania and Uganda. The organization was originally founded in 1967, but collapsed in 1977 and was officially revived on 7 July 2000. In 2008, after negotiations with the Southern Africa Development Community (SADC) and the Common Market for Eastern and Southern Africa (COMESA), the EAC agreed to an expanded free trade area including the member states of all three, thus becoming an integral part of the African Community. The East African Community is a potential precursor to the establishment of an East African Federation. In 2010 the EAC launched its own common market for goods, labor and capital within the region, with the goal of creating a common currency union and eventually a full political federation. In November 2013 a protocol was signed outlining the plans of the five member countries to launch a monetary union within ten years.

Kenya, Tanzania and Uganda have had a history of cooperation dating back to the early 20th century. The customs union between Kenya and Uganda in 1917, which the then Tanganyika joined in 1927, was followed by the East African High Commission from 1948 to 1961, later by the East African Common Services Organization from 1961 to 1967 and then by the East African Community until 1977. Inter-territorial cooperation between the Kenya Colony, the Uganda Protectorate and the Tanganyika Territory was first formalized in 1948 by the East African High Commission. This provided a customs union, a common external tariff, currency and postage; and also dealt with common services in transport and communications, research and education. Following independence, these integrated activities were reconstituted and the High Commission was replaced by the East African Common

Services Organization, which many observers thought would lead to a political federation between the three territories. The new organization ran into difficulties because of the lack of joint planning and fiscal policy, separate political policies and the dominant economic position of Kenya. In 1967 the East African Common Services Organization was superseded by the East African Community. This body aimed to strengthen ties between members through a common market, a common customs tariff and a range of public services so as to achieve balanced economic growth within the region.

Kenya, Tanzania and Uganda signed the Treaty for the establishment of the East African Community (EAC) in 1999, which entered into force in July 2000. In 2007 the Treaty was signed by Burundi and Rwanda, expanding the EAC to five countries. According to the Treaty, the EAC should first form a customs union, then a common market and a monetary union, and finally a political union. The Customs Union became operational in 2005, and was formally completed in 2010. The Common Market Protocol was signed in 2009, and the plan is that the creation of a common market, which includes free movement of goods, labor, persons, services and capital, and the right of residence and establishment, will be completed by 2015.

The work of forming a monetary union started early, but proceeded slowly. Thus, in 2007 the EAC member countries decided to fast track its establishment and aimed for 2012. The intention was to sign a protocol to establish the East African Monetary Union (EAMU) in 2012, which was finally signed in 2013, while actual implementation, though planned to be completed by 2015 is now expected to take several years. As evident from the experience of the European Monetary Union (EMU), forming a monetary union is a complicated project, and there is a non-negligible risk of failure. It is therefore necessary to ensure that the pre-conditions for forming the EAMU are adequate. This entails making sure that economic, political and institutional requirements are in place, since benefits are likely to be less visible than short-run costs.

3. Literature review

The Optimum Currency Area (OCA) theory is used to analyze the suitability of a monetary union for a given region; it explores the criteria as well as the costs and benefits of forming a common currency area. The concept of currency areas was founded by Mundell (1961) in his seminal paper titled “A Theory of Optimum Currency Areas”, followed by Mckinnon (1963), Kenen (1969). These authors are the founders of the traditional Optimum Currency Area Theory (OCA), which describes the characteristics that potential monetary union members should possess before they form a single common currency and surrender their national monetary policy and exchange-rate adjustment of their national currencies.

When it comes to African monetary unions, most of the literature is related to the current aim of building up a new currency known as the ECO. This currency union of anglophone West African countries could eventually come to exist in the near future under the name of the West African Monetary Zone (WAMZ). By using fractional integration, it has been established that some significant differences exist between these countries. It has been shown, for instance, that shocks to inflation in Sierra Leone are not mean-reverting, while the results for Gambia, Ghana, and Guinea suggest some inflation persistence (Alagidede, Coleman, and Cuestas, 2010). Balogun (2007) proved that independent monetary and exchange rate policies have been relatively ineffective in influencing domestic activities (especially, GDP and inflation) and, therefore, a currency union could benefit the region. After performing a fractional integration analysis on the West African Economic and Monetary Union (WAEMU), Gil-Alana and Carcel (2014) argued that the eight countries that share the CFA as a common currency are tied together not because of their shared economic homogeneity but rather due to their strong historical and traditional ties to France.

Several papers have been published with the purpose of promoting the East African Community. Using the Optimal Currency Area (OCA) approach, Mafusire and Brixiova (2013) test empirically the extent of the shock synchronization among the EAC members, concluding that if the countries in the union have major structural differences, a common monetary policy would have differential impacts that may not be helpful to some members. Durevall (2011) pointed out that the EAC has a number of convergence criteria, but these need to be improved and revised so that the union succeeds, and Kishor and Ssozi (2009) found that the proportion of shocks which are common across different countries is small, implying weak synchronization, with this degree of synchronization having improved after the signing of the EAC treaty in 1999. Several authors have studied the viability of a monetary union in the EAC. These studies have used different models and have reached different conclusions. For example, Buigut and Valev (2005) applied a two-variable SVAR model to test for shock correlation in the EAC countries; they found that forming a monetary union in the EAC is not feasible. Conversely, Mkenda (2001) and Falagiarda (2010) employed the G-PPP approach which uses cointegration analysis concluding that a monetary union in East Africa could be a viable option. Lastly, Sheikhet al. (2011) and Opolot and Osoro (2004) studied the feasibility of forming a monetary union in the EAC using the business cycle synchronization approach of Hodrick-Prescott and Baxter-King filters. They found a low degree of synchronization between EAC members, but with improved results in recent years.

Some studies have been conducted on macroeconomic issues in African countries using long memory techniques. For instance, focusing on the exchange rates, Mokoena, Gupta, and Van Eyden (2009) showed that, except for South Africa, none of the SADC (Southern African Development Community) real exchange rates are fractionally integrated. Fractional integration has also been used to analyze the stock market structure (Anouro and Gil-Alana, 2010; Rambaccussing, 2010), inflation (Gil-Alana and Barros, 2013), and housing prices (Gil-Alana, Gupta and Aye, 2013) in African countries. However none of these studies have focused on the East African Community. With this work we attempt to contribute to the literature on the formation of the Monetary Union in the East African Community by conducting a long memory fractional integration analysis of the persistence of the inflation rates in five East African countries.

4. Data and methodology

We use monthly data of the inflation rate from January 2004 to December 2013 from the five member countries of the East African Community (Burundi, Kenya, Rwanda, Tanzania and Uganda), obtained from Trading Economics website. With these time series we conducted a long memory and fractional integration analysis. As we can see in Figure 1, the five graphs show a considerable volatility in the inflation rates, each of them presenting a relatively different form, but with constant ups and down throughout the analysed period.

[Insert Figure 1 about here]

Long memory is a feature of the data which tells us that observations that are far from one another in time are highly correlated. Based on the frequency domain, we can provide the following definition of long memory. Let us suppose that $\{x_t, t = 0, \pm1, \dots\}$

is a covariance stationary process with autocovariance function $E(x_t - Ex_t)(x_{t-j} - Ex_t) = \gamma_j$, such that it has an absolutely continuous spectral distribution function, with a spectral density function, denoted by $f(\lambda)$, and defined as

$$f(\lambda) = \frac{1}{2\pi} \sum_{j'-\infty}^{j=\infty} \gamma_j \cos \lambda j, \qquad -\pi < \lambda \leq \pi. \qquad (1)$$

Then, x_t displays the property of long memory if the spectral density function has a pole at some frequency λ in the interval $[0, \pi)$, i.e.,

$$f(\lambda) \rightarrow \infty, \quad \text{as } \lambda \rightarrow \lambda^*, \quad \lambda^* \in [0, \pi). \qquad (2)$$

The empirical literature has mainly focused on the case where the singularity or pole in the spectrum occurs at the 0 frequency, i.e., ($\lambda^* = 0$). This is the standard case of I(d) models of the form:

$$(1 - L)^d x_t = u_t, \quad t = 0, \pm 1, \ldots, \qquad (3)$$

where d can be any real value, L is the lag-operator ($Lx_t = x_{t-1}$) and u_t is I(0), defined for our purposes as a covariance stationary process with a spectral density function that is positive and finite at the zero frequency. The polynomial $(1-L)^d$ in equation (3) can be expressed in terms of its binomial expansion, such that, for all real d,

$$(1-L)^d = \sum_{j=0}^{\infty} \psi_j L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - dL + \frac{d(d-1)}{2} L^2 - \ldots, \qquad (4)$$

and thus

$$(1-L)^d x_t = x_t - dx_{t-1} + \frac{d(d-1)}{2} x_{t-2} - \ldots,$$

implying that equation (3) can be expressed as

$$x_t = dx_{t-1} - \frac{d(d-1)}{2} x_{t-2} + \ldots + u_{t.},$$

In this context, d plays a crucial role since it indicates the degree of dependence of the time series: the higher the value of d is, the higher the level of association will be between the observations (Barros et al. 2011). The above process also admits an infinite Moving Average (MA) representation such that

$$x_t = \sum_{k=0}^{\infty} a_k u_{t-k},$$

Where

$$a_k = \frac{\Gamma(k+d)}{\Gamma(k+1)\Gamma(d)},$$

and $\Gamma(x)$ represents the Gamma function.

Given the parameterization in (3) we can distinguish several cases of interest depending on the value of d. Thus, if d = 0, $x_t = u_t$, x_t is said to be "short memory" or I(0), and if the observations are autocorrelated (i.e. AR), then they are "weak" in form, in the sense that the values in the autocorrelations are decaying exponentially; if d > 0, x_t is said to be "long memory", so named because of the strong association between observations which are distant in time. Here, if d belongs to the interval (0, 0.5) x_t is still covariance stationary, while d ≥ 0.5 implies nonstationarity. Finally, if d < 1, the series is mean reverting in the sense that the effect of the shocks disappears in the long run, contrary to what happens if d ≥ 1 with shocks persisting forever.

Two univariate methods of estimation of the fractional differencing parameter are employed: one is a Whittle parametric approach in the frequency domain (Dahlhaus, 1989), while the other is a semiparametric "local" Whittle approach (Robinson, 1995). In addition, a simple AR(1) model is also considered as an alternative approach to measure persistence by means of the AR coefficient.

In a multivariate framework, the bivariate relationships between the variables will be examined by means of fractional cointegration; two methods are employed here, the one based on Engle and Granger's (1987) methodology, and extended to the fractional case by Gil-Alana (2003), and a Hausman test for the null of no cointegration against the alternative of fractional cointegration, as developed by Marinucci and Robinson (2001).

5. Empirical results

The first thing we do in the paper is to check the persistence of the series by means of conducting a simple AR(1) model for each case. The estimated AR coefficients, displayed in Table 1 are in all cases extremely close to 1, ranging from 0.958 in the case of Kenya to 0.993 for Tanzania. This clearly indicates that the series are highly persistent and possibly nonstationary. Thus, we also conduct various unit root procedures. In particular, we use the ADF (Dickey and Fuller, 1979) and Phillips and Perron (PP, 1988) approaches. These are some of the most commonly employed tests in the literature.

[Insert Tables 1 and 2 about here]

As expected, the p-values reveal, in Table 2, that we cannot reject the null hypothesis of a unit root in any of the series. Therefore, according to these results, we should take first differences in order to make them I(0) stationary. Nevertheless, we should take into account that these unit root tests might have very low power when directed against specific alternatives such as trend-stationary models (DeJong et al., 1992), structural breaks (Campbell and Perron, 1991), regime-switching (Nelson et al., 2001), or fractionally integration (Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996). In what follows, we focus on the latter type of alternatives, noting that fractional integration includes the classic unit root models as particular cases of interest.

We estimate the fractional differencing parameter d in a model given by the following form:

$$y_t = \alpha + \beta t + x_t; \quad (1 - L)^d x_t = u_t, \quad t = 1, 2, ..., \tag{5}$$

where y_t is the observed time series for each country, α and β are the unknown coefficients corresponding to an intercept and a linear trend, and the resulting errors, x_t, are supposed to be I(d). First we employ a parametric approach, and thus, we need to specify a functional form for the d-differenced process. We considered three different cases corresponding to white noise disturbances, Bloomfield-type and monthly seasonal AR. The model of Bloomfield (1973) is a nonparametric approach that produces autocorrelations decaying exponentially as in the AR(MA) case. Finally, we use monthly AR(1) disturbances based on the monthly nature of the series examined.

In Table 3 we display the estimated values of d under three different specifications assuming a): no deterministic terms (i.e., α = β = in equation (5)), b): an intercept (α unknown and β = 0), and c): an intercept with a linear time trend (α and β unknown).

The most noticeable feature observed in Table 3 is that evidence of mean reversion (i.e. d < 1) is not found in any single case since all the confidence intervals include values of d equal to or higher than 1. For white noise and monthly AR(1) u_t, the unit root null cannot be rejected for the case of Burundi; however, this hypothesis is decisively rejected in favor of d > 1 in the remaining four series. If u_t follows the nonparametric exponential model of Bloomfield (1973) this explosive behavior is attached in the case of Tanzania but the unit root null cannot be rejected in any of the other four series.

[Insert Tables 3 and 4 about here]

In Table 4 we use a semiparametric method proposed by Robinson (1995) and modified later by Abadir et al. (2007) among many others. This is a "local" Whittle estimator in the frequency domain, using a band of frequencies that degenerates to zero.

The results here are consistent with the parametric ones and evidence of unit roots or values of d above 1 are obtained in all cases. This indicates that the series are not mean reverting, implying that if there is a negative shock, the series will not recover by themselves in any single country to their original long term projections, and strong policy measures should be implemented to recover the original values. In addition to this, the fact of having similar results across countries gives further support to the idea that a currency union could be beneficial for these countries.

Next we focus on the bivariate relationships bewteen the countries and the first thing we do is to employ the Engle and Granger's (1987) methodology to test for cointegration. We use this methodology rather than the system-based one of Johansen's (1988, 1991) since later we will extend Engle and Granger's method to the fractional case. The results using this approach are displayed in Table 5.

[Insert Table 5 about here]

Using the above approach, we only found evidence of no cointegration in the case of Tanzania with respect to the other four countries. In the rest of the cases, with the p-values above 0.05 indicating that we reject the null of no cointegration against the alternative of cointegration**.** However, as earlier mentioned with respect to the unit root procedures, this method can be seriously biased due to the fact that no fractional alternatives have been taken into account.

In what follows we use the methodology developed by Gil-Alana (2003), testing the null of no cointegration against the alternative of fractional cointegration. However, in a bivariate context, a necessary condition for cointegration is that the two individual series must be cointegrated of the same order. Thus, as a preliminary screening, we test the homogeneity condition of the order of integration in the bivariate systems (i.e., Ho: $d_x = d_y$), where d_x and d_y are the orders of integration of the two individual series, by using an adaptation of Robinson and Yajima (2002) statistic to the Whittle semiparametric estimation method. The results, though not reported, indicated that all series display similar integration orders.

Two different approaches are employed here to test for the possibility of fractional cointegration. First, following Gil-Alana (2003) we test the null of no cointegration against the alternative of fractional cointegration in the estimated residuals from the regression of one country against another. Here, we consider the two cases of uncorrelated (white noise) errors (in Table 6) and correlated (Bloomfield) disturbances (in Table 7). Then, we also consider the Hausman test for no cointegration of Marinucci and Robinson (2001).

[Insert Tables 6 and 7 about here]

Starting with Gil-Alana (2003) and assuming white noise disturbances (Table 6) we observe that the estimated degree of differentiation is within the unit root interval or even in some cases above 1, finding thus no evidence of cointegration of any degree. However, allowing for autocorrelated disturbances (Table 7) we observe a reduction in the degree of integration of the series and, though the unit root null hypothesis cannot be rejected, we observe many values below 1 suggesting some degree of fractional cointegration.

Next we perform the Hausman test of Marinucci and Robinson (2001). This method is based on the following:

$$H_{is} = 8\,s\left(\hat{d}^{*} - \hat{d}_{i}\right)^{2} \to_{d} \chi_{1}^{2} \qquad \frac{1}{s} + \frac{s}{T} \to 0, \tag{6}$$

where i = x, y refers to each of the series under examination, s is the bandwidth number (in our case we choose $s = (T)^{0.5}$), $\hat{d}_i$ are the univariate estimates of the parent series, and $\hat{d}^*$ is a restricted estimate obtained in the bivariate representation under the assumption that $d_x = d_y$. The results using this approach are displayed in Table 8.

[Insert Table 8 about here]

We observe that for some of the bivariate relationships we find some evidence of cointegration. In particular, for the cases of Burundi-Kenya, Burundi-Rwanda, Burundi-Uganda, Kenya-Rwanda and Kenya-Uganda. These results, together with the non-cointegration results obtained for the case of Tanzania with the rest of countries obtained with the Engle-Granger test show us that there has been traditionally more cointegration in inflation levels between Burundi, Kenya, Rwanda and Uganda than with Tanzania. The lack of cointegration for Tanzania can be explained by the fact that for many years this country has kept stronger economic and trading ties with the Southern African Development Community (SADC), an inter-governmental organization to which Burundi, Kenya, Rwanda and Uganda do not belong.

6. Conclusion

We conducted a long memory and fractional integration analysis on the inflation rate levels of the five member countries of the East African Community: Burundi, Kenya, Rwanda, Tanzania and Uganda. We did so in order to examine in this countries share similar characteristics for a future economic integration.

We first tested persistence by considering a simple AR(1) process and the AR coefficient was found to be very close to 1 in the five series examined. Due to this high degree of persistence, we also performed unit root tests and the results based on the ADF (1979) and the PP (1988) methods support the nonstationary I(1) specification again in the five countries examined. This result however could be questioned due to the low power of the unit root procedures if the series are fractionally integrated. Using this latter approach, the series displays orders of integration above 1, this being a clear sign that inflationary shocks in these countries will take permanent effects. Moreover, looking at bivariate fractionally cointegration relationships among the countries, the results support this hypothesis in all countries except Tanzania. This can be sustained by the fact that for many years Tanzania has kept stronger economic and trading ties with the Southern African Development Community (SADC), an inter-governmental organization to which Burundi, Kenya, Rwanda and Uganda do not belong. In addition to this, during the last decade Tanzania received immense foreign direct investment in its mining industries, mainly gold among other minerals, whereas in the other four countries agriculture remained the main economic sector. We believe it could be essential to take these features into account as monetary unions usually require homogeneity in the inflation levels of its country members. In addition to this, it could also be argued that joining into a monetary union would provide more macroeconomic and financial stability to the region, thus leading to possible lower levels of inflationary persistence in the future.

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Figure 1: Original time series

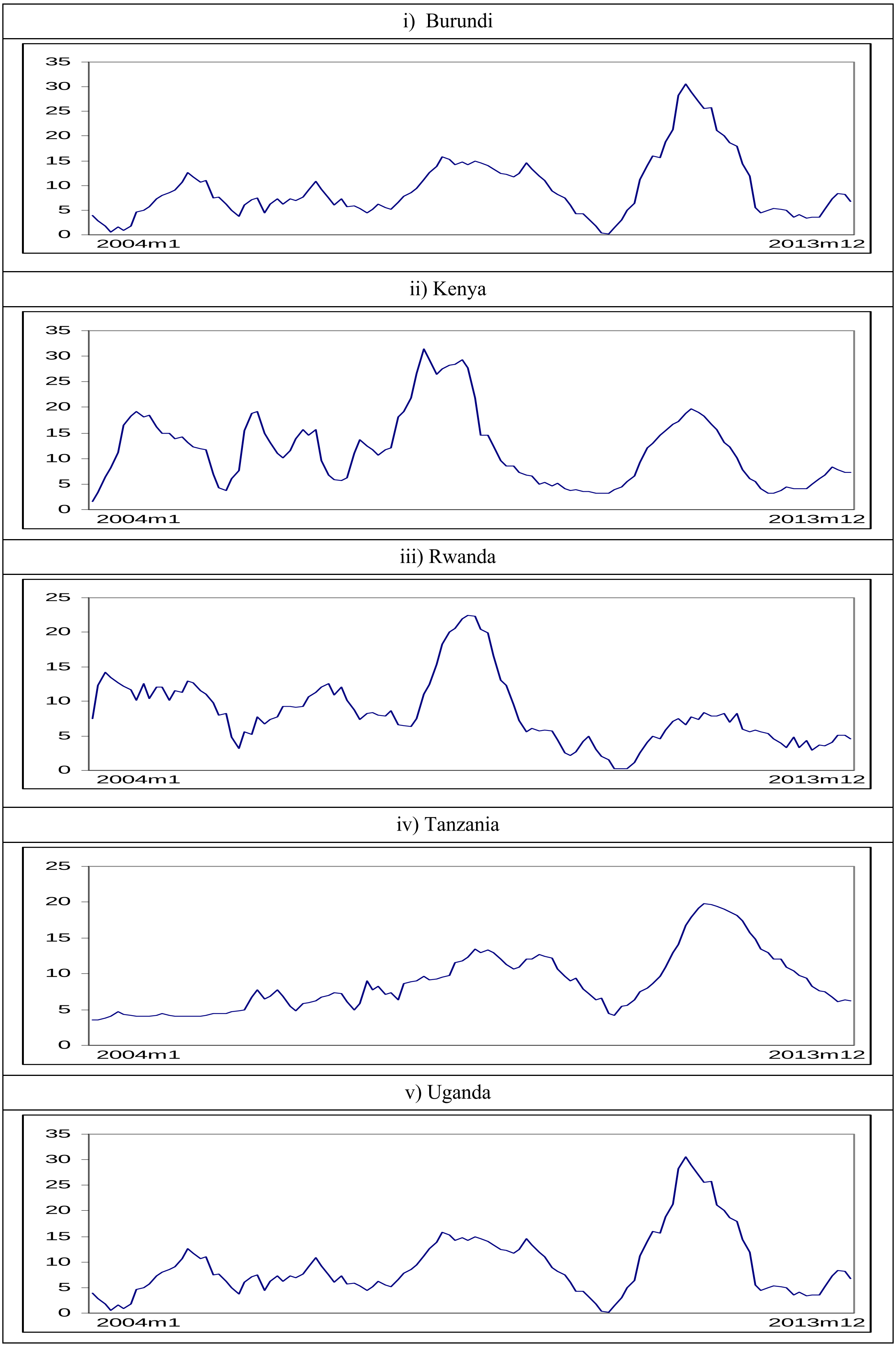

Table 1: Estimated AR(1) coefficients

Estimated AR (1)	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
	0.9747	0.9588	0.9774	0.9933	0.9705

Table 2: Unit root test results

ADF	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
No regressors	0.3922	0.3337	0.06076	0.391	0.4692
Intercept	0.117	0.2469	0.2986	0.4051	0.5100
Time trend	0.3593	0.4745	0.5007	0.8608	0.3685
PP	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
No regressors	0.1559	0.1997	0.2049	0.4111	0.1654
Intercept	0.0981	0.0802	0.1847	0.2921	0.1096
Time trend	0.2817	0.1575	0.2090	0.6886	0.3212

The values in the tables refer to the corresponding p-values.

Table 3: Estimates of d and 95% confidence intervals

	No regressors	An intercept	A linear trend
i) White noise disturbances			
Burundi	**0.98 (0.87, 1.12)**	**1.02 (0.91, 1.18)**	**1.02 (0.91, 1.18)**
Kenya	1.44 (1.28, 1.65)	1.44 (1.28, 1.66)	1.44 (1.27, 1.66)
Rwanda	1.24 (1.12, 1.39)	1.20 (1.09, 1.35)	1.20 (1.09, 1.35)
Tanzania	1.24 (1.13, 1.38)	1.27 (1.16, 1.40)	1.27 (1.16, 1.40)
Uganda	1.33 (1.21, 1.48)	1.37 (1.24, 1.53)	1.37 (1.24, 1.53º)
ii) Bloomfield disturbances			
Burundi	**0.97 (0.69, 1.32)**	**0.96 (0.68, 1.32)**	**0.96 (0.67, 1.33)**
Kenya	**1.04 (0.79, 1.42)**	**1.03 (0.68, 1.41)**	**1.03 (0.72, 1.39)**
Rwanda	**1.27 (0.96, 1.69)**	**1.32 (0.98, 1.74)**	**1.33 (0.98, 1.77)**
Tanzania	1.31 (1.03, 1.62)	1.34 (1.11, 1.66)	1.34 (1.11, 1.66)
Uganda	**1.26 (0.97, 1.58)**	**1.26 (0.97, 1.60)**	**1.26 (0.97, 1.61)**
iii) Monthly AR(1) disturbances			
Burundi	**0.97 (0.88, 1.10)**	**1.03 (0.93, 1.17)**	**1.03 (0.93, 1.17)**
Kenya	1.42 (1.25, 1.65)	1.42 (1.25, 1.65)	1.40 (1.24, 1.64)
Rwanda	1.25 (1.13, 1.42)	1.19 (1.07, 1.36)	1.19 (1.07, 1.36)
Tanzania	1.19 (1.09, 1.34)	1.24 (1.14, 1.38)	1.24 (1.14, 1.37)
Uganda	1.26 (1.16, 1.40)	1.31 (1.20, 1.46)	1.32 (1.20, 1.46)

In bold, evidence of unit roots (d = 1) at the 5% level.

Table 4: Estimates of d based on the Whittle semiparametric method

	10	11	12	15
Burundi	**1.095**	1.311	1.288	**0.939**
Kenya	1.358	1,457	1.342	**0.902**
Rwanda	**0.888**	**0.933**	**1.059**	**1.110**
Tanzania	**1.209**	1.335	1.476	1.253
Uganda	1.408	1.500	1.500	1.210
95% Lower Int.	0.739	0.752	0.765	0.787
95% Upper Int.	1.260	1.247	1.237	1.212

In bold, evidence of unit roots (d = 1) at the 5% level.

Table 5: Engle and Granger's (1987) cointegration results

	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
BURUNDI	---	---	---	---	---
KENYA	0.11860	---	---	---	---
RWANDA	0.23370	0.08135	---	---	---
TANZANIA	0.04972*	0.04570*	0.03636*	---	---
UGANDA	0.07771	0.15270	0.14660	0.00819*	---

*: Evidence of No cointegration at the 5% level.

Table 6: Gil-Alana's (2003) method of fractional cointegration with uncorrelated errors

	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
BURUNDI	---	---	---	---	---
KENYA	(0.89, 1.19)	---	---	---	---
RWANDA	(0.80, 1.13)	(1.11, 1.52)			---
TANZANIA	(0.86, 1.15)	(1.25, 1.63)	(1.08, 1.36)		---
UGANDA	(0.85, 1.18)	(1.19, 1.63)	(1.04, 1.32)	(0.94, 1.21)	---

Table 7: Gil-Alana's (2003) method of fractional cointegration with correlated errors

	BURUNDI	KENYA	RWANDA	TANZANIA	UGANDA
BURUNDI	---	---	---	---	---
KENYA	(0.37, 1.11)	---	---	---	---
RWANDA	(0.44, 1.08)	(0.44, 1.32)			---
TANZANIA	(0.61, 1.30)	(0.72, 1.40)	(0.97, 1.76)		---
UGANDA	(0.50, 1.16)	(0.57, 1.24)	(0.89, 1.69)	(0.75, 1.37)	---

Table 8: Haussman test of no cointegration

	BURUNDI	KENYA	RWANDA	TANZANIA
KENYA	H_{as} =15.093* H_{as} = 4.319* **d^* = 0.682**	---		
RWANDA	H_{as} =12.323* H_{as} = 10.071* **d^* = 0.722**	H_{as} = 0.236 H_{as} = 3.928* **d^* = 0.855**	---	
TANZANIA	H_{as} = 2.683 H_{as} = 1.209 d^* = 0.925	H_{as} = 1.220 H_{as} = 3.130 d^* = 1.026	H_{as} = 3.769 H_{as} = 1.079 d^* = 1.322	---
UGANDA	H_{as} = 8.152* H_{as} = 15.458* **d^* = 0.792**	H_{as} = 0.089 H_{as} = 10.131* **d^* = 0.877**	H_{as} = 2.871 H_{as} = 0.078 d^* = 1.246	H_{as} = 3.130 H_{as} = 3.163 d^* = 1.021

*: We reject the null of no cointegration in favour of fractional cointegration at the 5% level.