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LONG MEMORY IN ANGOLAN MACROECONOMIC SERIES: MEAN REVERSION VERSUS EXPLOSIVE BEHAVIOUR

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Abstract

This study examines the time series behaviour of several Angolan macroeconomic variables, using monthly data from August 1996 to June 2011. The series are the inflation rate, M1, M2, the exchange rate at the beginning and the end of the period, and the monthly average exchange rate. In the first stage univariate fractional integration models are estimated in order to determine whether shocks to the variables have transitory or permanent effects. In the second stage fractional cointegration techniques are applied to test for the existence of long-run equilibrium relationships between the variables of interest. The results suggest a high degree of persistence in the individual series (that are not mean-reverting) and the existence of bivariate long-run cointegrating relationships between prices and money, and prices and nominal exchange rates.

Keywords: Macroeconomics, Angola, fractional integration and cointegration, persistence

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1. Introduction

This study examines the time series behaviour of several Angolan macroeconomic variables using monthly data from August 1996 to June 2011. The variables examined are the inflation rate, M1, M2, the exchange rate at the beginning and the end of the period, as well as the monthly average exchange rate. Both univariate and multivariate analysis are carried out, the latter with the aim of testing for the existence of long-run equilibrium relationships between the variables of interest. Previous studies on African macroeconomics are rare; they include Aiolfi, Catão and Timmermann (2011), Fiess, Fugazza and Maloney (2010), Brito and Bystedt (2010), Cermeño, Grier and Grier (2010), Barros and Gil-Alana (2012), and Barros, Damásio and Faria (2012). Angola is a particularly interesting case to analyse given its oil and diamond wealth. Further, its macroeconomy has undergone two distinct phases, i.e. the war economy and then the oil period. The former was characterised by rampant inflation. After 1996, a restrictive monetary policy brought down the inflation rate. The world economic crisis in 2009 curbed oil demand and generated a terms of trade shock that considerably reduced the growth rate and brought about a fiscal crisis, forcing the country to sign a stand-by agreement of US$ 1.4 billion with the IMF, with the aim of alleviating liquidity restrictions and maintaining a sustainable macroeconomic position. The subsequent resurgence in global oil demand has led to a recovery in Angola's oil production and exports, which, combined with the IMF programme and the tightening of monetary and fiscal policies, has resulted in more solid macroeconomic foundations.

The present study makes a contribution by analysing the time series properties of some key macroeconomic series in Angola using fractional integration and cointegration techniques not previously applied to African countries. The layout of the paper is the following: Section 2 provides some background information. Section 3 briefly reviews the literature. Section 4 outlines the methodology, which is based on the concepts of fractional integration and cointegration respectively for the univariate and multivariate analysis. Section 5 presents the empirical results, while Section 6 contains some concluding comments.

2. The Angolan Economy

Angola, situated on the West coast of Africa above Namibia and below Congo, is a leading African country in terms of natural resources. After obtaining independence from Portugal in 1975, and a long war lasting 27 years (1975–2002), the country has been rebuilding its economy, relying largely on its oil production. However, poverty is still high among the urban population living in slums, allegation of corruption practices are common, the non-tradable and industrial sector is still underdeveloped, media control and persecution of journalists common and there is evidence that the country's wealth does not reach the general population. According to the IMF Country Report No. 11/346 of December 2011, Angola's fiscal accounts have shown large residual financing items, cumulatively equivalent to about US$32 billion (25 percent of GDP) from 2007 to 2010. The Angolan authorities put forward a number of explanations for this unaccounted money; however, the Human Rights Watch (2011) has also identified previous irregularities, with more than $4 billion in oil revenues disappearing between 1997 and 2002, pointing to mismanagement and suspected corruption. According to the corruption index 2011 from Transparency International [Guardian, 2011] Angola is among the most corrupt countries in the world. This supports the "rentier state" theory that explains underdevelopment in terms of the rent-seeking behaviour of the government. Angola is characterised by neopatrimonialism, i.e. a form of governance based on using state resources to secure the loyalty of "clients" in the population, with the distinction between public and private virtually disappearing. In such autocratic or oligarchic regimes the upper and middle classes are excluded from power, the leader enjoys absolute personal power, and the army is usually loyal to the leader rather than the nation (Weber, 1947). Previous research on Angola has focused mainly on productivity and efficiency issues (see, e.g., Barros and Dieke, 2008, and Assaf and Barros, 2011) whilst no macroeconomic studies have been carried out. This is the aim of the present paper.

3. Literature Review

The fractional integration or I(d) models applied in this paper to Angolan macroeconomic series have been widely used in the case of developed countries. For instance, regarding inflation persistence, Backus and Zin (1993) found a fractional degree of integration in US monthly data. They argued that aggregation across agents with heterogeneous beliefs results in long memory of the inflation rate. Hassler (1993) and Delgado and Robinson (1994) provided strong evidence of long memory or I(d) behaviour in the Swiss and Spanish inflation rates respectively. Baillie et al. (1996) examined monthly post-World War II CPI inflation in ten countries and found evidence of long memory with mean-reverting behaviour in all countries except Japan. Similar evidence was found by Hassler and Wolters (1995) and Baum et al. (1999). Bos et al. (1999, 2001) examined inflation in the G7 countries, finding that the evidence of long memory is quite robust to level shifts, although it disappears for a few inflation rates when allowing for structural breaks,.[1]

There is less evidence of I(d) behaviour in the inflation rates for developing countries (Kallon, 1994; Moriyama and Naseer, 2009). Almost all existing studies assume integer degrees of differentiation, testing stationarity/nonstationarity with unit

[1] Other recent papers examining long memory and structural breaks in inflation rates are Gadea et al. (2004), Franses et al. (2006) and Gil-Alana (2008), and forecasting issues are examined in Franses and Ooms (1997) and Barkoulas and Baum (2006).

root tests and cointegration techniques. Gaomab (1998) used an error correction model based on cointegration for the inflation rate in Namibia. Masale (1993) tested for stationarity of the inflation rate in Botswana and found evidence of non-stationarity and cointegration with South African prices. Moriyama and Naseer (2009) forecast the 3-month average inflation rate in Sudan with data from January 2000 to October 2008 using an ARMA(4,5) model. The estimated model tracked the surge in inflation during 2007 and 2008 once the external shock—bread price inflation due to world wheat price inflation—was taken into account. Chhibber et al. (1989) estimated a multivariate model for Zimbabwe, including both monetary factors and economic fundamentals as determinants of inflation. A similar model for Ghana was employed by Chhibber and Safik (1990) and Sowa and Kwakye (1991).

Regarding unit roots in monetary aggregates, a fractional integration approach has been taken in many works. Porter-Hudak (1990) started this line of research with a study using a seasonal long memory model for several US monetary series. Baum and Barkoulas (1996) also tested for fractional integration in US monetary aggregates and their various components. They found fractional orders of integration between 1 and 2 in all series. Similar results are reported in other papers by Gil-Alana (2006) and Caporale and Gil-Alana (2008).

There have also been a number of empirical studies investigating the dynamics of nominal and real exchange rates. Anthony and MacDonald (1999) analysed seven currencies in Exchange Rate Mechanism (ERM) using standard univariate unit root tests and found some evidence of stationarity, specifically mean reversion in both the wide band and the narrow band ERM. Mean reversion in real exchange rates was also tested by Bleaney et al. (1999), using data from countries with episodes of high inflation. Their results suggest that a unit root model is more appropriate than models with fixed rates of mean reversion.

Sollis et al. (2002) proposed new tests to analyse mean reversion, based on smooth transition autoregressive models. These included a test that forces mean reversion to be symmetric about the integrated process central case, and another that permits asymmetry. In comparison to the usual unit root tests, these provide stronger evidence against the unit root null hypothesis in a number of real exchange rates calculated using the US dollar and the Deutsche mark (DM). Similarly, using a sample of G7 countries, Campa and Wolf (1997) found strong evidence of mean reversion, and showed that the deviations of real exchange rates and trade from trend are uncorrelated.

In line with the above studies, Caporale and Gil-Alana (2004) examined whether real exchange rates are cointegrated with real interest rates and labour productivity differentials in the DM-US$ and Japanese Yen-US$ relationships. They found evidence of unit roots in all the individual series. Then, by estimating parametrically the memory of the cointegrating error, they conjectured the existence of fractional cointegration. A study with a similar focus is due to Dufrénot et al. (2006), who explored the real exchange rate misalignments of five European countries in the period 1979-1999. They posited an equilibrium relationship between real exchange rates and macroeconomic fundamentals (terms of trade, prices, foreign assets, fiscal wedge, interest rate differential), and given the evidence of unit roots in the individual series, they estimated the memory of the cointegrating residuals by means of the modified R/S statistic of Lo (1991), the Geweke and Porter-Hudak (GPH, 1983) estimator, and the exact maximum likelihood estimator of Sowell (1992). Their reported strong evidence of fractional cointegration for the Netherlands, and mixed results for France and the UK. The recent literature also explores linkages between exchange rates and fundamentals through fractional cointegration, but no such study has been carried out to date for African exchange rates.

4. The Methodology

We use both univariate and multivariate techniques based on long range dependence (LRD) or long memory processes. There are two possible definitions of long memory. Given a covariance stationary process $\{x_t, t = 0, \pm 1, \ldots\}$, with autocovariance function $E(x_t - Ex_t)(x_{t-j} - Ex_t) = \gamma_j$, according to McLeod and Hipel (1978), x_t displays LRD if

$$\lim_{T \to \infty} \sum_{j=-T}^{T} \left| \gamma_j \right| \tag{1}$$

is infinite. An alternative definition, based on the frequency domain, is the following. Suppose that x_t has an absolutely continuous spectral distribution function, implying that it has a spectral density function, denoted by $f(\lambda)$, and defined as

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

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.,

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

Most of the empirical literature in the last twenty years has focused on the case where the singularity or pole in the spectrum occurs at the 0 frequency, i.e.,

$$f(\lambda) \to \infty, \quad as \;\; \lambda \to 0^+.$$

This is the standard case of I(d) models of the form:

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

where 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. However, fractional integration may also occur at some other frequencies away from 0, as in the case of seasonal/cyclical models.

A natural extension of the I(d) models to the multivariate case leads to the concept of fractional cointegration. Although the original idea of cointegration, as in Engle and Granger (1987), allows for fractional orders of integration, all the empirical work carried out during the 1990s was restricted to the case of integer degrees of differencing. Only in recent years have fractional values also been considered. In what follows, we briefly describe the methodology used in this paper for testing fractional integration and cointegration in the case of Angolan data.

3a. Fractional Integration

There exist several methods for estimating and testing the fractional differencing parameter d. Some of them are parametric while others are semiparametric and can be specified in the time or in the frequency domain. In this paper, we use first a parametric approach developed by Robinson (1994). This is a testing procedure based on the Lagrange Multiplier (LM) principle that uses the Whittle function in the frequency domain. It tests the null hypothesis:

$$H_o: \; d \; = \; d_o, \qquad (4)$$

for any real value d_o, in a model given by the equation (3), where x_t can be the errors in a regression model of the form:

$$y_t \; = \; \beta^T z_t \; + \; x_t, \qquad t \; = \; 1, 2, ..., \qquad (5)$$

where y_t is the observed time series, β is a (kx1) vector of unknown coefficients and z_t is a set of deterministic terms that might include an intercept (i.e., $z_t = 1$), an intercept with a linear time trend ($z_t = (1, t)^T$), or any other type of deterministic processes. Robinson (1994) showed that, under certain very mild regularity conditions, the LM-based statistic ($\hat{r}$)

$$\hat{r} \;\rightarrow_{dtb}\; N(0,1) \qquad \text{as} \qquad T \;\rightarrow\; \infty,$$

where " $\rightarrow_{dtb}$ " stands for convergence in distribution, and this limit behaviour holds independently of the regressors z_t used in (5) and the specific model for the I(0) disturbances u_t in (3). The functional form of this procedure, which has been shown to be the most efficient one in the Pitman sense against local departures from the null, can be found in numerous empirical applications based on his tests (see, e.g., Gil-Alana and Robinson, 1997; Gil-Alana and Henry, 2003; Cunado et al., 2005, etc.).

As in other standard large-sample testing situations, Wald and LR test statistics against fractional alternatives have the same null and limit theory as the LM test of Robinson (1994). Lobato and Velasco (2007) essentially employed such a Wald testing procedure, and, although this and other recent methods such as the one developed by Demetrescu, Kuzin and Hassler (2008) have been shown to be robust with respect to even unconditional heteroscedasticity (Kew and Harris, 2009), they require a consistent estimate of d; therefore the LM test of Robinson (1994) seems computationally more attractive. In addition to Robinson's (1994) approach we also use maximum likelihood-Whittle methods in the frequency domain (Dahlhaus, 1989) and in the time domain (Sowell, 1992).

4b. Fractional Cointegration

Engle and Granger (1987) suggested that, if two processes x_t and y_t are both I(d), then it is generally true that for a certain scalar $a \neq 0$, a linear combination $w_t = y_t - ax_t$ will also be I(d), although it is possible that w_t be I(d - b) with b > 0. This is the concept of cointegration, which they adapted from Granger (1981) and Granger and Weiss (1983). Given two real numbers d, b, the components of the vector Z_t are said to be cointegrated of order d, b, denoted $Z_t \sim CI(d, b)$ if:

(i) all the components of Z_t are I(d),

(ii) there exists a vector $\alpha \neq 0$ such that $s_t = \alpha' Z_t \sim I(\gamma) = I(d - b)$, b > 0.

Here, α and s_t are called the cointegrating vector and error respectively.[2] This prompts consideration of an extension of Phillips' (1991) triangular system, which for a very simple bivariate case is:

$$y_t = \nu x_t + u_{1t}(-\gamma), \tag{6}$$

$$x_t = u_{2t}(-d), \tag{7}$$

for t = 0, ±1, ..., where for any vector or scalar sequence w_t, and any c, we introduce the notation $w_t(c) = (1 - L)^c w_t$. $u_t = (u_{1t}, u_{2t})^T$ is a bivariate zero mean covariance stationary I(0) unobservable process and $\nu \neq 0$, $\gamma < d$. Under (6) and (7), x_t is I(d), as is y_t by construction, while the cointegrating error $y_t - \nu x_t$ is I(γ). Model (6) and (7) reduces to the bivariate version of Phillips' (1991) triangular form when $\gamma = 0$ and d = 1, which is one of the most popular models displaying CI(1, 1) cointegration considered in the literature.

[2] Even considering only integer orders of integration, a more general definition of cointegration than the one given by Engle and Granger (1987) is possible, allowing for a multivariate process with components having different orders of integration. Nevertheless, in this paper we focus exclusively on bivariate cases and a necessary condition is that the two series display the same integration order.

Moreover, this model allows greater flexibility in representing equilibrium relationships between economic variables than the traditional CI(1, 1) formulation.

The method proposed here to examine the hypothesis of fractional cointegration includes the following three steps:

Step 1:

We first estimate individually the orders of integration of the series. For this purpose, we employ parametric and semiparametric methods. In the parametric context, we use first the method proposed by Robinson (1994) described above. As for semiparametric approaches, there are two main types, those based on the Whittle function and those that use log-periodogram-type estimators. We employ both of them. First, we use a "local" Whittle estimator in the frequency domain (Robinson, 1995a), with a band of frequencies that degenerates to zero. This estimator is implicitly defined by:

$$\hat{d} = \arg\min_{d} \left(\log \overline{C(d)} - 2\, d\, \frac{1}{m} \sum_{s=1}^{m} \log \lambda_s \right),$$

$$\overline{C(d)} = \frac{1}{m} \sum_{s=1}^{m} I(\lambda_s)\, \lambda_s^{2d}, \qquad \lambda_s = \frac{2\,\pi\, s}{T}, \qquad \frac{1}{m} + \frac{m}{T} \to 0, \tag{8}$$

where $I(\lambda_s)$ is the periodogram of the raw time series, x_t, given by:

$$I(\lambda_s) = \frac{1}{2\pi T} \left| \sum_{t=1}^{T} x_t\, e^{i \lambda_s t} \right|^2, \tag{9}$$

and $d \in (-0.5, 0.5)$. Under finiteness of the fourth moment and other mild conditions, Robinson (1995a) proved that:

$$\sqrt{m}\,(\hat{d} - d^*) \to_{dtb} N(0, 1/4) \quad as\ T \to \infty, \tag{10}$$

where d* is the true value of d. This estimator is robust to a certain degree of conditional heteroscedasticity (Robinson and Henry, 1999) and is more efficient than other semi-parametric competitors.[3]

Finally, we use the log-periodogram estimator of Robinson (1995b), which is defined as:

$$\hat{d}(l) \;=\; \sum_{j=l+1}^{m} \left(a_j - \bar{a}\right) \log I(\lambda_j) / S_l, \tag{11}$$

where

$$a_j = -\log\left(4\sin^2\left(\frac{\lambda_j}{2}\right)\right), \quad \bar{a} = \frac{1}{m-l} \sum_{j=1}^{p} a_j u_{t-j} + \varepsilon_t, \quad S_l \;=\; \sum_{j=l+1}^{m} \left(a_j - \bar{a}\right)^2, \quad \lambda_j \;=\; \frac{2\pi\, j}{T},$$

and $0 \leq l < m < n$.[4]

Step 2:

We test the homogeneity of the orders of integration in the bivariate systems (i.e., H_o: $d_x = d_y$), 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 $\hat{T}_{xy}$ to log-periodogram estimation. The statistic is:

$$\hat{T}_{xy} \;=\; \frac{m^{1/2}\left(\hat{d}_x - \hat{d}_y\right)}{\left(\frac{1}{2}\left(1 - \hat{G}_{xy} / (\hat{G}_{xx}\hat{G}_{yy})\right)\right)^{1/2} + h(T)}, \tag{12}$$

where $h(T) > 0$ and $\hat{G}_{xy}$ is the $(xy)^{th}$ element of

[3] There exist further refinements of this procedure, e.g., Velasco (1999), Velasco and Robinson (2000), Phillips and Shimotsu (2004), Shimotsu and Phillips (2005), and Abadir et al. (2007). The results obtainted using these methods (not reported) were very similar to those based on Robinson's method (1995a) that are presented in the paper.

[4] Extensions of this method can be found in Moulines and Soulier (1999), Velasco (2000), Phillips and Shimotsu (2002) and Andrews and Guggenberger (2003).

$$\hat{G} = \frac{1}{m}\sum_{j=1}^{m}\mathrm{Re}\left[\hat{\Lambda}(\lambda_j)^{-1} I(\lambda_j)\hat{\Lambda}(\lambda_j)^{-1*}\right], \quad \hat{\Lambda}(\lambda_j) = diag\left\{e^{i\pi\hat{d}_x/2}\lambda^{-\hat{d}_x}, e^{i\pi\hat{d}_y/2}\lambda^{-\hat{d}_y}\right\},$$

with a standard limit normal distribution.

Step 3:

In the third step, we perform the Hausman test for no cointegration set forth by Marinucci and Robinson (2001) comparing the estimate $\hat{d}_x$ of d_x with the more efficient bivariate one of Robinson (1995a), which uses the information that $d_x = d_y = d_*$. Marinucci and Robinson (2001) show that

$$H_{is} = 8s\left(\hat{d}_* - \hat{d}_i\right)^2 \quad \rightarrow_{dtb} \chi_1^2 \qquad as \quad \frac{1}{s} + \frac{s}{T} \rightarrow 0, \qquad (13)$$

with i = x, y, and where s < [T/2] is a bandwidth parameter, analogous to m introduced earlier; $\hat{d}_i$ are univariate estimates of the parent series, and $\hat{d}_*$ is a restricted estimate obtained in the bivariate context under the assumption that $d_x = d_y$. In particular,

$$\hat{d}_* = -\frac{\sum_{j=1}^{s} 1_2' \hat{\Omega}^{-1} Y_j v_j}{2 1_2' \hat{\Omega}^{-1} 1_2 \sum_{j=1}^{s} v_j^2}, \qquad (14)$$

with $Y_j = [\log I_{xx}(\lambda_j), \log I_{yy}(\lambda_j)]^T$, and $v_j = \log j - \frac{1}{s}\sum_{j=1}^{s}\log j$. The limiting distribution above is presented heuristically, but the authors argue that it seems sufficiently convincing for the test to warrant serious consideration.

5. Empirical Results

The dataset includes the Angolan CPI (Consumer Price Index), M1, M2 and the exchange rate of the Kuanza vis-à-vis the US dollar (TRY) at the beginning (TRYB) and the end of the period (TRYE), as well as the monthly average exchange rate (TRY), and covers the period from August 1996 to June 2011, comprising 179 observations. The series were obtained from the Central Bank of Angola.

5a. Univariate Results: Fractional Integration

We employ the parametric approach of Robinson (1994) described in Section 3a. In particular, we adopt the set-up given by (3) and (5), with $z^T = (1,t)^T$, $t \geq 1$, and 0 otherwise, testing H_o (4) for d_o-values equal to 0, (0.001), 2. In other words, the model under the null becomes:

$$y_t = \beta_0 + \beta_1 t + x_t; \qquad (1-L)^{d_o} x_t = u_t \qquad t = 1, 2, ..., \tag{15}$$

under the assumption that the disturbances are respectively white noise (Table 1), autocorrelated as in the model of Bloomfield (1973) (Table 2), or follow a seasonal monthly AR(1) process (Table 3).

[Insert Tables 1 – 3 about here]

We display in the tables the estimates of d using the Whittle function in the frequency domain (Dahlhaus, 1989) along with the 95% confidence bands of the non-rejection values of d using Robinson's (1994) parametric approach.[5] For each series, we report the three cases commonly examined in the literature, i.e., the cases of no regressors (i.e, $\beta_0 = \beta_1 = 0$ in (15)), an intercept ($\beta_1 = 0$), and an intercept with a linear time trend.

The results for each series can be summarised as follows:

a) Inflation rate: this is clearly an I(d) variable with d ranging between 0 and 1, and thus displaying long memory (d > 0) and mean-reverting (d < 1) behaviour. The estimate

[5] Very similar values were obtained with other methods in the time domain (Sowell, 1992; Beran, 1995).

of d is in most cases slightly below 0.5, also implying covariance stationarity. Consequently, log-prices, obtained from the inflation rates are clearly nonstationary, with values of d much above 1 and close to 1.4.

b) Monetary aggregates, M1 and M2: the estimates of d are above 1 with the exception of M1 with seasonal AR disturbances, for which the value is 0.980. In general, the unit root null hypothesis (i.e. d = 1) cannot be rejected for M1, but it is rejected in favour of higher orders of integration, i.e., d > 1 for M2. Thus, M2 displays a higher degree of dependence than M1.

c) Exchange rates: for this series the estimates are all above 1, being around 1.3 in the majority of cases, and the unit root null hypothesis (i.e. d = 1) is rejected in all cases.

According to these results inflation, monetary aggregates and exchange rates are clearly nonstationary variables with orders of integration predominantly above 1 for all three variables. This high degree of persistence indicates that shocks to the series have permanent effects and thus that active policies are required to counteract their effects.

5b. Bivariate framework: Fractional Cointegration

Here we analyse bivariate relationships among the three variables, in particular between

a) Monetary aggregates and prices

and

b) Exchange rates and prices.

Fractional cointegration in a bivariate system requires the two parent series to display the same degree of integration. For this purpose, we need to investigate the properties of the price series at first, since inflation is constructed as the first difference of log-prices. The estimated values of d are displayed in Table 4. It can be seen that the estimates of d are close to 1.5 in all cases, ranging between 1.427 (seasonal AR(1) with an intercept) and 1.541 (Bloomfield with a linear trend).

[Insert Table 4 about here]

In what follows, we focus on prices, M2 and the TCI -monthly average exchange rate, since these three variables have the closest orders of integration. We start by analysing the relationship between prices and M2.

a) M2 and Prices

[Insert Figure 1 about here]

Cointegration between money and prices has been widely tested in the literature together with other variables such as output and interest rates. Analysing long range dependence, Tkacz (2000) suggested that some of these variables may be cointegrated. Caporale and Gil-Alana (2005) examined money demand relationships in five industrialised countries by employing a two-step strategy testing the null of no cointegration against the alternative of fractional cointegration. Evidence of long-run equilibrium relationships was found in four out of the five countries considered. As for African countries, Benbouziane and Benamar (2004) examined cointegration between money and prices in the Maghreb countries. Their results do not tend to support the quantity theory of money, although, as Granger (1986) suggested, money and prices could still be cointegrated if other variables affecting prices are included in the cointegrating set.

Figure 1 shows that the two series move in a roughly similar way. Table 5 reports the semiparametric (Whittle) estimate of d for the two series. All the estimates are above 1 being close to 1.5 in the majority of cases. The Robinson and Yajima (2002) homogeneity test provides strong evidence that the order of integration is the same for the two series (Table 6).

[Insert Tables 5, 6 and 7 about here]

In light of the results in Table 6, the Hausman test for cointegration is employed (Step 3 in the methodology). Following Marinucci and Robinson (2001), we test the null hypothesis of no cointegration against fractional cointegration. Table 7 displays the results, which indicate that the null hypothesis of no cointegration is rejected in practically all cases. Further, the order of integration in the cointegrating regression is clearly smaller than 1.4 in all cases, but changes substantially depending on the bandwidth parameter m. In fact, for values of m higher than 12, the estimated values of d is smaller than 1, which implies mean reversion and hence provides some support to the quantity theory of money.[6]

b) Prices and Exchange Rates

[Insert Figure 2 about here]

Prices and exchange rates might also be linked. Sadorsky (2000) showed that energy future prices were cointegrated with exchange rates, and similar evidence is obtained when using oil prices (Hanson el al., 1993; Schnept, 2008; Abbott et al., 2008; etc.). Focusing on Africa, Vicente (2007) examined the long-run relationship between nominal exchange rates and South African prices to explain consumer prices movements in Mozambique.

Figure 2 shows plots of the two series. Cointegration is less apparent than in the previous case. Table 8 displays the estimates of d using the Whittle semiparametric method (Robinson, 1995a). As in the previous case, they are all above 1 and close to 1.5. In fact, when performing the Robinson and Yajima's (2001) test for homogeneity in the

[6] For $m = (T)^{0.5}$, the estimated d is equal to 0.86, implying mean reversion.

order of integration (Table 9) we cannot reject the null of equal orders of integration in any case.

[Insert Tables 8, 9 and 10 about here]

Table 10 reports the results of the Hausman test for cointegration, which provides evidence of cointegration in most cases. As before, the estimate of d in the cointegrating regression is clearly smaller than 1.5, being equal to or smaller than 1 depending on the choice of the bandwidth parameter. When $m = (T)^{0.5} \approx 13$, the estimated value d is 0.906 and the unit root null cannot be rejected, whilst it is rejected for values of m higher than 16.

5. Conclusions

Angola being a major oil producer, it is of interest to investigate the impact of economic and policy shocks on its macroeconomic variables. Some evidence is provided by the present paper, which focuses on inflation, monetary aggregates and exchange rates. In the first stage, univariate fractional integration or I(d) models are estimated; the values of d are found to be higher than 1 and close to 1.4 in the case of prices, which implies nonstationarity. As for the monetary aggregates, the unit root null hypothesis cannot be rejected for M1, but it is rejected in favour of higher orders of integration for M2, indicating a higher degree of time dependence for this variable. For exchange rates the estimates are all above 1, being around 1.3 in the majority of cases, again implying nonstationarity. In the second stage bivariate relationships among the variables are analysed, in particular between monetary aggregates and prices and between exchange rates and prices. It is found that the two pairs of variables are fractionally cointegrated at least for some bandwidth parameters, with some degree of (slow) mean reversion to the long-run equilibrium relationships.

Overall, this study makes a twofold contribution. First, the univariate analysis provides evidence on the persistence of several Angolan macroeconomic variables, an issue not previously analysed in the literature. The long-memory models estimated here are more general than the classical ones based on integer degrees of differentiation and thus allow for a much richer degree of flexibility in the dynamic specification of the series. The results imply a rejection of stationary I(0) or nonstationary I(1) models in favour of specifications wiht fractional degrees of integration, with orders of integration above 1 for all variables. This high degree of persistence implies that shocks to Angolan macroeconomic variables have permanent effects (the series not being mean-reverting), and therefore policy measures are required to retrieve the original trends. Second, the multivariate results suggest cointegration, at least for some bandwidth parameters, between prices and money, and prices and nominal exchange rates in Angola, with the rate of adjustment to the long-run equilibrium being hyperbolically slow.

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Table 1: Estimates of d assuming that the disturbances are white noise

Series	No regressors	An intercept	A linear time trend
INFLATION RATE	0.433 (0.377, 0.509)	0.399 (0.346, 0.475)	0.358 (0.288, 0.450)
MONETARY BASE M1	1.010 (0.965, 1.071)	1.010 (0.965, 1.071)	1.011 (0.961, 1.077)
MONETARY BASE M2	1.154 (1.095, 1.241)	1.154 (1.095, 1.241)	1.167 (1.104, 1.255)
EXCHANGE RATE TRYB	1.353 (1.265, 1.480)	1.353 (1.265, 1.480)	1.354 (1.266, 1.481)
EXCH. RATE (TRY)	1.381 (1.299, 1.495)	1.381 (1.299, 1.495)	1.381 (1.299, 1.495)
EXCH. RATE (TRYE)	1.394 (1.306, 1.520)	1.394 (1.306, 1.521)	1.394 (1.306, 1.521)

Table 2: Estimates of d assuming that the disturbances are autocorrelated as in Bloomfield (1973)

Series	No regressors	An intercept	A linear time trend
INFLATION RATE	0.529 (0.444,	0.464 (0.388, 0.577)	0.398 (0.283, 0.551)
MONETARY BASE M1	1.154 (1.078,	1.155 (1.078, 1.284)	1.178 (1.085, 1.310)
MONETARY BASE M2	1.180 (1.099,	1.180 (1.099, 1.308)	1.200 (1.109, 1.321)
EXCHANGE RATE (TRYB)	1.188 (1.074,	1.189 (1.074, 1.343)	1.189 (1.074, 1.350)
EXCH. RATE (TRY)	1.343 (1.222,	1.344 (1.222, 1.520)	1.344 (1.222, 1.520)
EXCH. RATE (TRYE)	1.288 (1.169,	1.288 (1.169, 1.453)	1.288 (1.169, 1.453)

Table 3: Estimates of d assuming that the disturbances are seasonal AR(1)

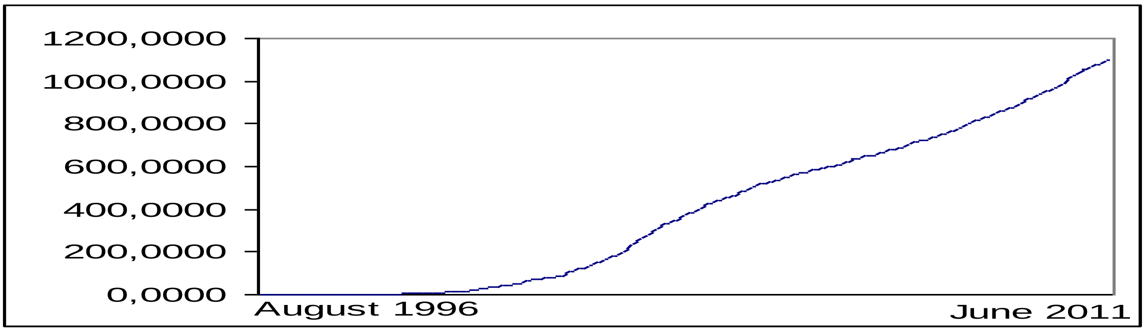
Series	No regressors	An intercept	A linear time trend
INFLATION RATE	0.406 (0.338,	0.382 (0.318, 0.465)	0.348 (0.275, 0.445)
MONETARY BASE M1	0.980 (0.919,	0.980 (0.919, 1.052)	0.978 (0.912, 1.057)
MONETARY BASE M2	1.134 (1.064,	1.134 (1.064, 1.230)	1.146 (1.071, 1.244)
EXCHANGE RATE(TRYB)	1.372 (1.273,	1.373 (1.273, 1.514)	1.373 (1.274, 1.515)
EXCH. RATE(TRY)	1.374 (1.288,	1.374 (1.288, 1.491)	1.374 (1.288, 1.491)
EXCH. RATE(TRYE)	1.389 (1.297,	1.389 (1.297, 1.518)	1.389 (1.297, 1.518)

Table 4: Estimates of d for the price series

Series: PRICES	No regressors	An intercept	A linear time trend
White noise	1.442 (1.382, 1.525)	1.442 (1.382, 1.525)	1.471 (1.419, 1.543)
Bloomfield	1.512 (1.410, 1.652)	1.511 (1.411, 1.651)	1.541 (1.453, 1.661)
Seasonal AR(1)	1.428 (1.363, 1.517)	1.427 (1.363, 1.516)	1.462 (1.404, 1.540)

Figure 1: Prices and M2

i) Prices



ii) M2

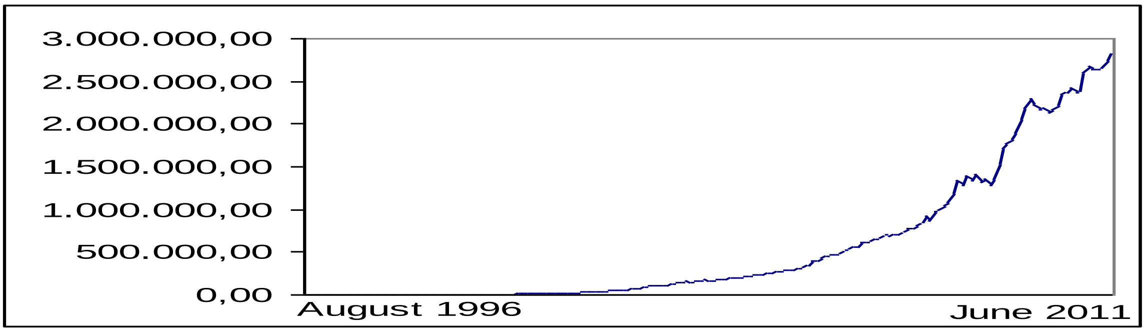


Table 5: Whittle semiparametric estimation method (Robinson, 1995a)

Series / m	5	10	12	13	15	20
PRICES	1.499	1.509	1.476	1.433	1.488	1.497
M2	1.319	1.427	1.385	1.349	1.415	1.228

Table 6: Tests of the Homogeneity Condition for the Orders of Integration

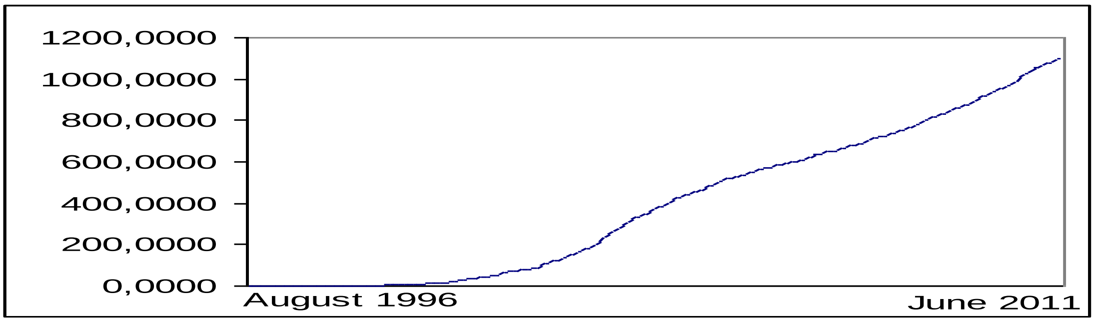
Series / m	5	10	12	13	15	20
PRICES	0.987	0.391	0.493	0.463	0.392	1.503

Table 7: Tests of the Null Hypothesis of No Cointegration against the Alternative of Fractional Cointegration

Series / m	5	10	12	13	15	20
H_x	28.666	26.577	17.532	4.603	0.659	
H_y	38.458	36.837	36.565	25.102	8-183	3.324
d^*	0.952	1.046	1.002	0.860	0.706	0.414

Figure 2: Prices and TCI (Exchange rate)

i) Prices



ii) TCI (Exchange rate)

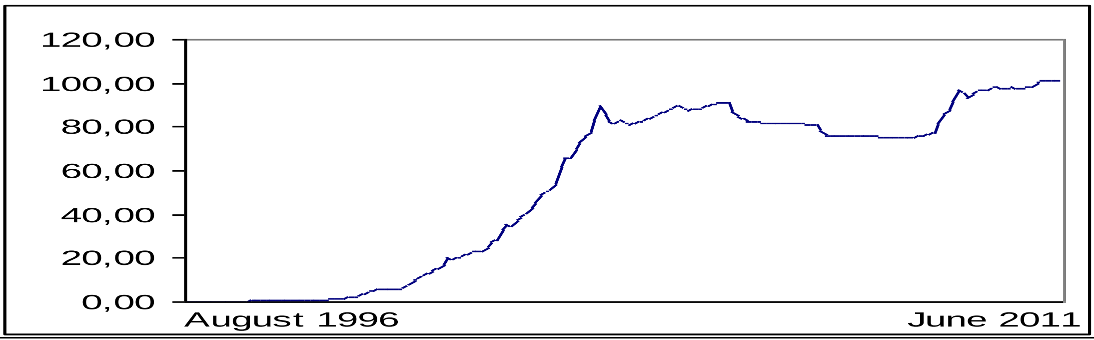


Table 8: Whittle semiparametric estimation method (Robinson, 1995a)

Series / m	5	10	12	13	15	20
PRICES	1.499	1.509	1.476	1.433	1.488	1.497
M2	1.504	1.501	1.450	1.396	1.387	1.412

Table 9: Tests of the Homogeneity Condition for the Orders of Integration

Series / m	5	10	12	13	15	20
PRICES	-0.006	0.014	0.175	0.261	0.702	0.582

Table 10: Tests of the Null Hypothesis of No Cointegration against the Alternative of Fractional Cointegration

Series / m	5	10	12	13	15	20
H_x	43.277	39.991	27.090	21.526	3.713	0.497
H_y	43.148	39.459	29.807	25.021	8.507	0.016
d^*	1.170	1.145	1.007	0.906	0.684	0.424