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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 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

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 results suggest

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

where L

where y_t

4b. Fractional Cointegration

Engle and Granger (1987) suggested that, if two processes x_t and y_t are both I(d), then it is

$_t = y_t \quad ax_t$ 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) $_t \quad Z_t \quad$ b), b > 0.

Here, 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:

$$y_t \qquad x_t \qquad u_{1t}(\quad), \tag{6}$$

$$x_t \qquad u_{2t}(\;d), \tag{7}$$

for t = 0, ±1, ..., where for any vector or scalar sequence w_t, and any c, we introduce the notation $w_t(c) = (1 \quad L)^c w_t$. $u_t = (u_{1t}, u_{2t})^T$ is a bivariate zero mean covariance stationary

0 $_t$ is I(d), as is y_t by

construction, while the cointegrating error $y_t \quad _t$

bivariate version of Phillips' (1991) trian

the most popular models displaying CI(1, 1) cointegration considered in the literature.

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[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-

Whittle estimator in the frequency domain (Robinson, 1995a), with a band of frequencies that degenerates to zero. This estimator is implicitly defined by:

$$d \quad \arg\min_{d} \quad \log \overline{C(d)} \quad 2d\,\frac{1}{m}\,{}^{m}_{s\;1}\log \quad {}_{s} \quad ,$$

$$\overline{C(d)} \quad \frac{1}{m}\,{}^{m}_{s\;1}\, I(\;{}_{s}) \quad {}^{2d}_{s}, \qquad {}_{s} \quad \frac{2\quad s}{T}, \qquad \frac{1}{m} \quad \frac{m}{T} \quad 0, \tag{8}$$

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

$$I(\;{}_{s}) \quad \frac{1}{2\quad T}\left|\,{}^{T}_{t\;1}\, x_t\, e^{i\quad {}_{s}t}\right|^{2}, \tag{9}$$

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

$$\sqrt{} \qquad 4 \quad /4) \qquad , \tag{10}$$

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:

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

$$G \quad \frac{1}{m} \overset{m}{\underset{j\ \ 1}{}} \mathrm{Re} \quad (\ \ _{j})^{\ 1} I(\ \ _{j}) \quad (\ \ _{j})^{\ 1*} \ ,$$

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 $z^T = (1,t)^T$, 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:

$$y_t \qquad {}_0 \qquad {}_1 t \qquad x_t ; \qquad (1 \quad L)^{d_o} x_t \qquad u_t \qquad t \qquad 1, 2, \ldots, \qquad (15)$$

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-

[5] For each series, we report the three cases commonly examined in the literature, i.e., the cases of no regressors ${}_0 \quad {}_1 = 0$ in (15)), an intercept ${}_1 = 0$), 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.

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 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.

[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 = $(T)^{0.5}$ 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

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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)