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Long-Run Exchange Rates Dynamics in Uganda: Evidence from the Sticky Price Monetary Model

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Abstract

The paper uses a hybrid framework that incorporates coffee prices, a major commodity export, as an additional variable to a standard sticky prices monetary model of exchange rate determination to examine the long-run behaviour of exchange rate in Uganda. The results suggest the existence of a significant influence of the standard variables of the sticky prices monetary model in explaining the long-run behavior of the exchange rate. In addition, the coffee prices, the major commodity export, is a highly significant explanatory variable, with the effect possibly propagated through the Balassa-Samuelsson effect. Given that world commodity prices induce exchange rate fluctuations, price signals in the world commodity markets may offer additional information for monetary policy making and inflation control in commodity dependent economies, like Uganda.

JEL Classification: F31, F37, F41

Key words: Exchange rate fundamentals, commodity exports, world coffee prices.

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1 Introduction

The question of exchange rate volatility and predictability has been a subject of academic and policy debate since the collapse of the Bretton-Woods fixed exchange rate system in 1973. Many exchange rate determination models have been developed. These models usually emphasize the importance of different fundamentals in explaining exchange rate behaviour. Indeed, Dornbusch (1980) argues that these theories can be regarded as providing partial explanations of the exchange rate behaviour. Furthermore, the empirical evidence supporting these models remains elusive. In fact Meese and Rogoff (1983) demonstrate that the exchange rate is a random walk since none of the fundamentals-based structural models could reliably out perform a simple random walk model.

In Uganda, the external value of the shilling and its volatility has continued to evoke spirited discussions among politicians and market participants alike. Interests are directed towards finding the underlying forces driving the value and volatility of the shilling. This contributes to these discussions by seeking to find a suitable model that can reliably explain the long-run behavior of the external value of the shilling. To do this, the study revisits the monetary model of exchange rate determination. The monetary model has been of a considerable influence on the empirical literature of exchange rate determination and forecasting because it finds its origin in theory and its relative simplicity and data requirement makes it more applicable on the Uganda data. The model's widespread use in the empirical literature has produced several modifications. The monetary model is one of the numerous structural or fundamental based models that have however, failed empirical regularity. In fact, since the seminal work of Meese and Rogoff (1983) showed that none of the fundamental based models could reliably outperform naïve random walk models in forecasting the exchange rates, none of the numerous attempts have persuasively overturn this finding.

Our approach is to augment the monetary model with additional underlying factors that potentially drive the value of the Uganda shilling and examine whether it improves its empirical performance. As a commodity exporting country and a net importer of petroleum products, the value of Uganda shilling is likely to be influenced by frequent fluctuations of world prices of its commodity export and main import. Given this insight, we augment the monetary model with the world prices of coffee and petroleum and examine whether the inclusion of these factors improves the empirical performance of the model. We attempt to examine the following empirical questions: i) Does the world price of Uganda's coffee exports explain movements in its exchange rates? ii) Controlling for any such co-movement, can the sticky price monetary model fit the data better?

We employ an autoregressive distributed lag model of Pesaran, Shin and Smith (2001) on quarterly data, which spans 2001Q4-2020Q4. Our results show that the sticky price monetary model of exchange rate determination augmented with world prices of a country's major commodity exports is successful in explaining the long-run dynamics of the exchange rate possibly through the Balassa-Samuelsson channel. The US dollar price of coffee appears to be a consistent factor explaining the Uganda shilling - US dollar exchange rate movements. In all the augmented models, a one percent increase in the international price of coffee appreciates the Uganda shilling - US dollar exchange rate by about 0.2 percent on average and this estimate is statistically significant at 5 percent. The rest of the paper is structured as follows. Section 2 presents an overview of the external trade and exchange rate regime in Uganda. Section 3 discusses the methodological issues while section 4 presents the empirical results. Finally, Section 5 presents the conclusion and policy implications.

2 Exchange rate policy in Uganda

Uganda's trade and payments regime has evolved considerably since independence in 1962. In the immediate post-independence era, Uganda pursued an inward looking import-substitution trade and industrial strategy, with a fixed exchange rate regime. Indeed, for most of the 1970s, the official exchange rate with the US dollar was held close to the original rate at which the East African shilling had been fixed, which the Uganda shilling inherited in 1966 after the dissolution of the East African Currency Board. Economic mismanagement at the time, and artificial shortages created by the fixed exchange rate regime led to the emergence of a parallel foreign exchange market. The premium on the parallel foreign exchange market increased dramatically and by 1981, the price of foreign currency in the parallel market was over 10 times higher than the official exchange rate (Atingi-Ego and Sebudde 2003).

In 1981, an adjustment program partly aimed at correcting the exchange rate distortion was initiated. Its centerpiece was a massive devaluation of the shilling, followed by a further devaluation in July 1982. In August 1982, a two-window system was introduced; with key transactions including exports of coffee, tea, tobacco and cotton; imports of petroleum; aid-financed projects; official loan and grant inflows; and the servicing of debts and arrears being carried out through Window I at the official exchange rate; and other transactions falling under Window II through an auction system. The two windows were subsequently merged in 1984 just before the collapse of the adjustment program.[1]

In 1986, there was a brief return to the two-window system before a fixed rate system was again established at the end of 1986. This further aggravated the external disequilibria in the economy. Consequently, a currency reform was undertaken in May 1987 in which one hundred shillings were exchanged for one new shilling, and the shilling devaluated by 77.0 percent in an attempt to address external imbalances. This reduced the parallel market premium substantially. In addition, various schemes, such as the Open General Licence System, the Special Import Programs and the Dual Licensing schemes for exporters wishing to import crucial inputs, were put in place to assist import-dependent industries.

In October 1989, the policy of maintaining the real effective exchange rate constant (a 'crawling peg' system) was introduced. As a result, the nominal exchange rate was adjusted on a monthly basis. In July 1990 the parallel market was legalized, leading to the establishment of foreign exchange bureaux. The bureaux were permitted to conduct spot transactions at freely determined exchange rates. Limits, albeit liberal, were placed on invisible payments in a bid to address concerns about capital flight.

In a further move towards a market based exchange rate regime, a foreign exchange auction system for import support funds was introduced in January 1992. Initially, commercial banks, and later foreign exchange bureaux, were permitted to bid in the auction, provided they were in a good financial footing with the Bank of Uganda. The auction was held weekly under the Dutch auction system, whereby each successful bidder paid its bid price. Eligibility of imports was based on a short "negative" list of goods jointly set by the Government and the donor community. The move to the auction system effectively ended the period of administered exchange rates. The private sector bought foreign exchange at market-determined rates in the bureau market or through the auction. Transactions through the official channel, including Government, were initially conducted at the auction rate and later at an average of the bureau rates. However, while exchange rates were market determined, the foreign exchange market remained segmented. By end October 1993, there was still a premium of about 5.0 per cent between the auction and the bureau rate. Thus, in spite of the introduction of the auction system, exchange rate convergence remained elusive.

[1] The adjustment program almost achieved the unification of the exchange rates before it collapsed.

In order to eliminate the segmented nature of the foreign exchange market and to bring about convergence of the exchange rates, an inter-bank foreign exchange market system was introduced in November 1993. This was expected to provide a more efficient and reliable mechanism for determining the official exchange rate and allocating scarce foreign exchange resources. Authorized dealers, including bureaux were free to set their exchange rates while trading among themselves. Subsequently, on fifth April, 1994, the government accepted the obligations of Article VIII, Sections 2,3 and 4 of the IMF's Articles of Agreement, expressing its commitment to a free and open exchange rate system.

The floating exchange rate system has, nonetheless, presented certain difficulties for the country. First, it has heightened the risk of exchange rate volatility, which is synonymous with the flexible exchange rate system. Second, the adoption of a flexible exchange rate system meant the loss of the exchange rate as a nominal anchor for domestic prices. Finally, the operation of an efficient foreign exchange market may not be technically feasible in a situation where financial markets are underdeveloped.[2] This is where Uganda finds itself today. In order to have a better understanding of the exchange rate fundamentals and dynamics, the next section highlights some theoretical and empirical insights of the theories of exchange rate determination.

3 Methodology

3.1 Specification of the empirical model

3.1.1 Basic Model

Two distinctive approaches standout in empirical literature: i) purely univariate or statistical models, which are based on the inherent structure of past observations; and ii) structural models, which are based on theoretical foundations and use empirical approximation to explain exchange rate dynamics and the future direction of the exchange rate. While univariate techniques are powerful predictors of short-term movements in the exchange rates, they do not provide deeper insights into the underlying forces that drive exchange rate dynamics. On the other hand, structural models, because of the sound theoretical foundations embedded therein, do provide not provide deeper insights into the underlying forces that drive exchange rate dynamics

[2] In view of the linkages between different financial operations, some degree of development in other financial markets is necessary to support the operation of a smoothly functioning foreign exchange market with a floating rate.

hand are based theory. Critical in structural models is the identification of the fundamental forces that determine the long-term behavior of the exchange rate.

The empirical literature on structural models has tended to focus the following theories of exchange rate determination or a hybrid, combining the salient features of the economy and the relevant hypotheses of the theoretical models. i) Purchasing Power Parity model, ii) Monetary model[3], and iii) Portfolio Balance models. We postulate a hybrid monetary models that combines the inherent features of the theoretical model and the salient features of the Ugandan economy.

The premise of the monetary model is that the fundamentals that dive the exchange rate originate from the disequilibria in the money market, with the assumption that domestic and foreign bonds are perfect substitutes. We follow the approach of Dornbusch (1976) and Frankel (1979), among others, which incorporate short-term rigidities. Consequently, the basic premise is that nominal output prices are sticky, thus the goods markets are not always in equilibrium. Furthermore, PPP holds only in the long-run.[4] Consequently, the exchange rate may in the short-run deviate from its long-run equilibrium value.[5] The real interest differential, which allows for changes in the long-run real exchange rate are assumed to be correlated with the unanticipated shocks to the trade balance.

The reduced form of the three variants of the monetary model of exchange rate determination is in a general specification. Thus,

$$s_t = \alpha_0 + \alpha_1(m_t - m_t^*) + \alpha_2(y_t - y_t^*) + \alpha_3(r_t^s - r_t^*) \tag{1}$$

$$+\alpha_4(\pi_t^e - \pi_t^{e*}) + \alpha_5(TB_t - TB_t^*) + u_t,$$

[3]There are three versions of the monetary model: the flexible-price model, which assumes that purchasing power parity (PPP) holds; the sticky-price model or the Dornbusch exchange rate overshooting model, which assumes that prices are sticky, in turn giving rise to substantial short-run overshooting of the nominal and real exchange rates in response to macroeconomic shocks.

[4] In the presence of short-term price stickiness, the PPP condition would be violated temporarily and the relation between interest rates and exchange rate needs to capture these short-term liquidity effects of monetary policy.

[5] MacDonald (1988) and Obstfeld and Rogoff (1996) argue that this type of framework is basically an extension of the Mundell-Fleming model. Consequently, this model is also known as the Mundell-Fleming-Dornbusch (MFD) model.

where s_t denotes the exchange rate of the domestic country at time t, defined as the domestic currency price of a unit of foreign currency, in which case, a higher value is a depreciation, while a lower value is an appreciation. In this study, we take Uganda as the domestic country and USA as the foreign country. m_t is the nominal money balances, y_t the real output, r_t^s the nominal short-term interest rate, π_t^e the expected rate of inflation, TB_t is the accumulated external trade balances, and u_t the error term. Foreign variables are shown by asterisks, $*$. If the flexible price variant of the monetary model holds, α_1, $\alpha_3 > 0$; $\alpha_2 < 0$; and $\alpha_4 = \alpha_5 = 0$. If the sticky price variant holds, then α_1, $\alpha_4 > 0$; $\alpha_2, \alpha_3 < 0$; and $\alpha_5 = 0$. If the real interest differential model holds, α_1, $\alpha_4 > 0$; and $\alpha_2, \alpha_3, \alpha_5 < 0$.

At the theoretical level, the monetary model has endured intense criticisms of its assumptions of perfect substitutability of domestic and foreign bonds, PPP and its reliance on the rational expectations-based equilibrium in the domestic and foreign money markets, which does not allow for the role of speculative bubbles, chartists or noise traders in the determination of the exchange rates. At the empirical level, studies initiated by Meese and Rogoff (1983) show that in the short-run, the monetary model does not outperform naïve random walk models in forecasting the exchange rates.[6] However, in the long-run, typically 1 or more years, the explanatory power of the monetary model of exchange rate determination increases and it outperforms the random walk models (see Mc Donald and Taylor, 1993; 1994).

In this paper, we concentrate on the sticky price variant of the monetary model for the bilateral exchange rate of the Uganda shilling-USA dollar. The empirical representation of the sticky price is as in the following unrestricted stochastic form:

$$s_t = \alpha_0 + \alpha_1 m_t + \alpha_2 m_t^* + \alpha_3 y_t + \alpha_4 y_t^* \qquad (2)$$

$$+\alpha_5 r_t^s + \alpha_6 r_t^{s*} + \alpha_7 r_t^l + \alpha_8 r_t^{l*} + u_t,$$

where r_t^l denote the nominal long-term interest rate, to serve as a proxy for expected inflation, π_t^e. Again as in equation (1), all variables are in logarithms, except for the short-term and long-term interest rates. If the sticky price holds on the Ugandan data set, then it is expected that $\alpha_1, \alpha_4, \alpha_6, \alpha_7 > 0$ (positive), while $\alpha_2, \alpha_3, \alpha_5, \alpha_8 < 0$ (negative).

[6] Random walk models suggest the behavior of the exchange rate can be represented as $NER = \alpha + NER_{t-1} + u_t$ or $\Delta NER = \alpha + u_t$, where α is a constant term and the error term, u_t is white noise with mean zero and constant variance.

3.1.2 Commodity price shocks and the exchange rate

Since our objective is to test the empirical performance of the sticky price monetary model of exchange rate determination when augmented with commodity prices, it is important to identify the channel through which commodity terms of trade shocks affect the exchange rate. We rely on the small open economy model of Chen and Rogoff (2002), in which it is demonstrated that an increase in the world price of a country's commodity exports exerts appreciation pressures on the real exchange rate through its effect on wages and demand for non-traded goods similar to the Balassa-Samuelson effect (see Balassa, 1964 and Samuelsson, 1964). Because domestic prices are sticky and will not adjust immediately to the increase in the prices of a country's commodity exports, the exchange rate will have to do the adjustment to preserve the relative prices of traded and non-traded goods or the efficient allocation mechanism.

A good measure of terms of trade into exchange rate analyses should thus be considered. As discussed in Chen and Rogoff (2002), the use of the traditional measures of terms of trade such as the relative price index of exports to imports are complicated by price stickiness and potential mechanical correlations and endogenous pricing behaviour, which calls for use of direct world prices of a country's commodity exports or imports. The commodity price-augmented monetary equation as below:

$$s_t = \alpha_0 + \alpha_1 m_t + \alpha_2 m_t^* + \alpha_3 y_t + \alpha_4 y_t^* + \alpha_5 r_t^s \qquad (3)$$

$$+\alpha_6 r_t^{s*} + \alpha_7 r_t^l + \alpha_8 r_t^{l*} + \alpha_9 p_t^{com} + u_t,$$

where p_t^{com} represents the world price in US dollar of a country's major commodity exports or imports. The parameter α_9 should enter with a negative sign is a country is a commodity exporter and a positive sign if a country is a net importer of the commodity as discussed above.

3.2 Empirical framework

We apply the autoregressive distributed lag (ARDL) model and ARDL bounds test of cointegration of Pesaran, Shin and Smith (2001). The ARDL models is a class of regression models which include both lags of explanatory variables and the dependent variables in the regression equation. They thus dynamic models in the sense that they include past values of both the dependent and independent series in the regression equations. Assuming only one independent variable, the ARDL model of orders *p* and *q* or ARDL(*p*, *q*), the ARDL model specified as:

$$\mathrm{Y_t} = u_0 + \sum_{\mathrm{i=1}}^{\mathrm{k}} \mathrm{\beta_{0i}\,X_{ti}} + \beta_{1i}X_{t-1)i} + \cdots + \beta_{pi}X_{t-pi)i} \tag{4}$$

$$+\gamma_1 Y_{t-1} + \cdots + \gamma_q Y_{t-q} + e_t$$

where u_0 is the constant, Y_t and X_{ti} are respectively dependent and independent variables, p_i is the lag order of independent variables, q is the autoregressive order of the model, and e_t is the innovations. The number of lags of the i^{th} independent series shown $p_i, i = 1, \dots, k$.

The long-run behavior of the exchange rate is examined using ARDL bounds test of Pesaran et al., (2001). The ARDL bounds test for cointegration is an efficient approach frequently used by practitioners due to its advantages in which, we do not need to have a stationary or $I(0)$ or difference stationary or $I(1)$ series to run the analysis and its efficiency over relatively small samples (Pesaran et al., 2001; Zhai et al., 2017). In order to observe the short and long-run dynamics of the variables, we can derive the unrestricted error correction model.

The ARDL bounds test is formulated in the following conditional error correction form:

$$\Delta \mathrm{Y_t} = \mu_0 + \alpha_0 \mathrm{Y_{t-1}} + \alpha_1 \mathrm{X_{1,t-1}} + \cdots + \alpha_k \mathrm{X_{k,t-1}} + \sum_{\mathrm{i=1}}^{\mathrm{q}} \gamma_{\mathrm{i}}\, \Delta \mathrm{Y_{t-i}} \tag{5}$$

$$+\sum_{\mathrm{j=0}}^{\mathrm{p1}} \beta_{\mathrm{k,j}}\, \Delta \mathrm{X_{1,t-j}} + \cdots + \sum_{\mathrm{j=0}}^{\mathrm{pk}} \beta_{\mathrm{k,j}}\, \Delta \mathrm{X_{k,t-1}} + \mathrm{e_t}$$

where μ_0 is the intercept and Δ is the first difference of the variables. The error correction term is EC_{t-1} is given as,

$$\mathrm{EC_{t-1}} = \mathrm{Y_{t-1}} - \sum_{\mathrm{i=1}}^{\mathrm{k}} \frac{\alpha_{\mathrm{i}}}{\alpha_0} \mathrm{X_{it-1}}. \tag{6}$$

The null hypotheses of cointegration based on the coefficients of the conditional error correction model given in equation (5) and the test is applied as follows:

$$H_0 = \alpha_0 = \alpha_1 = \cdots = \alpha_k = 0. \tag{7}$$

The rejection H_0 leads to the conclusion that there is a significant cointegration between variables. A Wald test statistic is computed, which is compared to the asymptotic limits given by Pesaran et al. (2001). If the test statistic is lower than the lower limit, H_0 is not rejected and if the test statistic is greater than the upper limit, H_0 is rejected and we conclude that there is cointegration between variables. If the test statistic lies in between the lower and upper limits, no conclusion can be reached and a specification of the model is advised (see Pesaran et al., 2001). Pesaran et al. (2001) presents five different cases for the inclusion of the intercept (μ_0) and trend (μ_1) coefficients in the error correction term as follows: i) case 1 - no intercept and no trend; ii) case 2 - restricted intercept and no trend; iii) case 3 - unrestricted intercept and no trend, iv) case 4 - unrestricted intercept and restricted trend; and v) case 5 - unrestricted intercept and unrestricted trend.

3.3 Data and Sample Characteristics

3.3.1 Data

We use quarterly data for the period 2001Q4-2020Q1, chosen to coincide with the period of floating exchange rate regime as discussed in section 2. Broad money, exchange rates and commodity price indices data sets are obtained from the International Financial Statistics database of the International Monetary Fund. The exchange rate is the Uganda shilling-US dollar average exchange rate. Broad money (M2) is used to as the nominal money balances and real output are the quarterly gross domestic product (GDP) of the two countries. The quarterly GDP and all nominal interest rates data sets for the USA are obtained from Federal Reserve Bank of St. Louis database. For USA, the 3-month interbank nominal interest rates are used as a proxy for the short-term interest rates, while the 10-year bond rates are used as long term interest rates. For Uganda, interest rate datasets are obtained from the Bank of Uganda database. The 3-month Treasury bill rates are used to proxy short-term interest rates and for lack of long-term interest rates, we use the 364 treasury bill rates are used to measure long-term interest rates. The quarterly GDP and broad money aggregates are seasonally adjusted. All values are in logarithms except for the interest rates.

3.3.2 Sample Characteristics

Figure 1 presents a visual representation of the series. The Uganda nominal exchange rate appears to capture more or less the frequent swings in the world coffee prices and little of the world crude oil prices. There are clear trends in nominal money balances and quarterly GDP series. Both GDP series for Uganda and USA seem to have had a structural break around 2008-09 possibly due to the 2008-09 global financial crisis. Except the huge jump and subsequent fall in the USA short-term interest rates around 2005-06, both the USA and Uganda's short-term interest rates seem to be stationary. The long-term interest rates of the USA represented by a 10-year bond rate seems to exhibit a downward trend. Note that the nominal exchange rate, coffee prices, and the other regressors, except short-term interest rates look like unit-root processes. But generally, we remain agnostic about the data generating processes of the series until a formal test of unit root is conducted using a more rigorous approach.

We test for unit root using the Zivot-Andrews and ADF tests.[7] The Zivot-Andrews test helps to identify unit root in the series even in the presence of structural breaks. The issue of structural breaks in macroeconomic time series data must be robustly addressed to minimise empirical results that are spurious. Of course structural breaks arise due to many factors, including circumstances such as economic and financial crises, policy or regime shifts, etc. because of this reason, it is crucial to test the null hypothesis of structural stability against the alternative of a one-time structural break. If the presence of structural breaks is not controlled for in empirical specifications, the results may be spurious, as they can be biased towards the erroneous non-rejection of the non-stationarity hypothesis (Perron, 1989, Perron, 1997, Leybourne and Newbold, 2003, Pahlavani, Valadkhani and Worthington, 2005, Harvie and Pahlavani, 2006).[8]

Zivot and Andrews (hereafter ZA) (1992) propose a testing procedure where the time of the break is estimated rather than assumed as an exogenous phenomenon. By endogenously determining the time of structural breaks they argue that the results of unit root hypotheses previously suggested by earlier conventional tests, such as the widely-employed Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979, 1981) test or the Perron (1989) methodology, may be reversed. The Zivot and Andrews (1992) model endogenises one structural break in a series (such as yt) as follows:

$$y_t = \mu_t + y_{t-1} + e_t \qquad (5)$$

$$y_t = \hat{\mu} + \hat{\theta} + DU_t(\hat{T}_b) + \hat{\beta}t + \hat{\gamma}DT_t(\hat{T}_b) + \hat{\alpha}y_{t-1} + \sum_{j=1}^{k} \hat{c}_j \Delta y_{t-j} + \hat{e}_t \qquad (6)$$

[7] The ADF test was conducted is for robustness check.

[8] Perron (1989) note that the ADF or PP tests for unit root could be biased toward non-rejection of unit root if there exists a one-time permanent change in the data.

Equation (6) accommodates the possibility of a change in the intercept as well as a trend break. DU_t is a sustained dummy variable capturing a shift in the intercept, and DT_t is another dummy variable representing a break in the trend occurring at time T_b where $DU_t = 1$ if $t > T_b$, and zero otherwise and is equal to $(t - T_b)$ if $(t > T_b)$ and zero otherwise. The null hypothesis is rejected if the coefficient is statistically significant.

The results are presented in Table 1. There is unit root in the series at the 5 percent level of significance, except the long-run interest rates and GDP for Uganda. Nonetheless, the ARDL bounds test framework is designed to handle series that exhibit a mixed order of integration.

Table 1: Unit root tests

	ADF			Critical values*		
Variable	Deterministic terms	Lag	Test	1%	5%	10%
s_t	constant, trend	1	-2.45	-4.04	-3.45	-3.15
m^*_t	constant, trend	1	-1.39			
m_t	constant, trend	1	-0.82			
y^*_t	constant, trend	1	-2.30			
y_t	constant, trend	1	-4.21			
$price\_c$	constant, trend	1	-1.82			
r_t^*	constant	1	-1.62			
r_t	constant	1	-2.87			
l_t^*	constant, trend	1	-4.51			
l_t	constant	1	-4.23			
	Zivot - Andrews					
s_t	constant, trend	1	-4.47	-5.57	-5.08	-4.82
m^*_t	constant, trend	1	-4.44			
m_t	constant, trend	1	-3.95			
y^*_t	constant, trend	1	-5.06			
y_t	constant, trend	1	-6.07			
$price\_c$	constant, trend	1	-4.33			
r_t^*	constant	1	-3.43			
r_t	constant	1	-3.62			
l_t^*	constant, trend	1	-5.71			
l_t	Constant	1	-5.40			

Notes: * The critical values are the ones used by Dickey and Fuller (1981) and Hamilton (1994).

Source: Author's computation

4 Empirical Results

4.1 Cointegrating Relationship

We fit the model with unrestricted intercept and no trend represented by case 3 as discussed above. The ARDL bounds test results for the classical and augmented monetary models are presented in Tables 2. As can been seen in Tables 2, the existence of a long-term relation among the variables is not rejected at 5 percent level of significance. The Walt test statistics are above the upper asymptotic limits. This confirms that there is a significant long-run relationship between the nominal exchange rate and its underlying exploratory forces.

Table 2: ARDL bounds test for cointegration

		MM3 model	
Significant levels	Test stat	Lower bound	Upper bound
		$I(0)$	$I(1)$
10%	5.58	2.558	3.654
5%		3.042	4.244
1%		4.168	5.548
	Test stat	MM3.3 model	
		$I(0)$	$I(1)$
10%	4.47	2.38	3.52
5%		2.80	4.07
1%		3.77	5.22

Source: Author's computation

4.2 Long-run and short-run coefficients

The empirical estimates are presented in Table 3. **The diagnostic tests are** plausible (see Figures 2 & 3 and Appendix Table 1)[9]. From the results in Table 3, USA dollar price of coffee appear to be a consistent factor explaining the Uganda shilling - US dollar exchange rate movements. In all the augmented models (mm1.1 – mm3.3), a one percent

[9]The Ljung-Box tests and the Ramsey RESET test show no autocorrelations and no model misspecification, respectively. The Studentized Breusch-Pagan test for homoscedasticity shows no heteroskedasticity of the residuals. All measures of coefficient stability and stability in the variances of the coefficients as measured by CUSUM and MOSUM are robust. However, the Shapiro-Wilk test shows absence of normality in the estimated residual. In the augmented models, except the Studentized Breusch-Pagan test for homoscedasticity, there are no issues with normality and autocorrelation. Also, the coefficients and variances of the coefficients are stable overtime.

increase in the international price of coffee appreciates the Uganda shilling - US dollar exchange rate by about 0.2 percent on average and this estimate is statistically significant at 5 percent. Further analysis of parameter stability (Figure 2) and possible regime switches (coefficient of the IT dummy) are robust and statistically significant, respectively.

Evaluating the estimation results from the point of view of economic theory (or the expected signs), in all the augmented models (mm1.1-mm3.3), the estimates most closely conform to the sticky price monetary model of exchange rate determination. In the mm3.3 model, all except the foreign income elasticity has the expected sign and are statistically significant. For example, the USA money supply has a negative coefficient (which implies that an increase in the USA money supply which depreciates the US dollar relatively appreciates the Uganda shilling-USA dollar exchange rate. An increase in the Uganda money supply has a positive coefficient, which implies that an increase in money supply leads to depreciation of the Uganda shilling-USA dollar exchange rate.

The elasticity of the Uganda income has a negative sign, which implies that an increase in domestic income leads to appreciation of the exchange rate. Also, the coefficient on the USA short-term interest rate is positive, implying that a rise in the USA short-term interest rates leads to a depreciation of the Uganda shilling-USA dollar exchange rate possibly due to short-term capital outflows. On the other hand, the elasticity of the Uganda short-term interest rate has a negative sign as expected.

We used the USA 10-year bond rates and Uganda 364-day Treasury bill rates to represent inflation expectations for the two countries. As shown in Table 3, the elasticity of the USA inflation expectation is negative, while the elasticity of the Uganda inflation expectation is positive. An increase in inflation expectation of the USA cause the US dollar to depreciate, relatively leading to an appreciation of the Uganda shilling-USA dollar exchange rate, whereas an increase in inflation expectation for Uganda causes the shilling to depreciate. The coefficient on the IT dummy variable is negative and statistically significant which implies that the adoption of the IT regime may have affected the value of the Uganda shilling-dollar exchange rate. Only the elasticity of the USA income is incorrectly signed, but is correctly signed in other regressions (such as mm1.1) and is statistically insignificant in regression mm2.2.

The coefficient of the error correction term is correctly signed and statistically significant. The coefficient of the error correction term increases in absolute value from

0.13 (classical monetary model) to 0.18 (augmented monetary model). The negative sign of the error correction term and its statistical significance implies that last or previous quarter's deviation of the exchange rate from a long-run equilibrium (or the error) influences the exchange rate's short-run dynamics. The -0.13 to -0.18 coefficient of the error correction term implies that the speed at which the exchange rate returns to its equilibrium after a change in its underlying factors or fundamentals is 0.18 at the maximum, which below average. Thus, it takes a longer time for the exchange rate to return to its equilibrium after a change in its fundamentals.

The empirical results demonstrate that the sticky price monetary model of exchange rate determination when augmented with world prices of a country's major commodity exports is a successful framework for explaining the long-run dynamics of the exchange rate possibly through the Balassa-Samuelsson channel. Commodity price fluctuations may affect the value of a country's currency through various channels. World coffee prices may be a crucial additional fundamental in the empirical exchange rate determination equation for Uganda because being a small open economy, the country is a price taker of the world market prices of its major exports and so that positive shocks to the world commodity prices exert pressure on the value of its currency and since its domestic prices (of non-traded) may be sticky, the increase in world commodity prices may instead be absorbed by the nominal exchange rate (overshooting) in order to preserve the efficient resource allocation mechanism (or prevailing relative prices) between traded and non-traded sectors. This justifies the inclusion of terms of trade in the monetary model of exchange rate determination for commodity dependent economies.

Turning to the classical model (mm1-model with short-term interest rates only), we note that without the coffee price variable, the model performs relatively poorly in terms of their overall fit and signs of the coefficients. The inclusion of coffee prices in the model (mm2.2) not only improves the fit of the model, but also leads to estimates with the expected signs. These results again lend support to the view that terms of trade shocks, measured by commodity price movements, are important in explaining the long-run exchange rate dynamics in commodity dependent economies. Their omission may thus explain some of the earlier empirical failures of the structural models of exchange rate determination.

Table 3: Long-run and short-run coefficients

Dependent variable is s_t						
	MM1	MM2	MM3	MM1.1	MM2.2	MM3.3
Con.	1.24***	0.97**	2.80***	1.32***	1.65**	3.36***
	(0.33)	(0.36)	(0.66)	(0.34)	(0.52)	(0.76)
m^*_{t}	-3.53*	-2.08>	-1.60***	-2.02***	-1.54**	-1.29***
	(1.55)	(1.10)	(0.37)	(0.49)	(0.50)	(0.23)
m_{t}	2.48**	1.99**	2.11***	1.38***	1.63***	1.79***
	(0.92)	(0.75)	(0.32)	(0.21)	(0.32)	(0.18)
y^*_{t}	3.44***	1.07***	-3.41***	1.67***	-0.42	-3.23***
	(0.84)	(0.27)	(0.77)	(0.13)	(0.40)	(0.51)
y_{t}	-4.95*	-3.92>	-1.42***	-1.95***	-1.91***	-0.73***
	(2.44)	(2.12)	(0.41)	(0.57)	(0.67)	(0.20)
r_t^*		0.03	7.92***		1.79**	6.91***
		(1.49)	(1.03)		(0.91)	(0.67)
r_t		-1.71*	-3.43***		-0.68>	-2.33***
		(0.77)	(0.43)		(0.35)	(0.24)
l_t^*			-8.13***			-6.07***
			(0.81)			(0.61)
l_t			3.95***			2.75***
			(0.39)			(0.24)
D08-9	-0.13*	-0.14*	-0.01	-0.06>	-0.06>	0.01
	(0.06)	(0.07)	(0.03)	(0.03)	(0.03)	(0.02)
IT-D	0.09>	0.29***	0.17***	0.04	0.14	0.11***
	(0.05)	(0.05)	(0.02)	(0.03)	(0.03)	(0.02)
$price\_c$				-0.23**	-0.25**	-0.17***
				(0.08)	(0.09)	(0.05)
ecm_{t-1}	-0.08***	-0.05**	-0.13**	-0.16***	-0.12**	-0.18***
	(0.02)	(0.02)	(0.03)	(0.04)	(0.03)	(0.04)
$\Delta(s_{t-1})$	0.22*	0.19	0.19>	0.25*	0.17*	0.11*
	(0.10)	(0.11)	(0.11)	(0.09)	(0.11)	(0.11)
$\Delta(m^*_{\mathrm{t}})$	-0.73>	-0.26	-0.56	-0.45***	-0.18	-0.13
	(0.39)	(0.37)	(0.41)	(0.36)	(0.36)	(0.34)
$\Delta(m_{\mathrm{t}})$	0.89***	0.83***	0.93***	0.83***	0.82***	0.93***
	(0.15)	(0.17)	(0.16)	(0.15)	(0.16)	(0.15)
$\Delta(y^*_{\mathrm{t}})$	-0.48	-1.06	-0.57	-0.39	-1.02	-0.96
	(0.75)	(0.91)	(0.92)	(0.73)	(0.87)	(0.86)
$\Delta(y_{\mathrm{t}})$	-0.14	-0.09	-0.13	-0.08	-0.06	0.05
	(0.14)	(0.15)	(0.16)	(0.14)	(0.15)	(0.14)
$\Delta\ (r_t^*)$		0.001	1.58**		0.44	1.77**
		(0.006)	(0.66)		(0.55)	(0.64)
$\Delta\ (r_t)$		0.001	-0.11		0.11	-0.06
		(0.001)	(0.11)		(0.08)	(0.10)
$\Delta(l_t^*)$			-0.85			-0.84
			(0.65)			(0.63)
$\Delta(l_t)$			0.33			0.30*
			(0.14)			(0.14)
D08-9	-0.004	-0.004	-0.003	-0.004	-0.001	-0.003
	(0.006)	(0.01)	(0.01)	(0.004)	(0.01)	(0.01)
IT-D	-0.004	-0.004	-0.003	-0.45***	-0.001	-0.001
	(0.004)	(0.005)	(0.005)	(0.36)	(0.01)	(0.004)
$\Delta(price\_c_{\mathrm{t}})$				-0.09*	-0.10*	-0.08*
				(0.04)	(0.04)	(0.04)
Adj. R^2	0.987	0.987	0.985	0.988	0.987	0.988
RSE	0.014	0.014	0.015	0.014	0.014	0.015
d.o.f.	53	49	43	51	47	41
F-stat.	318.1	245.1	170.5	291.9	233.2	64.4
p-value	< 2.2e-16	< 2.2e-16	< 2.2e-16	< 2.2e-16	< 2.2e-16	< 2.2e-16
ARDL	3,2,2,2,2	3,2,2,2,2,1,1	3,2,2,2,2,1,1,2,2	3,2,2,2,2,1	3,2,2,2,2,1,1,1	3,2,2,2,2,1,1,2,2,1

Significant codes: 0 '***', 0.001 '**', 0.01 '*', 0.05 '>' 0.1 ' ' 1

Source: Authors' Computation

4.3 In-sample foresting performance

We use a variant of error (residual) metrics to assess the goodness of fit of the models to assess the best performing model. These metrics are the mean absolute error (MAE), mean squared error (MSE), mean percentage error (MPE), symmetric mean absolute percentage error (sMAPE), mean absolute percentage error (MAPE), mean absolute scaled error (MASE), mean relative absolute error (MRAE), geometric mean relative absolute error (GMRAE), mean bounded relative absolute error (MBRAE), unscaled MBRAE (UMBRAE), Akaike Information Criteria (AIC) and BIC. The results are presented in Table 4. The results show that the augmented model (lower AIC, MAE) performs better than the classical monetary model.

Table 4: Goodness-of-fit measures

	Classical Monetary model (mm3)	Augmented monetary model MM3.3)
MAE	0.0101	0.0086
MPE	-1.474082e-05	-1.060663e-05
MAPE	0.00299	0.00257
sMAPE	0.00299	0.00257
MASE	0.755	0.640
MSE	0.000171	0.000125
MRAE	1135547.6	896178.4
GMRAE	1.248	1.113
MBRAE	0.542	0.599
UMBRAE	1.183	1.491
BIC	-368.995	-376.829
AIC	-396.315	-408.507
Observations	72	71

Source: Author's Computation

5 Conclusion and policy implications

In this paper, we use commodity prices as an additional fundamental in the standard monetary model of exchange rate determination to examine its in-sample performance in explaining the nominal exchange rate behaviour for Uganda. The evidence supports the view that for commodity dependent countries, world prices of their commodity exports are important in explaining the evolution of exchange rates. The evidence show that the world prices of coffee exports appear to have a stable and consistent impact on the external value of the over the floating rate periods. An increase in the coffee price index is associated with an appreciation of the domestic currency. The inclusion of commodity prices improves the in-sample fit of the standard monetary model of exchange rate determination. To the extent that world commodity prices induce exchange rate fluctuations, price signals in the world commodity markets may offer additional information for monetary policy making and inflation control in these economies.

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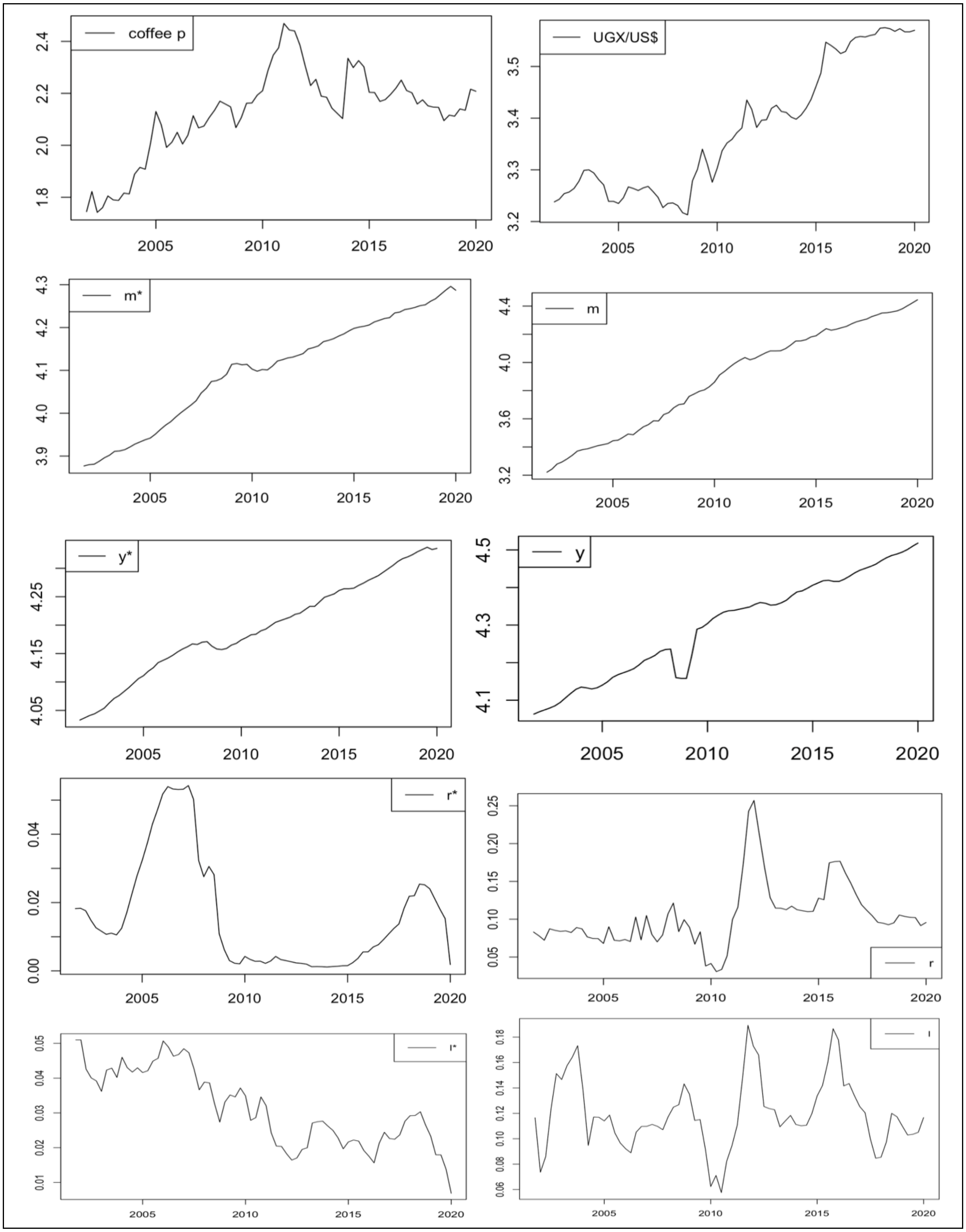
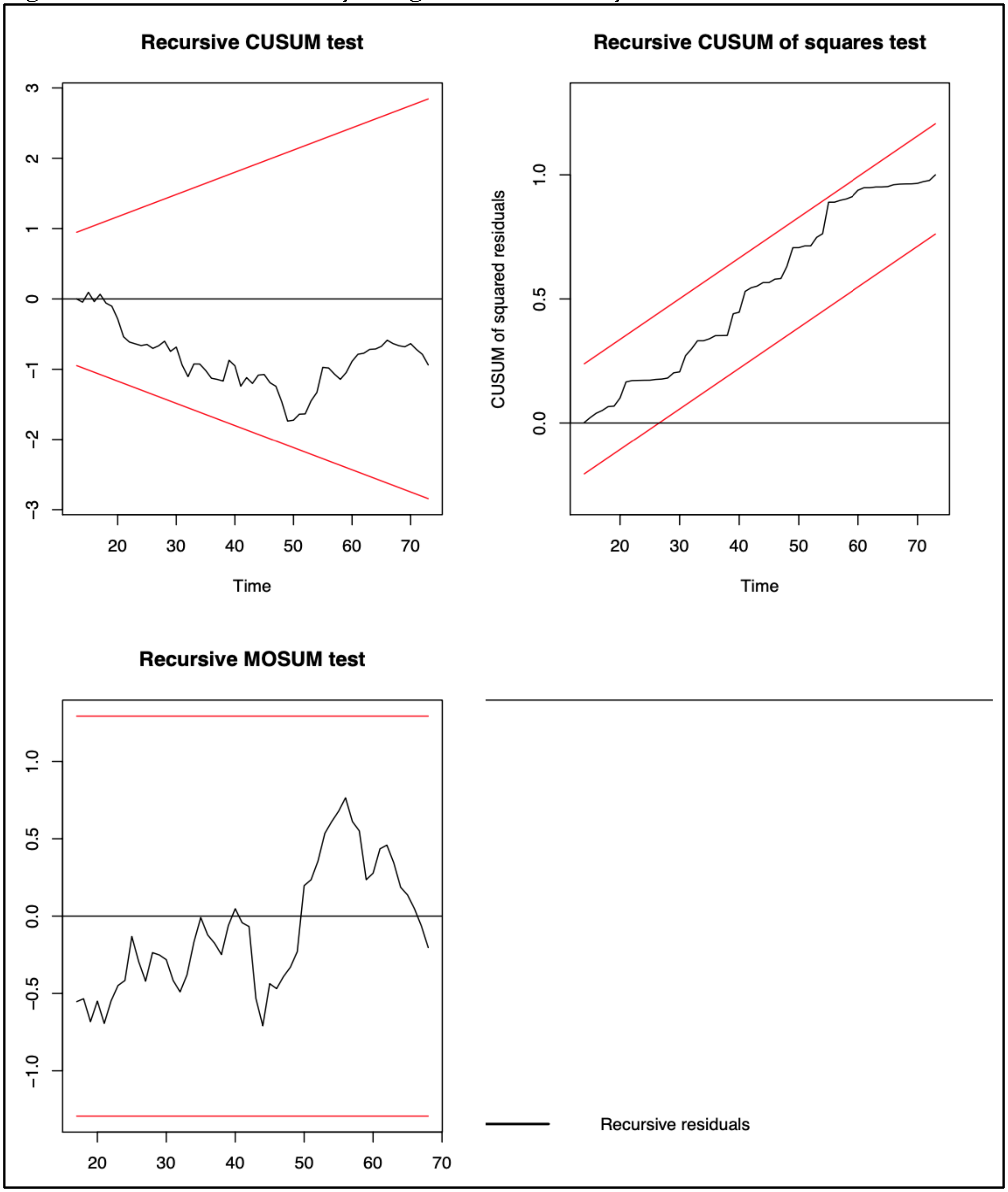
Figure 1: Visual impressions of the time series



Notes: *an asterisk stands for foreign country variable*

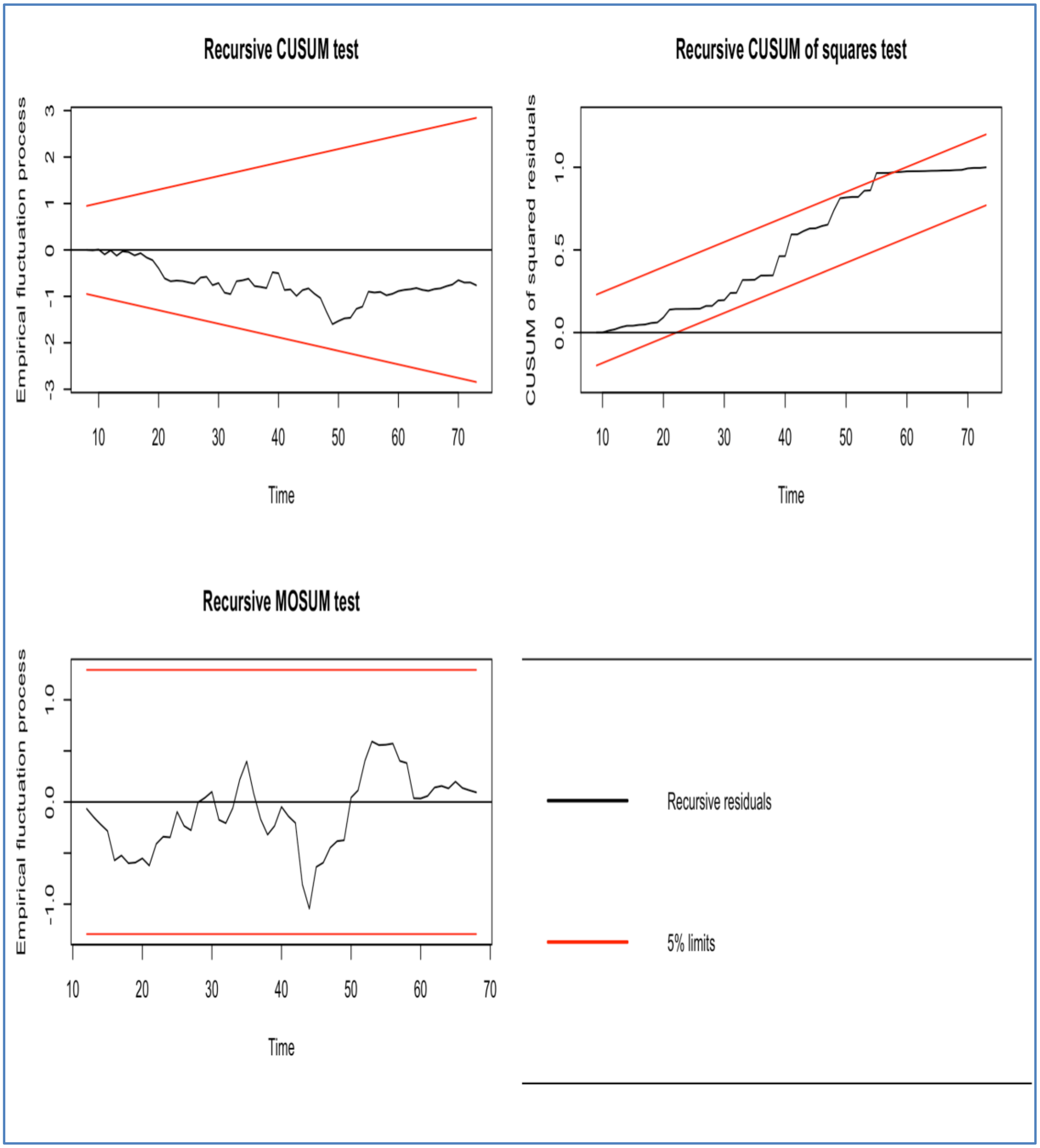
Source: Author's Computation

Figure 2: Parameter stability – augmented monetary model



Source: Authors' Computation

Figure 3: Parameter stability – classical monetary model



Source: Authors' Computation

Appendix Table 1: Diagnostic tests

	Test statistic	Df1	Df2	p-value
Classical monetary model				
Breusch-Godfrey LM Test	0.69213	1	58	0.4089
Ljung-Box X-squared	0.26216	1	…	0.6086
Studentized Breusch-Pagan	15.748	12	…	0.203
Shapiro-Wilk normality	0.93689	…	…	0.001
Ramsey's RESET Test for model specification	0.06584	2	62	0.9363
Augmented monetary model				
Breusch-Godfrey LM Test	0.0873	1	50	0.7688
Ljung-Box X-squared	0.0349	1	…	0.8517
Studentized Breusch-Pagan	30.356	20	…	0.06428
Shapiro-Wilk normality	0.9899	…	…	0.841
Ramsey's RESET Test for model specification	0.4704	2	58	0.6271

Source: Authors' Computation