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I. Ali¹

Estimating the demand for money in Libya: An application of the Lagrange multiplier structural break unit root test and the ARDL cointegration approach

This paper examines the demand for money in Libya using annual data for the period 1970–2010 by applying the Autoregressive Distributed Lag (ARDL) cointegration approach and allowing for endogenous structural breaks in cointegration equation. The results based on the bounds testing procedure confirm that a stable, long-run relationship exists between demand for money and its fundamental determinants; namely, real income, inflation rate and nominal exchange rate. The empirical results indicate that there is a unique cointegrated and stable long-run relationship among real money demand (M1), real income, inflation rate, and nominal exchange rate. The real income elasticity coefficient was found positive while the inflation rate elasticity and nominal exchange rate were negative. This shows that depreciation of domestic currency decreases the demand for money. The results also reveal that after incorporating the CUSUM and CUSUMSQ tests, M1 money demand function is stable between 1982 and 2010.

Keywords: money demand; ARDL; stability; Libya.

JEL classification: E4; E41; E44.

1. Introduction

The stability of the money demand function is a crucial issue in formulating and conducting monetary policy. It provides a reliable and predictable link between changes in monetary aggregates and changes in variables included in the money demand function (Deadman, Ghatak, 1981). It is also critical for selecting monetary policy instruments. Poole (1970) used IS-LM analysis to show that if the money demand is stable the money supply should be targeted. Otherwise, interest rate is the appropriate one. Considering this fact, a steady stream of empirical studies on demand for money function and its stability in developed as well as developing economies have been conducted.

In the context of Libyan economy, and from a policy perspective, it is fundamental to investigate whether there exists a stable long run relationship between real money balances and its determinants variables. Libya is a country heavily dependent on its oil sector, and it has being subjected to a number of shocks and regime shifts over the period of study (1970–2010). These include increasing oil prices in the late 1970s and early 80s, and then again in early 2000s; the changing

¹ Ali Issa — School of Economics, University of Benghazi, Benghazi, Libya; issa.ali@uob.edu.ly.

of economic regime in 1977; the declining in oil prices and production from the late 1980s to the end of 1990s; the USA embargo in the early 1980s and the imposition of the UN sanctions in 1992 after the Lockerbie crisis, isolating the country from the global economy; the devaluation of the official exchange rate in 1999; and the unification of the exchange rate in 2002. In other words, there have been frequent policy changes and/or external shocks to the Libyan economy possibly resulting in the occurrence of structural breaks in the economy and affecting the development of monetary aggregates. That is, these developments may have influenced the relationship between real money balances, real income, prices and other key economic variables making money demand function structurally unstable. Thus, it is crucial to determine if the money demand function is stable throughout the examined period. This is crucial for selecting monetary policy instruments, particularly, in Libya where the Central Bank cannot target the interest rate as monetary policy instrument since it is constant and subjected to regulation by policymakers.

The objective of this paper is two-fold: 1) to investigate the empirical relationship between real money balances (M1) and its determinants in Libya using Auto Regressive Distributed Lag (ARDL) technique; and 2) to examine the long-run stability of real money demand function. The significance of the paper lies in the urgent need to analyse the determinants of demand for money and its stability since no known study based on cointegration technique and structural breaks, in particular that of the autoregressive distributed lag approach, has been done on Libyan economy.

This paper is organised as follows: the empirical studies are discussed in Section 2. Section 3 presents the data and its sources. Here we also test the time series properties of the variables in the presence of endogenous structural breaks. The analytical framework employed to model equilibrium values of the money demand function is outlined in Section 4. In Section 5 we estimate the model by using Auto Regressive Distributed Lag modelling. Discussion on short-run dynamics and adjustment toward long-run equilibrium as well as stability tests are reported and analysed in this section. Section 6 contains the summary and conclusion of the findings.

2. Empirical studies

An extensive work of literature has examined the determinants of demand for money function and its stability in developing countries. Most of the recent studies have utilized ARDL cointegration approach in investigating the long run relationship between the demand for money and its determinants, see for example, (Bahmani-Oskooee, Chi Wing Ng, 2002; Bahmani-Oskooee, Rehman, 2005; Arize, 1994; Akinlo, 2006; Rao, Kumar, 2009; Sharifi-Renani, 2007; Achsani, 2010; Anwar, Asghar, 2013), among others.

Bahmani-Oskooee and Chi Wing Ng (2002) examined the long-run demand for money of Hong Kong using ARDL cointegration procedure on quarterly data over the period 1985Q1–1999Q4. The empirical results proposed that broad money is cointegrated with its determinants. Furthermore, the CUSUM and CUSUMSQ tests confirm the stability of the money demand function.

Bahmani-Oskooee and Shin (2002) estimated Korea's money demand function using ARDL approach to cointegration. The empirical results revealed that while the variables included in the money demand function are cointegrated, the parameters are not stable. They found a structural break in 1997, associated with the Asian financial crisis and claim that this external shock may have led to the instability.

Bahmani-Oskooee and Rehman (2005) estimated the demand for money for seven Asian countries: India, Indonesia, Malaysia, Pakistan, Philippines, Singapore and Thailand, using quarterly data from 1973 to 2000. Applying ARDL approach and CUSUM and CUSUMSQ tests, their results indicate that in some Asian countries even though real M1 or M2 monetary aggregates are cointegrated with their determinants, the estimated parameters are unstable.

Arize (1994) re-examined the money-demand function in three small open economies of Asia: Korea, Pakistan and Singapore. In addition to using the relatively new procedure of error-correction modelling, the roles of variables such as the expected change in the exchange rate, foreign interest rates, and foreign exchange risks on money demand are examined. In testing the importance of these variables in the money-demand function, special attention is paid to testing the assumptions of the classical linear regression model. The sample period for each country spans from 1973:1 through 1990:1. The empirical results suggest that the error-correction specification performs very well. In addition to the traditional variables, the results suggest that at least some measure of foreign monetary developments appear to have some significant effect on money-demand behaviour in these small developing economies.

Akinlo (2006) examined the cointegrating property and stability of money demand for Nigeria, using quarterly data from 1970 to 2002. Utilizing the ARDL approach combined with CUSUM and CUSUMSQ tests, the results indicate that the M2 is cointegrated with income, interest rate and exchange rate. Furthermore, the result revealed somewhat stable relation in particular the CUSUM test.

Sharifi-Renani (2007) estimated the demand for money in Iran utilizing the ARDL approach to cointegration analysis. The empirical results indicated that there is a unique cointegrated and stable long run relationship among narrow money, income, inflation and exchange rate. The empirical results also revealed that the income elasticity and exchange rate coefficient are positive while the inflation elasticity is negative. After incorporating the CUSUM and CUSUMSQ tests, results reveal that the M1 money demand function is stable between 1985 and 2006.

Rao and Kumar (2009) estimated the demand for real narrow money (M1) for Bangladesh for the period 1973 to 2003, using the Gregory and Hansen framework and allowing for endogenous structural breaks in the cointegration equation. The results have revealed that there exists a cointegrating relationship between real narrow money, real income and nominal interest rate. However, of the four possible structural breaks, the one with an intercept shift in 1989 yields meaningful cointegrating coefficients. The results have also indicated that there is a well determined and stable demand for money in Bangladesh from 1988 to 2003 and perhaps following the financial reforms in the 1980s, demand for narrow money has declined by a small amount.

Achsani (2010) employed the vector error correction model (VECM) and autoregressive distributed lag (ARDL) approach to investigate the broad money demand for Indonesia using quarterly data from 1990:1 to 2008:3. The empirical results show that the demand for real money is cointegrated with real income and interest rate. The real income has positive relationship with real money demand, both in the long-run and short-run. On the other hand, interest rate has a negative influence on M2 in the short-run, but has no statistically significant relationship in the long-run. Furthermore, the results show that the ARDL model is more appropriate in predicting stable money demand function of Indonesia in compare to VECM.

Anwar and Asghar (2013) attempted to analyze the long-run relationship between demand for money, real income, inflation rate and exchange rate using ARDL approach for Pakistan. The results of the study reveal that M2 monetary aggregate is cointegrated with its determi-

nants and its long-run relationship with its determinants appears to be stable. The study suggests that monetary authorities and policy makers should focus only on long-run stabilization policy in Pakistan.

Finally, Nitin and Asghar (2016) analysed the stability of the demand for money function in India over the period 1991:1–2014:9 using co-integration and Vector Error Correction Mechanism (VECM) framework. The results have indicated that a long-run relationship does exist between demand for real balances, national output, rate of interest and exchange rate. Two variables national income and exchange rate have been observed to be affecting demand for real balances positively, while the observed effect interest rate is negative.

3. Data sources data stationary

The relevant data were obtained from Central Bank of Libya. The study covers the period 1970–2010, which captures the most prominent structural changes in the economy. Therefore, it is necessary, before starting to perform any empirical estimations of the model, to analyse the time series data as to whether they are stationary or non-stationary. The stationarity properties of a time series (the absence of trend and long-run mean reversion) are scrutinised by carrying out the unit root test to avoid spurious or nonsense regressions. There are a number of methods available for conducting a unit root test², however, we have applied the Augmented Dickey–Fuller (ADF) unit-root test as a benchmark, and the two-break minimum LM unit root test proposed by Lee and Strazicich (2003). The minimum LM unit root test not only endogenously determines structural breaks but also avoids the problems of bias and spurious rejections.

3.1. Traditional ADF unit root test

The most popular and widely used test in the economics literature to examine the stationarity of a time series, in the absence of a structural break, is the Augmented Dickey–Fuller (ADF) test (Dickey, Fuller, 1979, 1981). In the following model, Dickey and Fuller test the null hypothesis against the alternative hypothesis:

$$\Delta y_t = \mu + \beta t + \alpha y_{t-1} + \sum_1^k c_t \Delta y_{t-1} + \varepsilon_t, \quad (1)$$

where Δ denotes the first difference, μ is an intercept, t is the time trend variable, and k is the number of lags which are included in the model to ensure that the error term ε_t is serially uncorrelated, hence obtaining an unbiased estimate of α (i.e. ε_t is white noise with zero mean and constant variance). The null hypothesis of the ADF test is $\alpha = 0$ (non-stationary series) against the alternative hypothesis of $\alpha < 0$ (stationary series), where $\alpha = \rho - 1$. Non-rejection of the null hypothesis implies that the time series y_t is non-stationary, and in this case the usual t -statistic cannot be used, hence the ADF statistic is used. On the other hand, rejection of the null hypothesis signifies the time series is stationary.

² For a detailed discussion about the methods of unit root test see (Ali, Reetu, 2012).

3.2. Minimum LM unit root test with two endogenous structural breaks

Lee and Strazicich (2003) develop a two-break minimum Lagrange Multiplier (LM) unit root test in which the alternative hypothesis unambiguously implies trend stationarity. The main advantages of the minimum LM unit root test suggested by Lee and Strazicich are as follows:

- 1) the break points are determined endogenously from the data;
- 2) the structural breaks are allowed under null and alternative hypotheses;
- 3) avoids the problem of bias and spurious rejections associated with previous tests;
- 4) the LM test enables accurate break point estimation.

Lee and Strazicich consider the data-generating process (DGP) (or parameterisation) as follows:

$$y_t = \delta' Z_t + e_t, \quad e_t = \beta e_{t-1} + \varepsilon_t, \quad (2)$$

where Z_t consists of exogenous variables and ε_t is an error term with mean zero and variance σ^2 . Two structural breaks models are considered by Lee and Strazicich (2003). Model (A) allows for two shifts in level and is given by $Z_t = [1, t, D_{1t}, D_{2t}]$, where $D_{jt} = 1$ if $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise. The term D_{jt} represents a dummy variable for a mean shift accruing at T_B . T_B denotes the time period when the break occurs. Model (C) allows for two changes in the level and trend and is described by $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]$, where $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and 0 otherwise. The term DT_{jt} is an indicator dummy variable for a trend shift accruing at time T_B . According to Lee, Strazicich (2003) the following regression can be used to obtain the LM unit root test statistic:

$$\Delta y = \delta' \Delta Z + \phi \tilde{S}_{t-1} + \mu_t \quad (3)$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta}$, $t = 2, \dots, T$; $\tilde{\delta}$ are coefficients in the regression of Δy_t on ΔZ_t ; $\tilde{\psi}_x$ is described by $y_1 - Z_1 \tilde{\delta}$; and y_1 and Z_1 denote the first observations of y_t and Z_t respectively. The unit root null hypothesis is given by $\phi = 0$, and the LM test statistics are described by $\tilde{\tau}$ (t -statistic) testing the null hypothesis $\phi = 0$. The critical values for the two break case are tabulated in Lee, Strazicich (2003).

3.3. Empirical results for the ADF and the minimum LM unit root tests

The regression results of the ADF, with an intercept term and a linear trend, and the minimum LM unit root tests are reported in Table 1. The inclusion of the trend can be justified in that most of the time series considered here have a trend. The graphs of the time series of interest are revealed in Appendix. The findings in Table 1, based on ADF unit root test, indicate that the null hypothesis of unit root is not rejected for all macroeconomic variables of interest at the five percent significance level. However, these results may be biased towards the non-rejection of the unit root test and the observed unit root behavior, and as Perron (1989) suggested, may have resulted from failure to account for a structural break in the data. Given this and by observing the graphs, it is likely that significant structural changes are very likely to have occurred in the Libyan economy. Therefore, the two-break minimum LM unit root test of Lee and Strazicich (2003) is utilized to analyse whether the time series is stationary or non-stationary, as well as to determine the major structural breaks that can be used in the regression of the model.

The regression results for the two-break minimum LM unit root test are contained in Table 1. One model is considered here; Model (C), which allows for two changes in the level and trend. This can be justified in that most of the time series considered here have a trend (see Appendix). All variables of interest are in log form. Due to the small sample size, a maximum of 4 lags was specified in GAUSS. The results of the LM test show a rejection of the unit root null hypothesis for 2 out of the 4 series. These are consumer price index (*CPI*) and exchange rate (*EXCH*). On the contrary, the rest of the variables are revealed to be non-stationary series. That is, applying the LM test apparently indicates that the other two variables: real money balances (*M*) and real income (*Y*) are non-stationary.

Table 1. Results for the ADF and LM unit root tests

Variables	ADF Test ($k = 2$)		LM unit root test								
	ADF test statistics ¹	Result ⁴	One-break unit root test ²				Two-break unit root test ³				
			t -statistic	T_{B1}	k	Result ⁴	t -statistic	T_{B1}	T_{B2}	k	Result ⁴
Real money balances (M)	−2.5858	NS	−3.7297	1978	0	NS	−4.9724	1981	2001	3	NS
Real income (Y)	−1.0901	NS	−4.4422	1984	3	NS	−4.5149	1984	1999	3	NS
Consumer price index (CPI)	−1.3431	NS	−5.6410	1983	3	S	−6.9750	1987	1999	3	S
Exchange rate ($EXCH$)	−2.0603	NS	−4.5850	1991	0	S	−9.6489	1989	2001	4	S

Notes. ¹ Critical value of $I(0)$ at the 5 percent level is -3.5279 .

² The critical values at the 5 percent significant level are as follows: for *Y* and *CPI* is $\lambda = (0.4) = -4.50$; for *M* is $\lambda = (0.2) = -4.47$; and for *EXCH* is $\lambda = (0.5) = -4.51$.

³ The critical values at the 5 percent significant level are as follows: for *GDP*, *CPI*, and *EXCH* is $\lambda = (0.4, 0.8) = -5.65$; and for *m* is $\lambda = (0.2, 0.8) = -5.71$.

⁴ S = stationary, NS = non-stationary.

On the whole, while the traditional ADF unit root test suggests that *CPI* and *EXCH* are non-stationary, results from the LM method suggest that these time series are trend stationary when the structural breaks are considered under both the null and alternative hypotheses at unknown time in trend function. The two-break points in the level and/or trend for the time series are significant for all time series. As can be seen from Table 1, the first significant break date for the macroeconomic variables in the Libyan economy took place in the early 1980s. These breaks are consistent with the changing of economic regime in the late 1970s and the embargo imposed by the USA in the early 1980s where parts of foreign assets were frozen. The breaks date of 1992 is in line with UN sanctions imposed in the early 1990s. Other second break dates of 1999–2003 correspond to economic reforms in the Libyan economy during the late 1990s and the beginning of this century. These reforms saw the restrictions upon the private sector alleviated and also the lifting of sanctions imposed by the United Nations took place. The breaks also correspond to the depreciation of the official exchange rate in 1999; the unification of the exchange rate in 2002; and the lifting of the UN sanctions and USA trade embargo in 2003 and 2004, respectively.

4. The M1 money demand and ARDL approach

According to conventional money market equilibrium the demand for real money balances depends upon real income as scale variable, and the nominal interest rate and exchange rate

as the opportunity cost of holding real balances. However, the special characteristics of Libya, like most other developing countries, should be considered when specifying the functional form of money demand. Libyan financial markets are immature and capital is restricted due to the constant nominal interest rate. Thus, there is a limited range of alternative financial assets (Ali, Harvie, 2013). Furthermore, the interest rate does not reflect the increase in price levels. As a consequence the interest rate does not reflect the true opportunity cost of holding money in Libya. In addition, since the interest rate is subject to regulation by policymakers it is no longer a good proxy for the costs of holding money but, rather, tends to show the restrictiveness of monetary policy. Therefore, the rate of inflation will be utilised as a proxy variable for the opportunity cost of holding money.

Following Bahmani-Oskooee (1996) and Bahmani-Oskooee, Rehman (2005), our proposed long-run model is based on the specification that real money balances, proxied by M1, is a linear function of real income (y), inflation rate (π) and nominal exchange rate ($exch$):

$$m_t = \alpha_0 + \alpha_1 y_t + \alpha_2 \pi_t + \alpha_3 exch_t + \varepsilon_t. \quad (4)$$

All variables, except the inflation rate, are converted into logs. Our data series is annually data for period 1970 to 2010, and are obtained from Central Bank of Libya (CBL). Nominal values of money balances and income are converted into real values using the consumer price index (CPI) and GDP deflator, respectively.

According to (Arango, Nadiri, 1981) and (Bahmani-Oskooee, Pourheydarian, 1990), the direction of effects of real income and inflation rate is well known and therefore needs no further discussion, however, the effect of exchange rate could be negative or positive. Given that, EXCH is defined as number of units of domestic currency per US dollar, a depreciation of the domestic currency in EXCH raises the value of the foreign assets in terms of domestic currency. If this increase is perceived as an increase in wealth, then the demand for domestic money rises yielding a positive estimate of α_3 (wealth effect). However, if an increase in EXCH spurs an expectation of further depreciation of the domestic currency, public may hold less of domestic currency and more of foreign currency. In this case, an estimate of α_3 is expected to be negative.

To empirically analyse the long-run relationships and dynamic interactions between the variables of interest, the model is estimated by utilizing the ARDL cointegration methodology developed by Pesaran and Shin (1998) and further extended by Pesaran et al. (2001). There are a number of advantages that ARDL has over the other cointegration techniques, such as that of the residual-based Engle and Granger and the cointegrating rank test by Johansen. The advantages of the ARDL approach are as follows. First, the ARDL method, unlike other multivariate cointegration techniques such as the maximum likelihood-based Johansen, does not require the pre-testing of the variables included in the model for unit root tests³ (Pesaran et al., 2001). Second, the ARDL procedure is relatively unbiased, and hence is a more statistically significant approach to determining the cointegration relation for a small sample size as is the case for this study. Third, the ARDL approach avoids the difficulties experienced by the Johansen cointegration technique such as deciding on the number of exogenous and endogenous variables to be included, the treatment of deterministic components, and the order of VAR and the optimal number of lags to be identified. The estimation procedures are very sensitive to such choices

³ However in this paper we test for unit roots to eliminate the possibility of I(2) variables, and also to identify the main structural breaks.

and decisions (Pesaran, Shin, 1998). Finally, the ARDL method is able to distinguish dependent and independent variables when cointegration exists.

The ARDL requires a two-step procedure for estimating the long-run relationships. The first step is to investigate the existence of a long-run relationship among the variables in the model of interest. This can be done by using the F -test. Once a long-run cointegrating relationship is found to exist the second step is to estimate the long and short-run elasticities. Applying ECM determines the short-run adjustment to its long-run equilibrium.

Following (Pesaran et al., 2001) and (Pesaran, Pesaran, 2009), and by including two structural breaks ($D1$ and $D2$) determined endogenously by LM unit-root test, the model can be expressed in the form of an unrestricted error correction model (UECM) format as follows:

$$\Delta m_t = \alpha_0 + \sum_{j=1}^n b_j \Delta m_{t-j} + \sum_{j=1}^n c_j \Delta y_{t-j} + \sum_{j=1}^n d_j \Delta \pi_{t-j} + \sum_{j=1}^n e_j \Delta exch_{t-j} + \delta_1 m_{t-1} + \delta_2 y_{t-1} + \delta_3 \pi_{t-1} + \delta_4 exch_{t-1} + \delta_5 D1 + \delta_6 D2 + \varepsilon_t, \quad (5)$$

where, the dummy variable $D1$ takes on a value of zero prior to the first break date of 1981 (which is consistent with the changing of economic regime in the late 1970s and the embargo imposed by the USA in the early 1980s) and unity thereafter up to the second break date that occurs in 2001 (unification of the exchange rate and the lifting of UN sanctions) when $D2$ takes on the value of one and zero otherwise. α_0 is a drift component, ε_t is white noise error, δ_i , $i = 1, 2, \dots, 6$ are the long-run multipliers, and b, c, d and e are the corresponding short-run dynamic coefficients of ARDL model.

To investigate the existence of a long-run relationship among the variables, unrestricted error correction model (UECM) regressions are estimated. That is, the null hypothesis is tested by considering the UECM for the model, restricting all the lagged variables. Specifically, the null hypothesis of cointegration amongst the variables is $\mathbf{H}_0 : \delta_1 = \dots = \delta_6 = 0$ is tested against the alternative hypothesis $\mathbf{H}_1 : \delta_1 \neq \dots \neq \delta_6 \neq 0$ by the means of F -test with an asymptotic non-standard distribution. Two sets of critical values are tabulated; lower bound critical values assuming the regressors are $I(0)$ and upper bound critical values assuming the regressors are purely $I(1)$. If the computed value is above the upper critical value, the null hypothesis of no long-run relationship is rejected irrespective of the orders of integration for the time series. On the contrary, if the computed value falls below the lower critical value, the null hypothesis is not rejected and the conclusion is that there is no long-run relationship between the independent variable and its determinants. On the other hand, if the F -statistic falls between the lower and upper critical values, the result is inconclusive. In the last case the conclusive inference cannot be made without knowing the order of integration of the regressors. The order of the lags in the ARDL model is selected by the Akaike Information Criterion (AIC).

The error correction version of ARDL model relating to the variables in equation (5) is as follows:

$$\Delta m_t = \alpha_0 + \sum_{j=1}^n b_j \Delta m_{t-j} + \sum_{j=1}^n c_j \Delta y_{t-j} + \sum_{j=1}^n d_j \Delta \pi_{t-j} + \sum_{j=1}^n e_j \Delta exch_{t-j} + \theta ecm_{t-1} + \varepsilon_t, \quad (6)$$

where θ is the speed of adjustment parameter and ecm is the residuals that are obtained from the estimated cointegration model of equation (5).

5. Empirical results and stability tests

Based on the bounds test the computed F -statistic is 7.95, which is above the upper critical bound ($UCB=5.6794$) at the five per cent significance level. This provides conclusive evidence of a long-run relationship between the real money balances and the relevant macroeconomic variables, namely the real income, inflation and exchange rate. Given the existence of a long-run relationship, in the next step we used the ARDL cointegration method to estimate the parameters of equation (4).

Equation (5) represents the demand for real money balances, defined as the nominal money stock deflated by the consumer price level. The AIC lag specification for m is ARDL(2,2,1,0). The results reported in Table 2, where the deterministic time trend was included, suggest that the long-run estimated parameter of the real income elasticity (1.45) has the right sign and is highly significant at the 1 percent level. The inflation rate elasticity (0.41) is negative and also significant at the 1 percent level, supporting the theoretical specification identified above. This means that 1 percent increase in the inflation rate will bring about a decrease in real money demand by 41 percent. And, as the nominal exchange rate coefficient is negative (-0.22) and highly significant, it appears that a depreciation of the Libyan Dinar decrease the demand for money.

The structural breaks for 1981 and 2000 are significant at the 5 percent and 10 percent levels, respectively. The results show that both breaks have a long-run positive effect for the former,

Table 2. Estimated long-run coefficients and short-run error correction model (ECM) for the real money balances

The long-run coefficient estimates based on ARDL(2,2,1,0) and selected lag based on AIC		The short-run coefficient estimates based on ARDL(2,2,1,0) and selected lag based on AIC	
Variable	Coefficient	Variable	Coefficient
Y	1.448*** (0.158)	Δm	0.230* (0.138)
m^{cap}	-0.408*** (0.115)	ΔY	0.511*** (0.113)
$exch$	-0.222*** (0.066)	$\Delta Y1$	-0.145 (0.099)
T	0.0209*** (0.0030)	$\Delta \pi$	-0.0789 (0.0594)
Constant	-3.095*** (0.785)	$\Delta exch$	-0.1387** (0.0499)
		ΔT	0.0131*** (0.0032)
D1982	0.0823* (0.0449)	$\Delta D1982$	0.0514* (0.0311)
D2000	-0.0642 (0.0454)	$\Delta D2000$	-0.0401* (0.0283)
		$ecm(-1)$	-0.625*** (0.103)
$R^2 = 0.82$		$DW = 2.03$	

Notes. ***, **, * represent significance at 1, 5 and 10 percent levels, respectively. Figures in parentheses are standard errors.

and negative effect for the latter on real money demand. Even though the elasticities of all variables experience a considerable decrease during the short-run, most of them still highly significant and have the right signs.

Table 2 shows that the ECM coefficient has the right sign and is highly significant. This confirms the existence of the cointegration relationship among the variables. The coefficient of ECM is equal to (-0.62) , which implies that 0.62 percent of disequilibrium in the previous year is corrected in the current year. This result indicates that the speed of adjustment is marginally high in the demand for money function in the Libyan economy.

In the final stage, the stability tests of CUSUM and CUSUM of Squares proposed by Brown et al. (1975) were applied. The plots of CUSUM and CUSUMSQ, as given in Figures 1 and 2 below, stay within the 5 percent critical bounds. This reveals that demand for narrow money in Libya is stable from 1982 to 2010, and this stability in the function implies the stability in money multiplier and, thus, ensures the changes in the monetary aggregates to have a specific predictable impact on the output and inflation. Consequently, the Central Bank of Libya can use the monetary aggregate as its intermediate target for monetary policy.

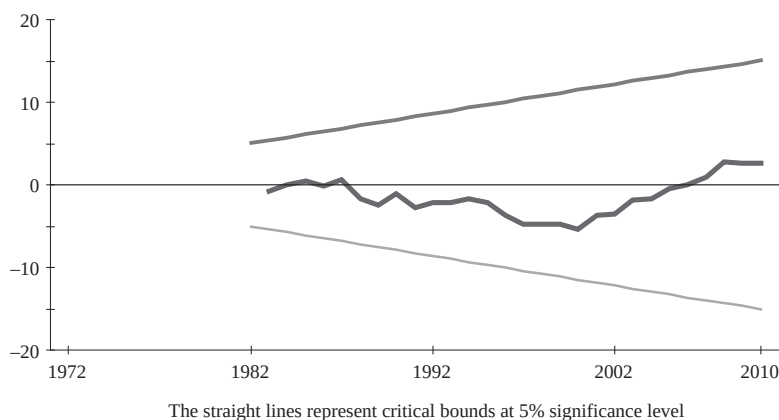


Fig. 1. Plot of cumulative sum of recursive residuals

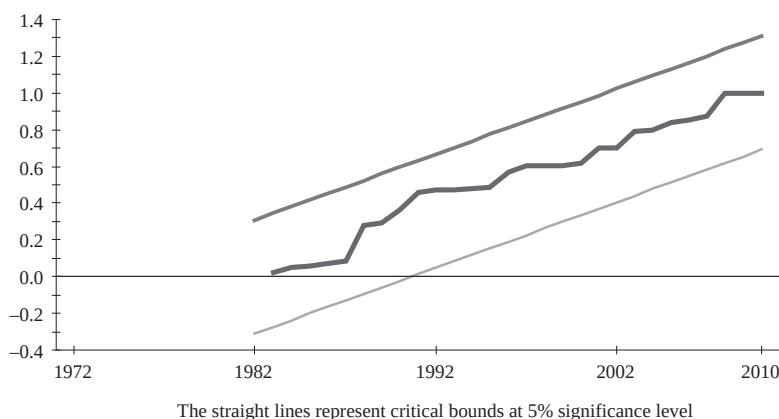


Fig. 2. Plot of cumulative sum of squares recursive residuals

6. Conclusion

In this paper, we have used LM technique for unit root test in presence of structural breaks, and ARDL technique to investigate the existence of a long-run relationship among the variables in the model as well as to estimate the long and short-run elasticities of the demand for money function in the Libyan economy for the period 1970–2010. The result reveals that there exists a cointegrating relationship between real narrow money, real income, inflation rate and exchange rate after allowing for structural breaks. The empirical results have shown that, all the variables in the model are statistically significant and consistent with the demand theory both in long-run as well as short-run.

The results indicate that real income is positively associated with real money balances while inflation rate and exchange rate negatively affect real money balances. The negative effect of inflation rate on real money balance supports our theoretical expectation that the inflation rises, the demand for money decreases. This shows that people may prefer to substitute physical assets for money balances. The negative effect of exchange rate on real money balances shows that the depreciation of the Libyan Dinar decreases the demand for money.

Furthermore, when the model was subjected to CUSUM and CUSUMSQ stability tests, both tests showed stability. This implies that demand for narrow money is temporally stable in Libya and, therefore, it can be said that monetary aggregate is the appropriate intermediate target for monetary policy for the Central bank of Libya.

References

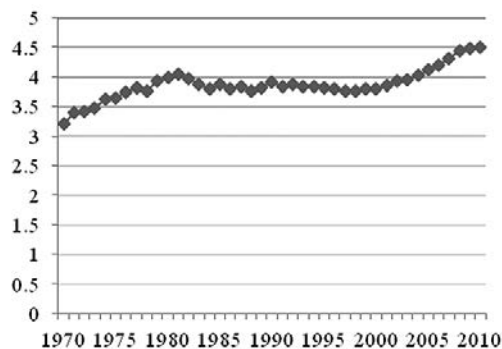
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Appendix. Plots of macroeconomic data for the Libyan economy

Real Money Balance



Real Income



Consumer Price Index (2003=100)



Nominal Exchange Rate

