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Estimating the Demand for Broad Real Money Balances of Saudi Arabia Using the Time Series Approach to Econometrics

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Abstract. This paper adopts the time series approach to present an econometric study for the demand for broad real money balances in Saudi Arabia for the period 1972.1-1989.4. We provide statistical evidence that the series entering the long run relationship are of unit roots. We provide an ECM and show that it is surprisingly stable over the sample period and therefore immune to Lucas critique.

Introduction

Finding a stable real money demand function has always been of particular interest for macroeconomic theory and for policy implications. The literature of estimating such function for both developed and developing countries is voluminous. The hypotheses underlying the estimation of such function, however, vary from one study to another. More recently and in particular after the publication of the pathbreaking article of Davidson, *et al.* [1], Error Correction Models (ECM) have been widely used in modelling both the consumption and the money demand function in the UK and elsewhere. The origins of the ECM's and their developments are given in Asseery [2], while surveys for the theory and empiricisms of the money demand function are given in, inter alia, Laidler [3], Goldfeld [4, 5] and Goodhart [6], Judd and Scadding [7], Asseery [2] and Asseery [8].

In this paper we will provide an empirical investigation into the broad real money of Saudi Arabia over the period 1972.1 to 1989.4. The paper is divided into four sections; while section 1 introduces the subject of the study, section 2 discusses the results of testing for unit roots for all the variables entering this relationship as well as testing for cointegration. Section 3 provides an ECM and the results of testing the arrived at model for stability over the sample period. Section 4 provides concluding remarks.

Testing for Unit Root and Cointegration

As it is well known, it has been shown that, in equilibrium, real money balances are a function of volume of transaction and an opportunity cost variables. The precise measures of these variables, however, are not universal. As is shown in the surveys presented by Laidler [3] and Judd and Scadding [7], a number of studies have used various definitions for monetary aggregate, the proxy of volume of transaction/wealth and the interest rate/inflation. In this study we used the broad definition of money (M3) as defined by the Saudi Monetary Agency (SAMA). For the interest rates we experimented at initial stages with a spectrum of foreign rates ranging from a very short one to a long run interest rate of the Seven Industrial Countries as well as the averages of the short and long interest rates of these Big 7.

We found, however, that a short interest rate and a long one of United Kingdom are the apropriate in terms of the signs and the magnitude of the coefficients as well as their comparability to those used by other studies, interalia, Artis and ~lewis [9], Longbottom and Holly [10]. For the proxy of transaction we used real non-oil-GDP seasonally adjusted. This variable is essentially the same one adopted by Asseery [8] in his extensive study to the demand for narrow real money balances in Saudi Arabia. In his study he demonstrated that the GDP itself usually used when estimating this function for Saudi Arabia is a "perronian" series, i.e., it is characterized by a single break with a change in the slope. Indeed his results demonstrate that the recent argument of Macdonald [11], which claims that the hypothesis of stationary velocity of Saudi Arabi can not be rejected only when the GDP is taken as the proxy of volume of transactions, is spurious.

In view of the results of Asseery [8], the precise definitions of the variables appearing in this study are: lm is the logarithm of broad real money balances (not seasonally adjusted), lq is the logarithm of real non-oil-GDP, lrs is the logarithm of a short interest rate and lrl is the logarithm of a long interest rate. The data are quarterly and they were abstracted from the publications of the International Monetary Fund (IMF) using the DATASTREAM.

The literature of cointegration provides many tests for unit roots. The most popular tests in applied work are Dickey and Fuller [12, 13], Phillips and Perron [14] and Perron [15]. The results of testing the levels of the series are shown in Table 1, while Table 2 shows the results of the same tests for the first differences of the series.

In both tables the reported values in the top line in front of each series are the t-ratios of the D-F test whithout the trend, followed by the t-ratios of the same test but with the trend. The third and the fourth lines of values in front of each series are the calculated values of P-P test without and with the trend, respectively. In comparing these results with the critical values given in Fuller [16] it is evident that the logarithm of real money balances, real non-oil-GDP and both interest rates have the same order of integratedness, i.e. ~ I (1). Therefore, the variables that enter the cointegrating regression for the demand

Table 1. Results for testing for unit roots: "Levels"

		p = 0	1	2	3	4
lm	D-F	−3.48	−2.43	−2.31	−2.09	−1.94
		1.10	.098	−.581	−.890	−1.09
	P–P	−3.54	−2.99	−2.44	−2.46	−2.32
		1.13	.479	1.43	.289	.142
lq	D–F	−2.48	−1.08	−.906	−.709	−.595
		−2.18	−2.42	−2.50	−2.71	−2.79
	P–P	−2.53	−1.96	−2.05	−2.00	−2.07
		−2.22	−2.55	−2.52	−2.72	−2.72
lrs	D–F	−3.49	−3.73	−3.22	−3.53	−3.15
		−3.36	−3.65	−3.24	−3.54	−3.19
	P–P	−3.45	−3.58	−3.62	−3.64	−3.65
		−3.43	−3.49	−3.54	−3.57	−3.59
lrl	D–F	−2.15	−2.03	−1.74	−1.96	−1.93
		−3.45	−3.30	−3.19	−3.45	−3.71
	P–P	−2.19	−2.24	−2.25	−2.29	−2.32
		−3.53	−3.52	−3.53	−3.57	−3.53

Table 2. Results for testing for unit roots: "First differences"

		p = 0	1	2	3	4
Δlm	D–F	−4.50	−2.56	−1.93	−1.58	−1.80
		−5.31	−3.38	−2.67	−2.26	−2.90
	P–P	−4.56	−4.31	−7.12	−4.55	−4.74
		−5.43	−5.28	−8.12	−5.48	−5.63
Δlq	D–F	−16.5	−8.13	−6.73	−5.72	−4.13
		−16.0	−7.52	−5.92	−4.73	−3.32
	P–P	−17.6	−18.0	−19.3	−20.5	−21.0
		−16.7	−16.8	−17.2	−17.4	−17.2
Δlrs	D–F	−7.32	−5.39	−4.43	−4.21	−4.08
		−7.31	−5.31	−4.37	−4.10	−4.03
	P–P	−7.43	−7.42	−7.43	−7.44	−7.44
		−7.48	−7.46	−7.48	−7.49	−7.48
Δlrl	D–F	−7.48	−6.14	−4.15	−3.71	−3.73
		−7.62	−6.30	−4.26	−3.83	−3.92
	P–P	−7.59	−7.59	−7.57	−7.58	−7.58
		−7.79	−7.79	−7.77	−7.77	−7.77

Note: D-F is the Dickey-Fuller test P-P is the Phillips-Perron test.

for real money balances are those which are I(1). We estimate the cointegrating regression by OLS and both the estimated coefficients and their t-values are shown in Table 3,a.

Furthermore, we applied Phillips and Hansen [17] procedure (P-H) which takes into account a semiparametric correction for serial correlation and endogeneity. This is asymptotically equivalent to maximum likelihood procedure. These fully modified estimators have asymptotic mixed normal distributions and thus permits quite general

Table 3. Testing for cointegration a: The cointegrating regression dependent variable lm

	Constant	lq	lrs	lrl	R^2	CRDW
OLS	2.20	1.01	.226	−1.20	.962	1.46
	(4.81)	(23.9)	(2.00)	(-4.79)		
Phillips-Hansen	2.39	1.02	.345	−1.55	.960	–
P–H Procedure	(5.23)	(23.1)	(1.86)	(−2.63)		

WALD Tests restrictions:
Price homogeneity, Wald (chi sq 1) = 3.040
Unitary income elasticity, Wald(chi sq 1) =.0408
Joint significance, Wald(chi sq 3) = 1075

b: Testing the residuals for unit roots:

		p = 0	1	2	3	4
OLSR	P–P	−6.36	−6.30	−6.66	−6.68	−6.55
P-HR	P–P	−6.61	−6.60	−6.90	−7.10	−6.98

Testing for cointegration using Johansen estimation procedure

Likelihood ratio test of Max r vectors:			5% Simulated critical values	
Test of r =0 versus r=1			75.22	47.21
Test of r<=1 versus r=2			40.59	29.86
Test of r<=2 versus r=3			19.82	15.41
Test of r<=3 versus r=4			2.887	3.762
Normalized cointegrating vectors:				
lm	1	1	1	1
lq	.972	−.083	1.95	1.25
Lrs	.350	7.63	11.2	−.873
lrl	−1.60	−4.62	−15.6	1.17

hypothesis tests using conventional methods. We started by examining the cross-correllogram between the fitted residuals and the first difference of both regressors as well as the correllogram of the residuals themselves.We found that a lag truncation of four is appropriate to construct $\hat{\Omega}$. The estimated long run covariance and its various parts are given below:

$$\Sigma = \begin{bmatrix} .035 & -.024 & -.002 & .0004 \\ -.024 & .045 & .002 & -.0001 \\ -.002 & .002 & .014 & .0048 \\ .0004 & -.0008 & .005 & .004 \end{bmatrix}$$

$$\Delta = \begin{bmatrix} .054 & -.007 & -.004 & -.007 \\ -.032 & .038 & .007 & .002 \\ -.009 & .003 & .015 & .007 \\ .001 & .002 & .004 & .004 \end{bmatrix}$$

$$\hat{\Omega} = \begin{bmatrix} .072 & -.072 & -.011 & -.006 \\ -.015 & .031 & .008 & .003 \\ -.011 & .008 & .015 & .007 \\ -.006 & .003 & .007 & .005 \end{bmatrix}$$

The P-H fully modified coefficients and their t-values calculated from their asymptotic standard errors are shown in Table 3,a. It should not be surprising that the magnitude of the classical t-ratios and the coefficient of determination are higher than their counterparts obtained via P-H procedure, since the formers are biased upward.The bias of OLS, furthermore, is not limited to the statistics but also involves the magnitude of the coefficients. The bias is downward and it is more pronounced with respect to the coefficients of interest rates. This finding has already been observed by Asseery [2].

The same table shows that cointegrating regression Durbin-Watson (CRDW) clearly indicates that the variables are cointegrated at the 1% level of significance. The critical values for this test for cointegration are given by Sargan and Bhargava [18] and by Engel and Yoo [19]. The residual obtained from the cointegrating regression was tested for stationarity using the drift-free D-F and P-P regressions with p also varying from 0 to 4. The results in this table clearly suggest that the residuals are stationary at least at the 1% level. The critical values of this test can be found in Phillips and Ouliaris [20) and MacKinnon [21].

Also in the same table we report Wald statistics for testing various hypothesis within the format (R- r) = θ, in which, θ is k-vector of regression coefficients, R is mx1 fixed matrix and r is mx1 fixed vector.

The statistics in Table 3,a show that all regressors are significant determinants for real money balances judging by both their calculated t-values and Wald tests for the joint significance. The price homogeneity is marginally accepted by the data and the rates of interest are significantly different from zero. The Wald statistics for testing the hypothesis that income elasticity is not different from unity is accepted. Since CRDW on its own is not enough evidence for cointegration between the variables involved we used Johansen [22] procedure. The reported results of the Johansen procedure shown in Table 3,b reject the hypothesis that there is no cointegrating vectors at the 5% level. The results, however, suggest that there are at most three cointegrating vectors. The normalized vectors are reported in the same table and they carry the expected signs. One of them, however, seems to be reasonable in terms of the magnitude of the coefficients which they do compare with the P-H cointegrating regression.

ECM and Testing for Stability Over the Sample Period

At this stage one may follow Phillips and Loretan [23], see Asseery [24], in formulating the Error Correction model or proceed to the second step of Engel-Granger two-step method following, inter alia, Enders [25], Kim [26], Asseery and Peel [27] and

Asseery and Peel [28]. In this study we preferred to follow the "Hendryfication" approach. The feed back model is estimated using the reduction procedure, starting from a general distributed lag model.The base line is of the form:

$$\mathrm{lm}_t = a + \sum_{i=1}^{5} b_i \ \mathrm{lm}_{t-1} + \sum_{j=0}^{5}\Big[c_{1j}\ \mathrm{lq}_{t-j} + c_{2j}\ \mathrm{lrs}_{t-j} + c_{3j}\ \mathrm{lrl}_{t-j}\Big]$$
$$+\ c_{4j}\ \mathrm{lp}_{t-j}\ \Big] + d_1 S1 + d_2 S2 + d_3 S3 + e_t \tag{3}$$

The final stage of the reduction process is reported in Table 4. The dynamics of the arrived at model were selected exclusively by reference to the data and the congruency of the specification is statistically supported by the fact that it parsimoniously

Table 4. The specification for real money demand function based on generalized instrumental variables estimation

Modelling lm_t of Eq(3) using GIVE, Sample 1972: 1-1989:4; Endogenous regressors: (lq_t)

Variables	Coefficient	t-value
Constant	-.095	-.949
S1	-.014	-1.67
S2	-.018	-2.02
S3	-.020	-2.36
lm_{t-1}	1.13	19.9
lm_{t-4}	-.143	-2.99
lq_t	,068	2.52
lq_{t-4}	-.055	-2.52
lrs_{t-1}	-.023	-2.27
lrs_{t-5}	.0275	1.60
lrl_t	-.0742	-1.50
lrl_{t-3}	.054	1.49

S.e. = .023 R^2 = .999 DW = 2.21
ARCH F(1,65) = 1.73 AR(1-5) F(5,48) = 1.96
SARGAN instrument test with 3 D.F. = .488
RESET test for linearity(4) = 8.0
Ljung-Box(12) = 9.96 (mild spike at lag nine)
B–J(2) = .297
F-test for restrictions:
F(21 , 34) = .762

Notes:

The normality assumption of the residuals is tested by B-J, see Jarque and Bera [30], and cannot be rejected. The Ljung-Box at lag 12 indicates that the residuals are white noise. AR 1-5 is the Lagrange Multiplier test for autocorrelation up to the fifth lag order. ARCH is the Engel's test for Autoregressive Conditional Heteroscedasticity. RESET is the Ramsey test for functional misspecification. Sargan is the Sargan's test for the validity of the instruments.

Long-run (Steady-State) solution:
lm_t = Const. + 1.00 lq_t + .345 lrs_t - 1.55 lrl_t

encompasses the base line; the F-test on the twenty one restrictions imposed by the final specification on the base line model is F(21, 34)=.762. The procedure used to estimate the final specification is the Generalized Instrumental Variables using lq_{t-1}, lq_{t-2}, lq_{t-3} as additional instruments for lq_t. The validity of the instruments is not rejected judging by Sargan test which is distributed as a Chi.Sq. (3).Therefore, as argued by Nelson and Startz [29], the high correlation between lq_t and its own lagged values rules out potential problems generated by poor instruments.

This final model has been subjected to a battery of test statistics. The long-run (steady state) shown at the bottom of the table is comparable with the cointegrating regression of the P-H and Johansen procedure.

At this stage we impose the long run solution as the Error Correction term and the equation reported in Table 5 is reparameterised as follows:

$$
\begin{aligned}
\Delta lm_t = {} & \underset{(4.16)}{.163}\ \Delta_3 lm_{t-1} + \underset{(4.43)}{.105}\ (\hat{l}q_t - lq_{t-4}) \\
& - \underset{(-2.48)}{.033}\ \Delta_4 lrs_{t-1} - \underset{(-1.70)}{.066}\ \Delta_4 lrl_{t-1} \\
& - \underset{(-2.71)}{.049}\ \big[lm_{t-4} - 1.03\ lq_{t-4} \\
& - .345\ lrs_{t-4} + 1.55\ lrl_{t-4} - .409\ \big] \\
& + \underset{(2.74)}{.021} - \underset{(-1.92)}{.017}\ S1 - \underset{(-2.29)}{.020}\ S2 - \underset{(-2.43)}{.021}\ S3
\end{aligned}
$$

S.e. = .022 $\quad R^2$ = .633 $\quad$ DW = 2.20

ARCH F(1,65) = 1.00 $\quad$ AR(1−5) F(5,48) = 1.02

RESET test for linearity F(4,57) = 1.22

B-J(2) = .029

The above equation looks statistically reasonable in terms of the magnitude of the estimated coefficients and their corresponding t-ratios. The coefficient of the Error Correction term is on the low side which is consistent with most recent studies of money demand function. The most striking result is the presence of the short-run dynamics of the interest rate. This coupled with the low value of the coefficient of the Error Correction term probably suggests that deviation of real balances from equilibrium is maintained through time by the adjustment in short-run interest rates rather than by velocity.

Since most designed tests for cointegration can only be taken as guidance for cointegration and to ensure that the variables under investigation are cointegrated the

Error Correction term should be statistically significant. It is obvious from the equation presented above that this term is statistically significant at the 1% level which is another evidence that real money balance, real non-oil-GDP and interest rates are cointegrated. Other variables included are also statistically significant at the 1% level and the equation passed impressively all diagnostic tests.

Two issues here concerning the above estimated equation are worth noting. The first is concerning the dynamic misspecification and structural break and the second is the performance of this equation over subsamples. To deal with the first issue we reestimated the equation over the sample up to 1986.4 reserving the last 12 data points for the predictive accuracy test, Hendry [31].The reestimated equation is:

$$
\begin{aligned}
\Delta lm_t = \; & .169 \; \Delta_3 lm_{t-1} + .106 \; (\hat{l}q_t - lq_{t-4}) \\
& (4.19) \qquad\qquad (4.19) \\
& - .034 \; \Delta_4 lrs_{t-1} - .069 \; \Delta_4 lrl_{t-1} \\
& (-2.36) \qquad\qquad (-1.70) \\
& - .055 \; [lm_{t-4} - 1.03 \; lq_{t-4} \\
& (-1.91) \\
& - .345 \; lrs_{t-4} + 1.55 \; lrl_{t-4} - .409 \,] \\
& + .024 - .024 \; S1 - .028 \; S2 - .027 \; S3 \\
& \;(2.93) \quad (-2.61) \qquad (-2.99) \qquad (-2.91)
\end{aligned}
$$

S.e. = .022 $\quad R^2 = .643 \quad$ DW = 2.19

ARCH $F(1,57) = 1.30$ $\quad$ AR(1 − 5) $F(5,38) = 1.07$

RESET test for linearity $F(4,45) = .742$

$Xi(12) = 9.9$

It is clear from the value of the test Xi that the equation is dynamically stable over the sample period. This result, however, should not be taken as to say that the estimated equation is immune to structural breaks. It is known that the period of estimation has witnessed many political and economic events. It witnessed three monarchs, two oil shocks, boom, recession and the Iraq/Iran conflict. Therefore, one might argue that the equation can never be structurally stable. For this reason we ran the recursive regresion, see Chow [32] and Hendry [31], and computed the MeanChow test, see Hansen [33] and the MaxChow test. The computed values are, 3.20 for the MaxChow and 1.50 for the MeanChow. On the basis of these tests we can conclude that there is no sign of instability of the ECM specification for broad real money demand function over the sample period.

For the second issue we reestimated the equation for the period 76.3-86.4 and the result is:

$$
\begin{aligned}
\Delta lm_t = {} & .163 \ \Delta_3 lm_{t-1} + .134 \ (\hat{l}q_t - lq_{t-4}) \\
& (2.26) \qquad\qquad (3.72) \\
& - .029 \ \Delta_4 lrs_{t-1} - .096 \ \Delta_4 lrl_{t-1} \\
& (-1.70) \qquad\qquad (-1.98) \\
& - .055 \ [lm_{t-4} - 1.03 \ lq_{t-4} \\
& (-2.42) \\
& - .345 \ lrs_{t-4} + 1.55 \ lrl_{t-4} - .409 \ ] \\
& + .021 - .017 \ S1 - .031 \ S2 - .029 \ S3 \\
& (1.78) \ (-1.39) \qquad (-2.57) \qquad (-2.43)
\end{aligned}
$$

S.e. = .026 $\qquad R^2 = .660 \qquad$ DW = 2.19

ARCH $F(1,65) = 1.00 \qquad$ AR(1 − 5) $F(5,21) = 1.23$

RESET test for linearity $F(4,27) = .171$

B - J(2) = .442

It is clear from this result that all variables retained their explanatory powers, and no cause for alarm has sprung out of the diagnostic tests.Therefore, we can conclude that the specification of the ECM is robust to changes in sample size.

A further note on the structural stability of the ECM is the impact of misspecification of the marginal process. It has been demonstrated by Hendry [31] that the stability of the parameters in the ECM, when the parameters of the marginal process are not stable, rules out the interpretation of the ECM as a reduced form of the forward looking model EVEN IF the structural instability of the marginal process is caused by the omission of some important variables. Here a test procedure should be designed in a way that it is robust when the information set used by the econometrician in the specification of the marginal process is a subset of the information used by agents in forming their expectations.This issue is quite important since the formulation of the ECM's has less restrictive lag structure and they typically let the parameters be chosen by the data generation process.These parameters may be interpreted as a mixture of agent's decision rules and the parameters of the historically contingent structure of the exogenous variables. Consequently, they seem to be quite vulnerable to Lucas critique, see Lucas [34]. Therefore this work may be complemented by testing for the stability of the marginal process. We will investigate this issue in a separate study.

Conclusion

This paper provides an empirical investigation into the broad real money demand function of Saudi Arabia taking into account that the variables entering the equation are, a priori, potentially non-stationary integrated variables. Recently an argument against the

ECM's approach has been proposed in the literature, favoring the Euler equation approach to structural models of money demand function. Even if the latter fit the data well they are not reliable for policy simulation because of their vulnerability to Lucas critique. The evidence from Saudi data cast doubt on this argument. A line of future research implied by this study is the microfoundation of the ECM.

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تقـديـر دالة الطلب على النقود الواسعة الحقيقية للمملكة العربية السعودية باستخدام مدخل السلاسل الزمنية للاقتصاد القياسي

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ملخص البحث. هذه الدراسة تستخدم مدخل السلاسل الزمنية للاقتصاد القياسي لتقديم نموذج تقديري لدالة الطلب على النقود بمفهومها العريض أو ما يسمى بـM3 . في هذه الدراسة قدمنا شاهدًا إحصائيًا على أن السلاسل الزمنية المعبرة عن المتغيرات التي تدخل في هذه الدالة في الأجل الطويل تتصف بجذور وحدوية . بناءً على ذلك فقد قمنا بوضع نموذج ECM والذي أثبت مناعته ضد نقد LUCAS .