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Empirical analysis of the elasticity of real money demand to macroeconomic variables in the United Kingdom with 2008 financial crisis effects

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This research work has employed vector error correction and cointegration techniques in order to estimate the elasticity of real money demand to macroeconomic variables such as industrial production index, exchange rates and short-term interest rates in the United Kingdom. Also, global financial crisis was introduced as an impulse variable to capture structural breaks inherent in the series. Empirical results showed that long-run relationships existed between real money demand and industrial production index, short-term interest rates, and exchange rates in the United Kingdom. The study showed that in the long-run, real money demand had more than unity elasticity with industrial production index in both economies. Real money demand has an inelastic relationship with short-term interest rates and exchange rates. Furthermore, results indicated that it would take long time for real money demand to adjust to its long-run equilibrium. Impulse response analysis revealed that any increase in short term interest rates will have negative effects on the real money demand in the medium to long-term. Whilst real money demand in the United Kingdom tend to be more significant in forecasting the Euro zone money demand, the latter tends to be negatively statistically significant in the former real money demand model. The financial crisis witnessed globally had negative effects on real money demand in the United Kingdom.

Key words: Vector error correction, Cointegration, impulse response analysis, macroeconomic variables, long-run equilibrium, real money demand and financial crisis.

INTRODUCTION

Money demand models provide a structure, which helps to explain changes in money explained by advances in macroeconomic variables. They symbolize a normal yardstick against which tends to measure monetary advancements. This therefore, having a firm long-run and short-run money demand is very imperative, as the presence of a well-specified and stable relationship between money and macroeconomic variables can be seen as requirement for the use of monetary aggregates in the conduct of monetary policy.

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The stability of this relationship is usually assessed in a money demand framework, where money demand is linked to other macroeconomic variables like industrial production index (used as a proxy for real economic output) and interest rates. This research work focuses on studying the elasticity by estimating long-run and short-run money demand function for the United Kingdom by adopting the method of cointegration and error correction analysis. Many factors affect the demand for money. These factors include, level of prices, level of interest rates, the level of real national output (real GDP) and speed of financial innovation. In addition to the variables, which are usually considered within money demand analysis, exchange rates play crucial role. During periods of high inflation, some countries experience partial replacement of domestic currencies by foreign currencies, either as a store of value or a medium of exchange. Hence, the exchange rate is an important factor explaining money demand. It is assumed that the interest rates are significant in money demand models. According to Keynes theory (1936), there are three justifications for the demand of money; transactionary, precautionary and speculative motives. Keynes (1936) theory implies that interest rates have an inverse relationship with the speculative money balances and there exist an indirect transmission mechanism, which depends strongly on the interest rates effect on investment and through the multiplier effect on real sector of the economy. Theoretically, the income velocity of money is not stable and does not depend upon the rate of interest. Keynes (1936) also showed that the transactionary demand for money is positively linked to real incomes and inflation. Hence, the quantity of nominal money demand is proportional to the price level in the economy. Similarly, the precautionary is positively correlated with real incomes and inflation. The total demand for money is obtained by summing the transaction, precautionary and speculative demands.

Prior to this research, many studies have looked at the relationships between macroeconomic variables and the real money demand in the United Kingdom. This research has gone further to study the impact of real money demand in the Euro Area on the United Kingdom. This is important as it will add to literatures on the significance of the real money demand in the United Kingdom to the European Monetary Union or vice versa. There are extensively rich literatures on the relationships between money demand and determinants such as real economic activities, exchange rate, long-term interest rates and inflation. For example, Hendry and Ericsson (1991) using recursive procedures to derive cointegrated model showed that money demand model is uniquely different from models of prices because constancy holds only conditionally on long-run prices in the United Kingdom and United States. Similarly, by employing cointegration and error correction techniques, Skrabic and Tomic-Plazibat (2009) emphasized that in addition to industrial production index, exchange rate explains the most variations of money demand in the long-run while interest rates is significant only in the short-run in Croatia. Drilsaki showed that interest rate causes the largest shift in money demand in addition to industrial production in Turkey between 1989 and 2010. Frait and Komárek (2001) argued that in a monetary model of the exchange rate, a depreciation of the domestic currency is likely to induce extra demand for domestic goods from abroad and the induced rise in domestic production implies higher domestic inflation rate and a need for more money in the economy as the amount of transactions increases. However, according to the currency substitution approach, depreciation reduces the confidence in the domestic currency, thereby lowering money demand via a substitution effect with foreign money. Hence, its coefficient should be negative. Orlowski (2004) also stressed the implication of exchange rates risks for Hungary, Poland and the Czech Republic, countries well known for their inflation targeting monetary policies. Doornik et al. (1998) using practical cointegration rank under restrictive dynamics showed that the long-run ratio of money demand is negatively related with interest rates and inflation rates in the UK. Similarly, based on a correlation analysis, Antczak (2003) pointed out the importance of money growth for steadying inflation rates in some transition economies of Europe. Further, Bahmani et al. (2013) by studying the impact of economic and monetary uncertainty on money demand in emerging economies of six Central and Eastern European countries showed that money demand is transitory and monetary targeting irrespective of output and monetary uncertainty can be effectively stable. These empirical studies suggested the following functional form for the money demand function as M/CPI where M represents a narrow monetary aggregate, CPI is the consumer price index (which is CPI deflator).

This paper intend to use cointegration and error correction with unrestrictive dynamic techniques to justify the presence of contemporaneous relationships between real money demand, industrial production index, short-term term interest rates and exchange rate in the United Kingdom. In addition, the relationship between real money demand in both the UK and Euro Area was also studied. Also, impulse response function analysis was adopted to ascertain the responsiveness of real money demand to shocks in the macroeconomic variables. The pound sterling to dollar exchange rates was considered in the analysis.

The source of data

The data employed in this research work are monthly observations of industrial production index (IPI), consumer price index (CPI), short-term interest rates (INR), exchange rates (EXR), and narrow money supply M1.

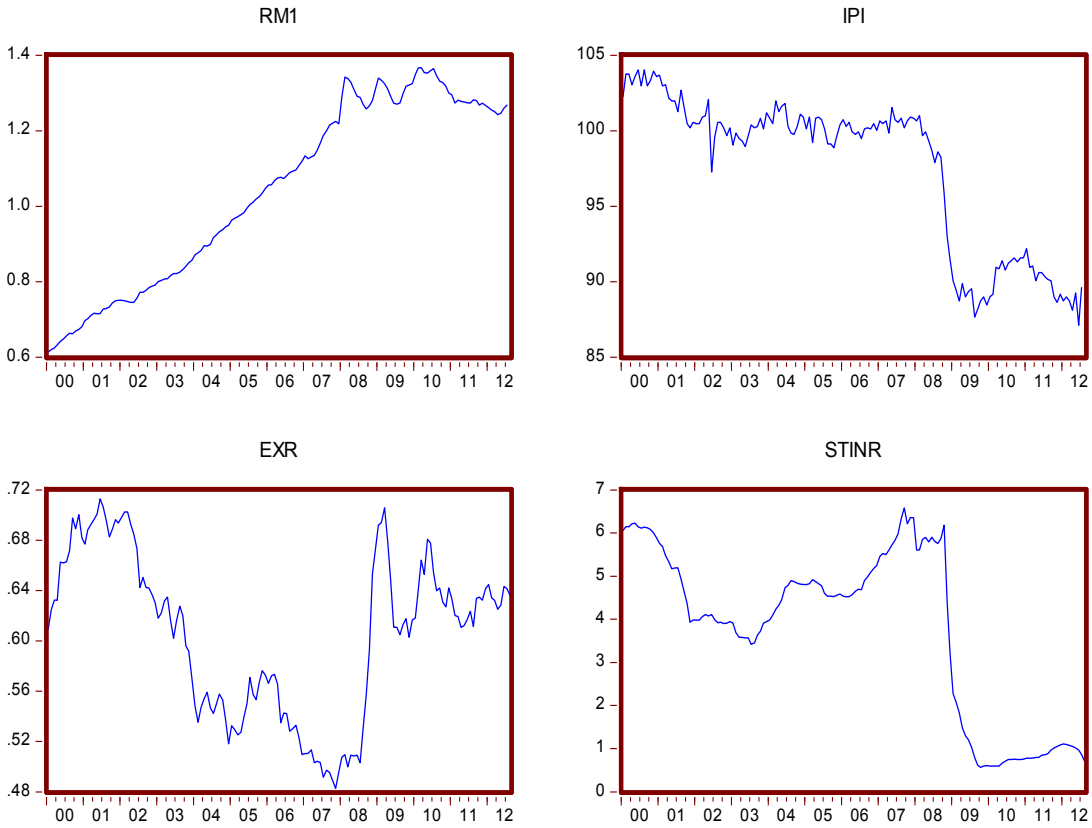

Figure 1. Time plots on Real Money Demand RM1_UK, IPI_UK, EXR_UK, INR_UK

The data on UK's Narrow Money Supply M1, Industrial Production Index, Exchange rates and Consumer Price Index were sourced from the database of Organization for Economic Cooperation and Development (OECD)[1]. While data on Euro Area's Narrow Money Supply (M1) and CPI were obtained from European Central Bank Database. Monthly data from 2000 to 2012 to build two vector autoregressive models, one for each economy was used. All data were transformed to log so that they can have same magnitude and to improve the data analysis.

DATA ANALYSIS AND RESULTS

The variables included in the analysis are short-term interest rates, exchange rates, narrow money supply, industrial production index and consumer price index. Financial crisis was introduced as a dummy variable to capture structural breaks in the models especially due to the global recession. All macroeconomic variables were converted to log. Narrow Money Supply M1 was deflated by CPI using the formula, *RM1= M1/CPI* to obtain the Real Money Demand, RM1_UK and RM1_EA in the United Kingdom and Euro Area respectively. Based on the time plots in Figure 1, we can assume random walks for all endogenous variables in this research work. To check the stationarity of our series, Augmented Dickey-Fuller unit root test (Table 1) was conducted on all the endogenous variables.

Augmented Dickey-Fuller Test (ADF)

Augmented Dickey-Fuller test (Said and Dickey, 1984) accommodates general Autoregressive Moving Average (ARMA (p, q)) models with unknown orders. The ADF tests, the null hypothesis showed that a time series y_t is I (1) against the alternative that it is I (0), assuming that the dynamics in the data have an ARMA structure. The ADF test is based on estimating the test regression:

$$Y_t = \beta' D_t + \Phi y_{t-1} + \sum_{j=1}^{p} \Psi_j \Delta y_{t-j} + \varepsilon_t \qquad 2.1$$

where Y_t and Δy_{t-j} represents level and first difference of each endogenous variable respectively, D_t is a vector of deterministic terms (constant, trend etc.). The *p*-lagged difference terms, Δy_{t-j}, are used to approximate the

[1] Organization for Economic Cooperation and Development was established in 1960 by European leaders with aim of encouraging cooperation and reconstructions after World War II. It currently spread across Europe, Americas, and Asia-Pacific regions with 34 memberships.

Table 1. Unit root tests results

United Kingdom	Level		First Difference	
	Without trend	With trend	Without trend	With trend
Lrm1_uk	-2.7064	0.0603	*-6.9358	*-7.6539
Lexr_uk	-1.7404	-1.6724	*-8.9426	*-8.9311
Linr_uk	-0.4665	-2.0943	*-4.8332	*-4.8892
Lipi_uk	-0.3848	-1.7308	*-15.3851	*-15.3421
Euro Area				
Lrm1_ea	-1.5351	-1.5548	-2.6486***	-2.9460

Values from ADF tests. *, *** represents no unit root at the first difference either at 1 per cent or 10 per cent level of significance respectively

ARMA structure of the errors, and the value of *p* is set so that the error ε_t is serially uncorrelated. The error term is also assumed to be homoskedastic. The specification of the deterministic terms depends on the assumed behaviour of y_t under the alternative hypothesis of trend stationarity (Said and Dickey, 1984). Under the null hypothesis, y_t is I (1) which implies that φ = 1. The ADF t-statistic and normalized bias statistic are based on the least squares estimates of (2.1) and are given by

$$ADF_t = t_{\varphi=1} = (\hat{\varphi}-1)/SE(\hat{\varphi}) \quad 2.2$$

$$ADF_n = T(\hat{\varphi}-1)/(1-\hat{\Psi}_1-\ldots\ldots\ldots-\hat{\Psi}_p) \quad 2.3$$

Or alternatively, the ADF formulation is

$$Y_t = \beta' D_t + \pi y_{t-1} + \sum_{j=1}^{p} \Psi_j \Delta y_{t-j} + \varepsilon_t \quad 2.4$$

where π = φ − 1. Under the null hypothesis, Δy_t is I(0) which implies that π = 0. The ADF t-statistic is then the usual t-statistic for testing π = 0 and the ADF normalized bias statistic is $T\hat{\pi}/(1-\hat{\psi}_1-\cdots-\hat{\psi}_p)$. The test regression (2.4) is often used in practice because the ADF t-statistic is the usual t-statistic reported for testing the significance of the coefficient y_{t-1} (Said and Dickey, 1984). Many studies including Skrabic and Tomic-Plazibat (2009) have shown that economic variables behave like random walks or at least have random walk components by using unit roots tests such as ADF. As shown in Table 1. The above table reveals that after considering with and without trend, the unit root tests accepted the null hypothesis of unit root. Therefore, a further ADF and PP tests on the first differences concluded that all variables are integrated of order one that is, I (1).

2008 Global Financial Crisis Effects

From the time plots (Figure 1), it can be observed that United Kingdom experienced an increase in the real money stock over the period under consideration. Economic growth as measured by industrial production has dropped significantly in the United Kingdom especially since the periods after 2008 financial crisis. The exchange rates have stabilized in the UK after recoveries from their low figures in 2008 (it rose sharply in early 2009 and has almost stabilised to its early 2000 figures). The financial crises lead to significant drop in the short-term interest rates, in an effort to stimulate and feed economic expansion, the Bank of England intervened by consistently reducing the prime lending rates. The “credit crunch” and market liquidity made the economic down-turn more protracted. Cash became “king” as investors avoided a variety of risky assets. Several financial corporations filed for bankruptcy in the United States, notably are Lehmann Brothers, IndyMac Bank, Merrill Lynch and the purchase of banking assets of Washington Mutual by JP Morgan Chase (ostensibly the biggest bank failure). Also, the insurance giant American International Group (AIG) sought an abridged loan ($US85 billion rescue package) from the Federal Reserve. Furthermore, a syndicate of 10 banks created an emergency fund of at least ($US70 billion) following the demise of Lehmann Brothers. In the United Kingdom, the government bailed out Northern Rock through nationalisation after un-successful take-over bids. Spanish Group Santander Bank bought Bradley and Bingley after its nationalisation in late 2008. Similarly, UK government acquired a major stake (about 84%) through partial nationalisation of the Royal Bank of Scotland Group in 2009. Mortgage Bank like Halifax Bank of Scotland, UK largest mortgage lender was merged with Lloyds TSB Group and the UK government took a 43.4% ownership in the combined group. Furthermore, in an attempt for safe-haven, most euro area banks especially from Central and Eastern European countries suffered significant capital flights inform of outflows of cross-border interbank deposits, mainly as non-affiliated depositors withdrew. There was significant drop in venture capital funding which generally results in slowed job creation and rise in unemployment rate. Below, potential growth impacted negatively on the labour force by steadily increasing the unemployment rates. There has been marginal drop in unemployment rates since the beginning of 2012 in the United Kingdom.

Table 2. Chow breakpoint test: Nov. 2007.

H_0: No breaks at specified breakpoint		
Test	**Value**	***P*-Value**
F-*statistic* : F(10,128)	1.6475	0.1004
Log likelihood ratio Chi-Square(10)	17.9197	***0.0563
Wald Statistic Chi-Square(10)	16.4755	***0.0868

Null hypothesis that there is no breaks at specified breakpoints can be rejected ***10% level of significant

This is attributable to significant downward pressure on the growth of labour earnings in the UK. On the other hand, consumer price index has shown steady rise during the period under review. The rise in inflation is partly responsible for the lower labour costs (lower-pay rises whereby companies are seen to be "hoarding labour" by retaining highly skilled staff to keep training cost down) particularly in the United Kingdom.

In 2009, there was a spike in the financial market activities in both economies. These were partly due to combination of news and improved companies earnings reports. This development was short-lived as investors failed to distract from worries about the economies. European Union banks witnessed increased reduction in assets through deleveraging Global Financial Stability Report (GFSR, 2012). Also, there was increasing market fragmentations and financial repression, which threatened the unified monetary policy of the euro area. In 2011, the ECB introduced a special scheme called the Long-term Refinancing Operations to boost the economies in the area. As a consequence of deteriorating economic conditions, persistent global financial turmoil especially in the Euro Zone, the money demand have experienced slow growth rate, which has impacted negatively on the United Kingdom economy. Furthermore, between May 2011 and July 2012, the European Union introduced some temporary and permanent financial assistance mechanisms such as the European Financial Stability Facility (EFSF), European Financial Stability Mechanism (EFSM) and European Stability Mechanism (ESM). These are measures geared towards ensuring good economic governance and fiscal discipline amongst member countries.

Presence of structural changes induced by 2008 Global Financial Crisis

The financial crisis of 2008 was included in the analysis, as a measure of the structural breaks observed in the series. This is essential because when there are breaks in the data, the regular ADF test tends to discover unit roots (non stationarity) that are inexistent. Structural change may occur for many reasons. The European integration has resulted in structural change in location, regional trade, regional fiscal coordination and economic governance. It could also occur by accident like the collapse of Lehmann Brothers and other financial institutions in late 2008. According to IMF World Economic Outlook (1998), crises may be considered to be an outcome of financial disturbances when markets suffer from a high degree of susceptibilities. These susceptibilities factors could be loss of confidence in banking system, sharp decline in assets and failure of financial institutions and financial corporations and so on. Chow Breakpoint tests (Table 2) were carried out on UK model respectively to ascertain where impacts of the global financial crisis were initially felt. After which an appropriate dummy variable was set up in the model reflecting this date. Chow breakpoint test involve comparing results of three tests statistic F-Statistic, log likelihood ratio and Wald Statistic. We tested whether there is structural change in the series before and during the 2008 financial crisis. Therefore, November 2007 was set as the breakpoint. The results of the three tests are as shown in the Table 2 below.

Test of parameter constancy

The reparameterized model is;

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Phi_i * \Delta y_{t-i} + AD_i + \varepsilon_t \qquad 2.5$$

where $\Pi = \alpha\beta'$, AD_i is the deterministic trend term which either equal to zero or not equal to zero. The dummy variable is as specified below:

Financial crisis = { 0, for t= 01/2000……..10/2007 ; 1 for t= 11/2007,……..08/2012

Dummy variables are sometime referred to as indicator variables whose presence in a model may remove the impacts of outliers or in this case residuals exceeding about 2σ in absolute values or 95% confidence interval.

Table 3. Granger causality test using Toda-Yamamoto procedure.

Dependent variable: L(RM1_UK)			
	X^2	df	Prob.
L(IPI_UK)	5.2495	2	**0.0725
L(INR_UK)	6.5135	2	** 0.0385
L(EXR_UK)	1.3970	2	0.4973
All	14.7048	6	**0.0227
Dependent variable: L(IPI_UK)			
L(RM1_UK)	2.9582	2	0.2278
L(INR_UK)	0.3639	2	0.8336
L(EXR_UK)	11.3172	2	* 0.0035
All	14.6450	6	** 0.0232
Dependent variable: L(INR_UK)			
L(RM1_UK)	2.3171	2	0.3139
L(IPI_UK)	18.1382	2	* 0.0001
L(EXR_UK)	3.0881	2	0.2135
All	26.1797	6	* 0.0002
Dependent variable: L(EXR_UK)			
L(RM1_UK)	0.6961	2	0.7061
L(IPI_UK)	7.2401	2	**0.0268
L(STINR_UK)	1.2625	2	0.5319
All	7.8436	6	0.2498
Significant at *1%, **5% level of significant			

By their inclusion we may obtain a better estimate of the innovation/shock variance. Doornik et al (1998, pp. 550) suggested three ways of including “impulse” dummies. “They could be ignored, introduced unrestrictedly, restricted to the cointegration space or a mixture of the last two”. However, they emphasized that the size of their effect matters sometimes when we have sample size increasing asymptotically after several Monte Carlo simulations.

Modelling real money demand in the United Kingdom

In order to avoid the situation of modelling a spurious regression amongst the endogenous variables, and the loss of long-run relationship usually associated with VAR*(p)* of random walks using the first difference, vector error correction models (VECM) was developed for the four endogenous variables using global financial crisis as an exogenous variables.

Test of granger non-causality

Granger non-causality test using the alternative procedure by Toda-Yamamoto (1995) was carried out to ascertain the causal relationship amongst the endogenous variables. This procedure was adopted because the use of Wald test statistic of linear restrictions on parameters of a VAR model where some of the series are non-stationary will not follow the usual asymptotic chi-square distribution under the null hypothesis (Engle, 1984). This is because the test’s asymptotic distribution involves nuisance parameter which cannot be observed directly. In light of this reason, this research work adopted the method proposed by Toda and Yamamoto (1995). One tested for the absence of Granger causality by estimating the following VAR model:

$$Y_t = \gamma_0 + \gamma_1 Y_{t-1} + \ldots\ldots + \gamma_p Y_{t-p} + \varphi_1 X_{t-1} + \ldots\ldots + \varphi_p X_{tp} + \omega_t \quad 2.6$$

$$X_t = \upsilon_0 + \upsilon_1 X_{t-1} + \ldots\ldots + \upsilon_p X_{t-p} + \varphi_1 Y_{t-1} + \ldots\ldots + \psi_p Y_{t-p} + v_t \quad 2.7$$

Then, testing H_0: $\varphi_1 = \varphi_2 = \ldots\ldots = \varphi_p = 0$, against H_A: 'Not H_0', is a test that *X does not* Granger-cause *Y*. Similarly, testing H_0: $\varphi_1 = \varphi_2 = \ldots\ldots = \varphi_p = 0$, against H_A: 'Not H_0', is a test that *Y does not* Granger-cause *X*. In each case, a rejection of the null hypothesis implies there is Granger causality. In Summary, Granger non-causality test results show that there exist unidirectional causality from LIPI_UK to LRM1_UK, LINR_UK to LRM1_UK and not vice versa (Table 3). However, there is a reasonable evidence of granger causality from all the independent endogenous variables to LRM1_UK when considered together at 5 per cent significant level.

Cointegration analysis of non-stationary series

The result of Augmented Dickey-Fuller Test, Engle and Granger (1987) pointed out that a linear combination of two or more non-stationary series may be stationary. The stationary combination may be interpreted as the cointegration, or equilibrium relationship between the variables. Regressing one random walk against another can lead to spurious results that is, conventional significance tests will tend to indicate a relationship between the variables when in fact there is none. To avoid this we may run regression with the stationary variables. However, if the variables are non-stationary (random walks) but are cointegrated running a regression with the first difference variables may lose the long-run information as the first difference regression results is for short-run. If the random walks are found to be cointegrated the regression result with variables at level are non-spurious and it also measure the long-run relationship between the variables. Therefore, the vector error correction model (VECM) was performed to investigate the short-run relationship including the Granger Causality relationship. We considered the vector autoregressive process with Gaussian white noise defined by

$$Y_t = \sum_{i=1}^{p} \Phi_i Y_{t-i} + \varepsilon_t \quad (2.8)$$

$$\Phi(B) y_t = \varepsilon_t \quad (2.9)$$

where $\mathbf{y}_{-p+1}, \ldots, \mathbf{y}_0$, are fixed and the shock or innovation ε_t is a Gaussian white noise. Since the AR operator $\Phi(B)$ can be re-expressed as $\Phi(B) = \Phi^*(B)(1-B) + \Phi(1)B$ where $\Phi^*(B) = I_k - \sum_{i=0}^{p} \Phi_i * Bi$ with $\Phi_i^* = -\sum_{j=i+1}^{p} \Phi_j$, the vector error correction model is

$$\Phi^*(B)(1-B)y_t = \alpha\beta' y_{t-1} + \varepsilon_t \quad (3.0)$$

$$\Delta y_t = \alpha\beta' y_{t-1} + \sum_{i=1}^{p-1} \Phi_i * \Delta y_{t-i} + \varepsilon_t. \quad (3.1)$$

Where Δy_t represents the first difference of endogenous variables at time t. Furthermore, Δy_{t-1} represeents the first difference of exogenous variables at time t-1 (this otherwise referred to as the short-run variables. One impulse for the VECM (p) form is to consider the relation $\beta' y_t = c$ as defining the underlying economic relations and assume that the regressors react to the disequilibrium error $\beta' y_t - c$ through the adjustment coefficient α to restore equilibrium; that is, they satisfy the economic relations. The cointegrating vector β is sometimes called the *long-run parameters*. Considering we have a vector error correction model with a deterministic term. The deterministic term D_t contains a constant and a linear trend.

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Phi_i * \Delta y_{t-i} + AD_i + \varepsilon_t \quad (3.2)$$

where $\Pi = \alpha\beta'$. The alternative vector error correction representation considers the error correction term at lag t-p.

If the matrix Π has a full rank (r=k), all components of $\mathbf{y}_t$ are I(0). On the other hand, $\mathbf{y}_t$ are stationary in difference if rank(Π). When the rank of the matrix Π is $r < k$, there are k-r linear combinations that are non-stationary and r stationary cointegrating relations. The cointegration rank test determines the linearly independent columns of Π. Johansen (1991) and Johansen and Juselius (1990) proposed the cointegration rank test using the reduced rank regression. When there are deterministic cointe-grated relationships among variables, deterministic terms in the VAR (p) model will not be present in the VECM (p) form. On the other hand, if there are stochastic cointe-grated relationships, deterministic terms appear in the VECM (p) form via the error correction term or as an independent term in the VECM (p) form. In some cases, a linear combination of variables removes the stochastic trend(s), but not the deterministic trend, so there is need to account for a linear trend in the cointegration space. There are different specifications of deterministic trends. Johansen (1988) suggested two test statistics to test the null hypothesis that there are at most r cointegrating vectors. One of them is the likelihood ratio trace statistics and the other one is maximum eigenvalue statistics, to determine the presence of cointegration vectors in non-stationary time series. The trace statistics and maximum eigenvalue statistics are shown in equation (3.1) and (3.2) respectively

1. Trace Test

$$\lambda_{trace} = -(n-p)\sum_{i=r+1}^{k} \ln(1-\lambda_i) \quad (3.3)$$

2. Maximum Eigenvalue Test

$$\lambda_{max} = -(n-p)\ln(1-\lambda_{r+1}) \quad (3.4)$$

where n is the sample size, λ_i is the i^{th} largest canonical correlation between residuals from the n-dimensional processes and residual from the n-dimensional differentiate processes. After carrying out Johansen Test of cointegration on the four endogenous, one long-run relation was generated. The test of cointegration was done excluding an intercept in the VAR. This was to ensure the validity of the critical values of the test associated with Johansen Cointegration test. The error corrections term (Table 6) as this long-run relation is sometimes referred, was computed based on the trace and maximum eigenvalue as depicted in the Table 4 below. The trace test tests the null hypothesis of at most r cointegration vector against the alternative hypothesis of full rank cointegration vector, the null and alternative hypothesis of maximum eigenvalue statistics is to check the r cointegrating vectors against the alternative hypothesis of at least one cointegrating vectors. The tests could not reject the hypothesis that the rank (Π) is at most one in both cases. Toda (1994) in an experiment using limited stochastic simulation showed that both tests

Table 4. Unrestricted Johansen Cointegration Rank Test.

Trace				
Hypothesized	**Trace**		**1%**	
No. of CE(s)	**Eigenvalue**	**Statistic**	**Critical Value**	**Prob.****
	0.248700	86.36395	71.47921	0.0002
At most 1	0.148050	44.04324	49.36275	0.0384
At most 2	0.126546	20.32964	31.15385	0.2097
At most 3	0.002061	0.305326	16.55386	1.0000

Maximum Eigenvalue				
Hypothesized	**Max-Eigen**		**1 %**	
No. of CE(s)	**Eigenvalue**	**Statistic**	**Critical Value**	**Prob.****
None *	0.248700	42.32071	37.48696	0.0020
At most 1	0.148050	23.71360	30.83396	0.0926
At most 2	0.126546	20.02431	23.97534	0.0404
At most 3	0.002061	0.305326	16.55386	1.0000

Table 5. Lag order selection.

Lag	LogL	LR	FPE	AIC	SC	HQ
0	585.4621	NA	$3.65*10^{-09}$	-8.0764	-7.9106	-8.0090
1	1525.994	1802.138	$8.85*10^{-15}$	-21.0069	-20.5097	-20.8048
2	1580.845	102.0307	$5.14*10^{-15*}$	-21.5502*	-20.7215*	-21.2135*
3	1590.901	18.1433	$5.60*10^{-15}$	-21.4672	-20.3069	-20.9957
4	1597.965	12.3494	$6.36*10^{-15}$	-21.3422	-19.8504	-20.7360
5	1611.581	23.0422	$6.61*10^{-15}$	-21.3088	-19.4855	-20.5679

are similar but emphasized that if r_{Ho}=0, there is a significant difference. Lutkepohl et al. (2000), considering different deterministic terms, showed that powers of the two tests are similar. However, with small sample sizes, the trace test power performance is higher. As a result of these likelihood ratio tests, Johansen tests depend only on completely specified autoregressive process for levels of data series. It tends to find cointegration more often in finite sample than in the asymptotic distribution and is more sensitive to the misspecification of lag length than in the non-normality of the disturbances (Wen, 1995). Hence, particular emphasis was placed on the lag length selection (Table 5) and adequate use of diagnostic testing for the residuals was ensured to avoid over acceptance of cointegration (see model misspecification analysis subsection 2.5.3). Therefore, prior to the estimation of VECM with the accompanying cointegrating vector, optimal lag length of initial Vector Autoregressive (VAR) model was ascertained. Different information criteria were calculated for various lag lengths. After calculations based on different criteria, two lags was selected by the Final Prediction Error (FPE), Akaike Information Criterion (AIC), Schwarz Criterion (SC), Hannan-Quinn (HQ) methods (Table 5). One advantage of this approach is that it can be applied to set of variables containing possibly a mixture of I (0) and I (1) regressors.

The long-run analysis shows that the cointegrating relation or error correction term with coefficients or speed of adjustment to equilibrium as measured by the multiplier α = -0.02446 is significant at 1% level. This is an indication that one can expect the LRM1_UK to converge to its long-run equilibrium at a very slow rate so as to allow the short-run dynamics. Specifically, we expect the equilibrium to be achieved in about three years after shock of real money demand. In order to evaluate the long-run relations, the cointegrating vector was normalized on LRM1_UK. The result is as shown in Table 6 below. A global test value of 13.46 is significant at 1%, which is high implies that all the endogenous variables are important in forecasting real money demand in the United Kingdom when considered together. Surprisingly, negative and significant relationship exists between real money demand and industrial production index in the long-run. However, the coefficient indicates a high responsiveness of money demand to a unit change in industrial production in the United Kingdom. The normalized coefficients indicate that all are statistically

Table 6. Normalized cointegrating coefficients.

Cointegrating Eq:	L(RM1_UK (t-1))	L(IPI_UK(t-1))	L(INR_UK(t-1))	L(EXR_UK(t-1))	@TREND (OOM 01)
CointEq1	1	-6.546987	0.326513	0.891143	-0.00633
S.E		-1.37901	-0.06105	-0.2539	-0.0013
t-statistics		[-4.74760]	[5.34795]	[3.50985]	[-4.86591]

Table 7. VECM coefficients for United Kingdom real money demand.

	Coefficient	Std. Error	t-Statistic	Prob.
α	-0.02446	0.007097	-3.4457	*0.0007
D(LRM1_UKt-1)	0.2952	0.07562	3.9030	*0.0001
D(LIPI_UKt-1)	-0.2625	0.08091	-3.2443	*0.0015
D(LINR_UKt-1)	-0.03907	0.01388	-2.8161	*0.0056
D(LEXR_UKt-1)	0.08220	0.03012	2.7294	*0.0071
Constant	0.003754	0.001062	3.5357	*0.0005
FINANCIALCRISIS	-0.003195	0.001687	-1.8943	***0.0602
R-squared	*0.3625*			
Adjusted R-squared	*0.3355*			
F-statistic	*13.4550*	Durbin-Watson statistic		1.8665
Prob(F-statistic)	**0.000000*			

Null hypothesis that estimated coefficient is equal to 0 can be rejected at *1% or ***10% level of significant

significant in the long-run. This cointegration vector relates money demand positively to short-term interest rates and exchange rates with low inelasticity (less than one). This finding confutes Doornik et al (1998) claim that the contemporaneous relation between money demand and long-term interest rate is negative in the UK. Hence, it may be argued that while long-term interest rates have a negative contemporaneous relation, short-term interest rates have a probable positive relation with real money demand in the United Kingdom. This may be attributable to higher rate of growth of M1 relative to CPI (a key component of real money demand function) despite the lingering tight liquidity especially during the financial crisis period. However, a partial test on individual endogenous variable revealed that only the first lagged variables of all the independent variables are significant in the short-run. Hence a general to specific model with only lag 1 was postulated in the case of UK real money demand. Specifically, a 1% increase in the lagged variable $LIPI\_UK_{t-1}$ results in 0.2625% drops in LRM1_UK. While LRM1_UK declines by 0.039% for every 1% increase in $LINR\_UK_{t-1}$, it increases by 0.08219% in the case of 1% increase in $LEXR\_UK_{t-1}$. The decline in real money demand is line with the theoretical evidence. Between 2009 and 2010, the BoE loosened monetary policy through large-scale purchase of assets (quantitative easing), which lead to increase in broad money by about 8% (Bridges and Thomas, 2012). Also none of the variables are weakly exogenous. If any of the endogenous/independent variable(s) is (are) weakly exogenous then parameters of this (these) variable(s) will have marginal density function bearing no relation to the parameters that determine the conditional density function of the dependent variable that is, $f_j(y,x)=f_c(y|x, \beta_i).f_m(x| \Phi_r)$. The global financial crisis has a significant negative impact on the United Kingdom real money demand during the period. If the coefficient is significant either at 1, 5 or 10% level, it can be concluded that the crisis has important impact on the real money demand. This is identified by a marginal coefficient of -0.003195 (Table 7). The VECM allows for the findings that the other endogenous variables Granger-Causes LRM1_UK or vice-versa as long as the error correction terms are statistically significant irrespective of the joint significance of the estimated coefficients.

Model misspecification analysis – Real money demand for United Kingdom

A Durbin-Watson value of 1.8665 indicates no serial correlation in the VECM system error term and confirms long-run relationships that exist between the endogenous variables. One of the major problems associated with the Johansen test of cointegration is the insensitivity to the non-normality of residuals/innovations. Therefore, in order to ensure the avoidance of over acceptance of cointegration, residual diagnostics were conducted for serial correlations, normality, ARCH effect and Heteroske-

dasticity. ARCH effect (*obs* R squared* = 0.3274, *p*-value = 0.2060) are insignificant at 10% level. After conducting the Breusch-Godfrey LM test of serial correlation (*obs*R squared* = 3.05, *p*-value = 0.2166) on the residuals one could not reject the null hypothesis of no serial correlation. Jarcque-Berra value of 1035.878 (*p*-value of 0.000) indicates the residuals are not multivariate normally distributed probably due to some remaining outliers otherwise the system is consistent with available evidence. The evaluation of the historical simulations using Theil inequality coefficient shows that a value of 0.01334 is close to zero, covariance accounted for 93.85%, variance 4.2% while bias proportion is 1.93% indicating a strong correlation between the actual and forecasted values. Forecast errors which are largest when most are happening in the economy usually reflect external shocks. In the case of real money demand in the UK, these shocks were observed mostly at the beginning of 1st and end of 2nd quarters of 2001, 2002, 2003, 2004, 2008, 2009 and 2010.

Impulse response function

The dynamic behaviour of the initial VECM model by studying impulse response function of money demand to Cholesky one standard deviation innovation or shock from independent variables was conducted. Impulse Response Function *(IRF)* helps to determine how each endogenous variable responds over time to shock in that variable and in every other endogenous variable by tracing the response of endogenous variables to such shocks. It allows one to identify shocks with specific endogenous variables so in order to ascertain how an unexpected change in one variable affects all variables over time. Therefore, an impulse response function shows the interaction between/among the endogenous variables sequence. Impulse response function (IRF) of a dynamic system is its output when presented with a brief input signal, called an impulse. More generally, IRF refers to the reaction of any dynamic system in response to some external change. A VAR can be written in the form of vector moving average (∞)

$$y_t = \mu + a_t + \varphi_1 a_{t-1} + \varphi_2 a_{t-2} + \ldots = \mu + \varphi(B)a_t \quad 3.6$$

$$\text{Where } \mu = E(y_t) = \psi_0/(1 - \psi_1 - \psi_2 - \ldots - \psi_p) \quad 3.7$$

The expression in equation 4.2 shows explicitly the impact of past shock a_{t-i} ($i > 0$) on the current y_t. The φ_i's are known as the impulse response function of the model. If a series is weakly stationary the φ_i coefficients decline exponentially. Below is a condensed form of the procedure for the computation of IRF.

Procedure:

1. To be able to compute the IRF, the model has to be in equilibrium. This can be achieved by holding the exogenous variable constant and allowing simulation over a long period of time so that the endogenous variables stop changing.
2. Introduce a one standard deviation shock to one of the endogenous variables say real money demand, ε_1 at time t = 0. This one period shock is what is referred to as the 'impulse'. This impulse will filter through the model affecting all the variables.
3. Then introduce one period shock to the next endogenous variable and so on until the last variable.

One important use of this type of analysis is that if the variables are cointegrated that is, move together in the long-run, effects of a temporary shock tend to dissipate after several years rather than been permanent. The IRF was calculated by increasing for one month only, the error terms in the four system equations of our VECM by one standard deviation and then calculate the immediate effect and future effects of this change on LRM1_UK. The impulse response function (IRF) was computed using the covariance matrix $\sum_{4x4}$ among the four error terms ε_{lrm1_uk}, ε_{lipi_uk}, ε_{linr_uk}, ε_{lexr_uk}. Please note that these error terms represent shocks from LRM1_UK, LIPI_UK, LINR_UK, and LEXR_UK respectively.

Variance-covariance matrix of shocks on VECM System

The matrix of variance-covariance above shows the response of LRM1_UK to one standard deviation disturbance on LRM1_UK, LIPI_UK, LINR_UK, and LEXR_UK that is, a one period increase of 0.008087, 0.008950, 0.04460 and 0.02161 respectively. An initial effect on LRM1_UK was concentrated largely on LRM1_UK. Shocks on the LRM1_UK had positive effect on LRM1_UK throughout the 24-month periods. Increasing the standard error term ε_{lipi_uk} and ε_{linr_uk} by 0.008950 and 0.04460 respectively produced negative reaction from LRM1_UK. Shocks on LEXR_UK resulted in in a positive response from real money demand in the second period. While the effect of shock on exchange rate produced a negative response from real money demand from period-10, one standard deviation innovation on industrial production index resulted in positive response from money demand in period-7. However, responses from shocks on industrial production index and exchange rates were the same in period-8(equilibrium period). Please note that short-term interest rates will produce the highest response from real exchange rate during the forecast period. As shown in Figure 2.

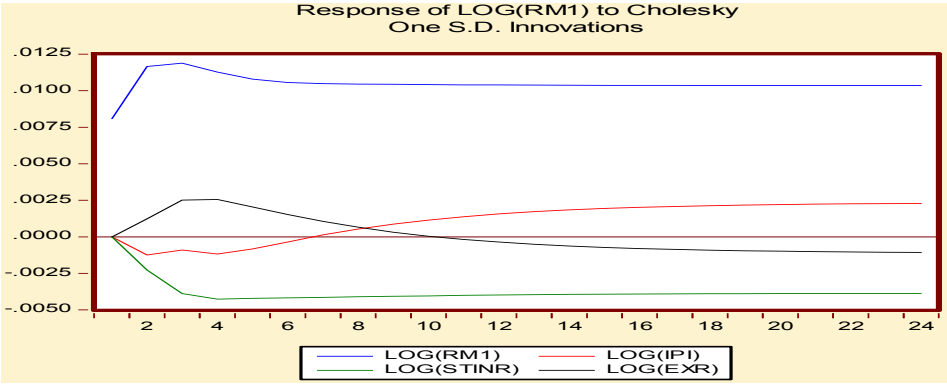

Figure 2. Response of LRM1_UK to One Standard Deviation Innovation.

Table 8. Granger causality between RM1_UK and RM1_EA.

Dependent variable: LOG(RMD_Uk)	Chi-sq	df	Prob.
LRMI_EA	12.07	5	0.034
AII	12.07	5	0.034

Dependent variable: LRMD _EA	Chi-sq	df	Prob
LOG(RMI_UK)	35.987	5	0.000
All	35.987	5	0.000

Relationships between real money demand in the United Kingdom and Euro Area

In order to study the relationships between the Real Money Demand in UK and Euro Area, we remodelled the Vector Error Correction Model system equation on United Kingdom's real money demand introducing the Euro Area's real money demand as additional exogenous variable. This type of analysis is essential considering the fact that United Kingdom is not committed to the "third EMU stage" which stipulates conditions necessary to adopt the euro currency[2]. Firstly, a test of granger non-causality (Table 8) between real money demand in United Kingdom and Euro Area was conducted using Toda-Yamamoto procedure (1995) as usual. Although, both variables granger causes each other, the chi-square values signifies that the past and current values of United Kingdom real money demand (with *chi-squared-value* 35.99, *p-value* = 0.0000) is highly significant in forecasting the future values of Euro Area real money demand than does the latter (*chi-squared-value* of 12.07 and *p-value* of 0.034).

[2] EMU is the Economic and Monetary Union which was established by the European Council in the Maastricht at the end of 1991 for the integration of European Union economies through coordination of economic, fiscal and independent monetary policies and adoption of single currency (Euro currency).

Null hypothesis of no granger causality significant at 1% and 5%

Though the introduction of the first difference of LRM1_EA in the UK real money demand VECM system equation increased *R-squared* by 2.38% from 36.5% to 38.88% (adjusted R-squared 35.84%), its coefficient in the model is negative and statistically significant at all levels with a value of -0.1759 and *p-value* of 0.0020 (Table 9). Furthermore, its presence in the model had a negative effect on the speed of adjustment to equilibrium in the long-run. Specifically, the speed reduced from about three years to six years in the presence of Euro Area Real Money Demand.

CONCLUSION

In this research work, we employed vector error correction and cointegration techniques in order to estimate the responsiveness (elasticity) of money demand to macroeconomic variables in the United Kingdom between 2000 and 2012. Long-run co-movement amongst the endogenous variables was established with very low speed of adjustment from disequilibrium caused by shocks on the real money demand. The adjustment to equilibrium in the long-run is expected to take about three years in UK so as to allow short-run dynamics. Real money demand in the long-run depended on industrial

Table 9. VECM Short-run coefficients for United Kingdom real money demand.

	Coefficient	Std. Error	t-Statistic	Prob.
α	-0.0140	0.0050	-2.7728	*0.0063
D(L(RM1_UK $_{(t-1)}$))	0.3405	0.0748	4.5529	*0.0000
D(L(IPI_UK $_{(t-1)}$))	-0.2336	0.0784	-2.9787	*0.0034
D(L(INR_UK $_{(t-1)}$))	-0.0369	0.0135	-2.7339	*0.0071
D(L(EXR_UK $_{(t-1)}$))	0.0811	0.0297	2.7289	*0.0072
Constant	0.0058	0.0010	5.7055	*0.0000
D(L(RM1_EA)	**-0.1759**	**0.0560**	**-3.1420**	***0.0020**
FINANCIALCRISIS	-0.0069	0.0016	-4.4210	*0.0000
R-squared	0.3888	Durbin-Watson statistic		1.85
Adjusted R-squared	0.3584			
F-statistic	9.6251			
Prob	*0.0000			

Null hypothesis that estimated coefficient is equal to zero can be rejected at *1% level of significant D(LOG(RM1)) = -0.0140*(L(RM1_UK(-1)) - 8.3299*L(IPI_UK (-1)) +0.4112*L(INR_UK (-1)) + 0.9906*L(EXR_UK (-1)) -0.0091*@TREND + 38.8767) + 0.3405*D(L(RM1_UK (-1))) - 0.2334*D(L(IPI_UK (-1)))-0.0369*D(L(INR_UK (-1)))+0.0811*D(L(EXR_UK (-1))) +0.0058-0.1759*D(L(RM1_EA)- 0.0069*FINANCIALCRISIS

production index, interest rates and exchange rates with relative elasticity. During the review period, the global test of the combined effects of all the endogenous and exogenous variables turned out to be significant in forecasting the UK real money demand. The impact of the 2008 global financial crises was evaluated. The effect of this dummy variable was negative and statistically significant in United Kingdom. This exogenous variable was introduced to evaluate the effect of inherent structural breaks in the economy observed especially from November 2007 in United Kingdom. Long-run dynamics showed a more than unity elasticity between industrial production index and real money demand. The short-run dynamics revealed that only the first lagged variables of the endogenous variables are statistically significant. Increases in industrial production index resulted in decline in the real money demand. This may probably be attributable to rising inflation rates or (and) low growth rate in narrow money supply M1 due to tight liquidity particularly during the financial crisis period. Increase in exchange rates resulted in increase in the real money demand. The economic theory plays strong role in determining the models' long-run and short-run properties, which are largely data-determined. Analysis of the relationship between the United Kingdom and Euro Area real money demand revealed that while there might be long-run relationships, real money demand in the Euro Area has a negative and statistically significant effect on the United Kingdom's real money demand. But there is currently debate on the future of its membership of the European Union. In 2012, more than half of British public and some members of the current Conservative Party led coalition are strongly in support of the UK leaving the European Union. Analysis of forecast error signified the impact of important shocks from external forces on real money demand. The global financial crisis, which began in November of 2007 as a result of significant downturn in the US economy, was revealed in the forecast errors of the model. Similarly, current financial crisis in the Euro Area was also accounted for in the forecast errors. However, each economy has peculiar shocks from external forces distorting the forecast errors at various points. These were attributable to tight market liquidity caused by the 2008 global financial crisis resulting in euro area banks suffering significant outflows of cross-border interbank deposits, Spain's announcement of austerity budget which resulted in highest inflation rate in 2 years, worsening unemployment rate in the Euro Area. Further analysis of the forecast error using impulse response indicates that shocks on interest rates will have the most effects on real money demand in the United Kingdom. Future increase in interest rates will likely explain decrease in money demand. The implication is that any increase in interest rates from their current levels will probably result in a significant reduction in the money demand in the United Kingdom in not too distant future. This is certainly a monetary policy concern for Bank of England.

Conflict of Interests

The authors have not declared any conflict of interests.

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