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REAL STOCK PRICES AND THE LONG-RUN MONEY DEMAND FUNCTION IN MALAYSIA:
Evidence from Error Correction Model

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This study adopts the error correction model to empirically investigate the role of real stock prices in the long run-money demand in the Malaysian financial or money market for the period 1977: Q1-1997: Q2. Specifically, an attempt is made to check whether the real narrow money (M1/P) is cointegrated with the selected variables like industrial production index (IPI), one-year T-*Bill rates (TB12), and real stock prices (RSP). If a cointegration between the variables, i.e., the dependent and independent variables, is found to be the case, it may imply that there exists a long-run co-movement among these variables in the Malaysian money market. From the empirical results it is found that the cointegration between money demand and real stock prices (RSP) is positive, implying that in the long run there is a positive association between real stock prices (RSP) and demand for real narrow money (M1/P). The policy implication that can be extracted from this study is that an increase in stock prices is likely to necessitate an expansionary monetary policy to prevent nominal income or inflation target from undershooting.*

Keywords: error correction model (ECM); real stock prices; money demand

Introduction

A study on direct relationship between stock prices and money demand functions is not very common.[1] It does not mean, however, that there was not any particular scheme of thoughts that had been used to explain such relationship. In fact, indirectly, Friedman (1988) has included stock prices in total non-human wealth,[2] and he empirically showed that the prices played significant roles in determining money demand function. On the other hand, money demand function, apart from being induced by stock prices, is also conjectured to be influence by other intimately but indirectly related factors. In particular, it is assumed that there exists an underlying stationary long-run equilibrium between real money balances, real income or real wealth and the opportunity of holding money balances. The last factor is generally represented by interest rates.

Friedman in the same article has further demonstrated that there are at least two effects of stock prices (wealth) movement on the demand for money, namely positive wealth effect and negative substitution effect. The first effect, which may positively affect wealth and in turn demand for money, works as follows: (i) any increase (decrease) in stock prices is an indication of increasing (decreasing) in nominal wealth; (ii) any increase (decrease) in stock prices is a reflection of rising (declining) in the expected returns from risky assets relative to safe assets. This will then induce the economic agents to hold larger amount of safer assets, such as money; and, (iii) any increase (decrease) in stock prices is a signal of inducing a rise (shrink) in the volume of financial transactions, which in turn will lead to higher money demand balances.

Meanwhile, the negative substitution effect of stock returns on money demand implies that as stock prices rise, the economic agents may preferably hold larger equities to other components of the portfolio, because the equities became more attractive or profitable (Thorton 1998, and Choudhry 1996). It is obvious from the discussion that the net effect of stock prices affect on money demand can be positive or negative, depending on which of the two effects is more dominant than the other. If the positive wealth effect of the stock prices dominates, then the higher the stock prices imply that the monetary authorities should foster monetary growth. Conversely, if the negative substitution effect dominates, the higher stock prices implicate the need for the monetary authorities to tighten monetary policy.

In the developed economies such as America, Japan and Germany, both positive and negative effects of stock prices on demand for money have been documented, therefore, had provided the supporting evidence to the theory of Friedman (1988). However, for the emerging markets, and Malaysia is no exception, to the best of our knowledge no empirical analysis has been done in this area. Thus, it is intriguing to investigate whether or not such relationship between stock prices and the demand for money does exist in Malaysia. For this reason, an attempt will be made by this paper to investigate the roles of stock prices in the long-run demand for money

[1] However, there are many studies on money demand alone, such as by Choudhry (1995), Jansen (1991), and Miller (1991).

[2] The studies on developed countries have included the volume of transaction or the return on securities as variables in the money demand function [see for example, Hamburger (1966); Keran (1971); McCornac (1991); etc.]. For a more detail discussion on this, refer Thorton (1998).

in Malaysia. Perhaps, by doing so we are able to compare the experience of Malaysia with that of the developed countries and eventually provide some reasons for the similarities and/or differences, if any. Furthermore, since the issue on the roles of stock returns and money demand are very crucial for policy makers, international fund managers, and other institutional investors who are seeking to diversify their portfolios in Malaysia's stock market, it is hoped that they will reap the benefits from the paper's findings.

Having said this, the objective of the present paper is three-fold: (a) to explore whether there exists a stationary long-run relationship between money demand and real stock prices in Malaysia; (b) to examine the size and direction of the effect of real stock prices on demand for money; and, (c) to appraise the temporal causality between the real money stock and the determinants of the long-run real money demand.

The remaining parts of the paper are constructed as follows. In Section 2, the data and methodology of cointegration and Granger causality on which the analysis is based are presented. Then, in Section 3, the reports of the statistical tests together with the results of empirical findings are displayed. The paper ends with a section on summary and conclusion.

Data and Model Specification

Data

In an attempt to provide a robust result, the study utilizes quarterly and seasonally unadjusted data[3] covering the period 1977: *Q1* to 1997: *Q2* (82 observations). The data gathered are from Bank Negara (the Central Bank of Malaysia) reports. Seven variables are tested in the model: (i) the Consumer Prices Index (CPI). It represents the price level (P); (ii) both measures of money, that is the narrow measure of money (M1) and broad measure of money (M2). They represent the nominal money balances. *M1* and *M2* are then deflated by price level (P) in order to provide real money balances. We note that *M1* and *M2* are treated in the present study as two different variables; (iii) the 3-month *T*-Bill rates (TB3) and one-year *T*-Bill rates (TB12). They are chosen as proxies for interest rates (opportunity cost of holding money). As in the case of *M1* and *M2*, *TB3* and *TB12* are also treated as two different variables;[4] (iv) for reason that the real *GDP/GNP* data for the study period are lacking, twenty-one years of quarterly Industrial Production Index (IPI) is used as a proxy for national output; and (v) the real stock prices (RSP) are calculated by dividing the Kuala Lumpur Composite Index (KLCI) by price levels. All variables are expressed in logarithmic terms.

Model Specification

Following a standard specification of the model, money demand function can be written as $(M/P)^d = f(y, i)$, where y is national output and i is interest rates, respectively. Meanwhile, a money demand function that includes stock prices can be simply formulated as follows:[5]

$$\boxed{(M/P)^d = \beta_0 + \beta_1 IPI_t + \beta_2 TB_t + \beta_3 RSP_t + \varepsilon_t \ldots\ldots\ldots\ldots\ldots\ldots\ldots (1)}$$

[3] This seasonal and unadjusted data are superior to the dynamic properties of the model (Ibrahim, 1998).

[4] In Malaysian economy, *T*-bill rate can be a good proxy for interest rate since it has been extensively used by Semudram (1981). In a most recent work, Ibrahim (1998) also used similar proxy.

[5] See for example, a model introduced by Choudhry (1996) which was then followed by Thorton (1998).

where
$(M/P)^d$ represents real money balances,
M = nominal money,
P = the prices level,
IPI = the Industrial Production Index,
TB = *T*-Bill rates,
RSP = stock prices, and
ε_t = the error random terms.

Theoretically, the demand for money function, $(M/P)^d$, is positively related to real income (proxied by IPI) and negatively related to the rate of interest, which in this paper is proxied by *T*-Bill rate. As indicated earlier, the rate of $(M/P)^d$ can be positively or negatively related to stock prices (SP).

Stationary and Cointegration Tests

Most macroeconomic variables are found to be non-stationary which has in turn resulted in a spurious regression (Sarletis 1992). To avoid this problem, the present paper conducts a unit root test in order to examine whether the variables are stationary or not. If the variables are non-stationary, they have to be differenced either by once or more until the stationary of the variables are achieved. The integration order of the variables shows how many times they have been differenced to be stationary.[6] Generally, most of the economic data will be stationary after taking once differencing.[7] To this end, the standard Dickey-Fuller (DF) and Phillips-Perron (PP) tests of unit root are applied in this paper.[8]

Once the order of integration is identified, a consideration needs to be given to the possible cointegration of the variables in the model. To be more precise, the objective is to know whether the variables have long-run equilibrium or inherent tendency to move together in the long run. The existence of cointegration among variables in the model indicates that there exists a long-run equilibrium. This information is extremely crucial, as it would suggest for specification of the correct model. If all variables under consideration are integrated of order one, I(1), but they are not cointegrated, the only valid regression model is to estimate the equation using the first difference of the variable (Engle and Granger 1991). If, instead, they found to be cointegrated, the error correction model (ECM) should be adopted to predict stock prices and the demand for money as it combines both short-run dynamics and long-run equilibrium conditions (Ibrahim 1997; 1998: 56-57).

Next, in the analysis of the cointegration of the variables, the two most commonly used tests, namely the residual-based Engle and Granger (1987) and the Johansen (1988) and the Johansen and Juselius (1990) approaches, are adopted to examine the hypothesis of a stationary long-run money demand function. Both tests, for brevity, are called EG and JJ tests, respectively.[9]

Granger Causality Tests

To evaluate the temporal causality between real stock prices and money demand, the Engle and Granger (EG) tests are adopted. The tests consider the possibility of the past variable level, for in-

[6] Integrated of order one, I(1) indicates that such variables have been differenced by once in order to achieve stationary.

[7] It has been recognized by many researchers, for example, Ibrahim's (1997) work on Indian economy.

[8] As it is well known and commonly adopted in recent empirical studies, these procedures or methods will not be explained in detail.

[9] To save space and due to the fact that it is commonly adopted in recent empirical studies, these procedures will not be explained in detail.

stance (Y), to explain the current changes in other variables (X), even though *Y* in the past did not change. This being the case and in view of the fact that the lagged values play such an important role in determining the dependent variable, they are included in the model. As stated earlier, if the variables are cointegrated, the error correction model (ECM) has to be applied in examining the variable relations. The error correction model (ECM) can simply be written as follows:

$$\Delta \ln(M/P)^d = \sum_{i=1}^{r} \phi_i \Delta \ln(M/P)^d_{t-1} + \sum_{i=1}^{s} \alpha_i \Delta \ln IPI_{t-1} + \sum_{i=1}^{k} \beta_1 \Delta \ln TB_{t-1} + \sum_{i=1}^{l} \delta_i RSP_{t-1} + \gamma_i EC_{t-1} + \nu_t \quad \text{..................................(2)}$$

where *EC* is the error correction terms or residual that is saved from the cointegrating regression equation (1). The *EC* terms are included in the model specification (Equation 2) to form an error correction representation of the model, which is based on the Engle and Granger's (1987) representation theorem. The application of error correction terms in the model is aimed at capturing the potential departure effects of the model's variables from the long-run equilibria. The size and significance of the error correction term in each equation implies the tendency of each variable to restore equilibrium in the money market (Choudhry 1996, and Thorton 1998).

We note at this juncture that the lagged values, if arbitrarily included to the lag length, may result in inefficiency or biased parameter estimates. For example, if the lag length is too large, the inclusion of irrelevant variables causes inefficiency of the estimated coefficients. By the same token, if it is too small, the estimated coefficients will be biased due to the omission of relevant variables.[10] To disentangle this problem, the Final Prediction Error (FPE) criterion is used to determine the lag lengths.[11] The present paper will, however, consider the lag lengths by examining across all possible lag combinations with the maximum lag lengths for each variable set up to 8.[12]

Finally, in order to judge the goodness of the chosen model, a battery of most common diagnostic tests are utilized. They are; (i) the RESET test. It is adopted to examine the model specification error; (ii) Jarque-Bera (JB) test is used for normality test; (iii) Durbin-Watson (DW) test. It is adopted to test the presence of autocorrelation; (iv) the HET test is used for testing heteroskedasticity; and finally, (v) the *CHOW* test is used to test for structural stability of the model.

[10] Another weakness of including arbitrary lag lenghts is that, it generally yields insignificant *F*-statistics. See, for example, Ibrahim (1999c) and Abd. Majid (2002).

[11] The smaller the *FPE* values, the better the model would be. Therefore, comparing with many possibilities of lag-length specification in the model, the smallest value of FPE of a model is considered and chosen as best-fit lag-length specification in the model.

[12] The main reason of choosing maximum of 8-lag-lengths is because of quarterly data analyzed in this paper. It is a commonly that for a quarterly data, the possible lag combination included in a model is set to 4, 6, and 8. To provide the most believable results, this paper considers the maximum possible lag combination, that is 8.

Stationary and Cointegration Test Results

Stationary Test Results

The results of the Dickey-Fuller (DF) and Phillips-Perron (PP) tests are reported in Table 1. In the test for the null-hypothesis where $\delta = 0$, it is found that unit root (nonstationary) exists. This is obvious from the non-rejection of the null-hypothesis because the critical values exceed the test statistics.[13] Specifically, when the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests for order of integration of each variable were conducted, with variable one-year *T*-bill rate (TB12) being the only exception, all other variables are nonstationary in log levels either with constant and no trend or with constant and trend regression models. Meanwhile, after taking first differences for Phillips-Perron (PP) test, all variables become stationary at 1 percent significance level or integrated of order 1, I(1). Using Augmented Dickey-Fuller (ADF) test, only 3-month *T*-Bill rate (TB3) and real stock prices (RSP), which is proxied by Kuala Lumpur Composite Index (KLCI), are stationary at 1 percent significance level either with constant and no trend or with constant and trend models. The real broad money demand (M2/P) is also stationary at 1 percent level of significance with constant and no trend model after taking the first differences, $I(1)$. However, based on Augmented Dickey-Fuller test (ADF) there exists the possibilities of being integrated of order 2, $I(2)$ for two variables, i.e., output (proxied by IPI) and real narrow money demand (M1/P). Since almost all variables are stationary in the first differenced series we conclude, therefore, that they are integrated of order 1, $I(1)$.[14]

The Cointegration Test Results

After examining the stationary properties of the data, the cointegration tests

Table 1. **Unit Root Test Results**

Variables	Log Levels				First Differences			
	ADF		PP		ADF		PP	
	Constant & No Trend	Constant & Trend	Constant & No Trend	Constant & Trend	Constant & No Trend	Constant & Trend	Constant & No Trend	Constant & Trend
M1/P	1.398	-1.079	0.812	-0.799	-2.213	-2.549	-9.803 *	-9.806 *
M2/P	0.931	-0.623	0.925	-0.662	-2.609 ***	-2.471	-8.706 *	-8.651 *
IPI	-1.711	-1.494	0.425	-2.366	-2.268	-2.678	-9.493 *	-9.456 *
TB3	-2.251	-2.451	-2.423	-2.670	-3.780 *	-3.769 *	-7.355 *	-7.302 *
TB12	-2.771 ***	-3.450 **	-2.127	-2.422	-	-	-8.264 *	-8.220 *
KLCI	-2.020	-2.778	-2.017	-2.812	-4.779 *	-4.767 *	-7.670 *	-7.838 *

Note: *, **, *** denote significance levels at 1 percent, 5 percent, and 10 percent, respectively.

[13] Many econometric textbooks explain this procedures of testing. For more detail, for example see Gujarati (1995).

[14] The paper mainly focuses on Phillips-Perron test for further specification of the models.

are now conducted.[15] The null-hypothesis of no cointegration is rejected if it is found that the residuals are stationary process. Recall that there are two alternative proxies used for real money demand; (M1/P) and (M2/P), and 3-month *T*-Bill (TB3) rates and one-year *T*-Bill rates (TB12); for interest rates. Thus, there are four cointegrating regression models that are subject to analysis and the results are reported in Table 2. Both Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests statistics either with constant and no trend or with constant and trend are tested. The Eigenvalues and Trace tests are reported for the test of cointegration using *JJ* test. As can be deduced from Table 2, the Engle and Granger test of Augmented Dickey-Fuller (ADF) test statistics tends to suggest that the hypothesis of having cointegration among the variables is rejected in each case. Meanwhile,

Table 2. **Cointegration Test Results**

Model	**Variables**	***EG*-Tests**		***JJ*-Test**		
		ADF	**PP**	**Lags in the VAR=4**		
		Statistics	**Statistics**	**Vectors**	**Trace**	**Eigenvalues**
1	(M1/P), IPI, TB3, RSP	(i) -1.330	(i) -4.155**	r = 0	45.062	21.059
				r £ 1	24.003	13.728
		(ii) -1.424	(ii) -4.387***	r £ 2	10.275	8.203
				r £ 3	2.072	2.072
2	(M1/P), IPI, TB12, RSP	(i) -1.355	(i) -4.150**	r = 0	48.364***	24.115
				r £ 1	24.250	14.287
		(ii) -1.460	(ii) -4.396***	r £ 2	9.963	7.737
				r £ 3	2.226	2.226
3	(M2/P), IPI, TB3, RSP	(i) -2.139	(i) -2.739	r = 0	44.340	16.647
				r £ 1	27.692	15.051
		(ii) -1.440	(ii) -1.468	r £ 2	12.642	7.993
				r £ 3	4.649	4.649
4	(M2/P), IPI, TB12, RSP	(i) -2.113	(i) -2.837	r = 0	45.143	17.693
				r £ 1	27.450	15.631
		(ii) -0.948	(ii) -1.638	r £ 2	11.819	6.935
				r £ 3	4.883	4.883

Note: *, *, *** denote significance levels at 1 percent, 5 percent, and 10 percent, respectively.
(i) and (ii) in both ADF and PP tests indicate the regression models with constant and no trend and constant and trend, respectively.
For the trace tests, the critical values at 10 percent and 5 percent significance levels for the null hypothesis that r = 0 and r £ 1 are 45.2 and 48.4, respectively. The corresponding critical values for the Eigenvalues tests are 24.9 and 27.3.

[15] In this test, the seasonal dummies were included to eliminate the seasonal effect.

based on Engle and Granger test of PP test statistics,[16] the *first model* [(M1/P) regressed on IPI, TB3, and RSP] and the *second model* [(M1/P) regressed on *IPI*, TB12, and RSP] for both constant without trend and constant with trend models are respectively found to be cointegrated at 5 percent and 10 percent level of significance.

The last three columns of Table 2 report the results of Model 2, [(M1/P regressed on IPI, TB12, and RSP], which are based on Trace test. They are cointegrated statistically with a single non-zero vector at 10 percent level of significance. However, Models 1, 3, and 4, [(M1/P) regressed on IPI, TB3, and RSP], [(M2/P) regressed on IPI, TB3, and RSP], and [(M2/P) regressed on IPI, TB12, and RSP], are found not to be cointegrated statistically. Specifically, the null hypothesis of non-stationary residual in Model 2 is rejected by the Trace test at 10 percent significance level when the order of cointegration is set to 4. Conversely, the maximum Eigenvalues test statistics shows evidence of noncointegration of the six variables in the four considered models (Models 1, 2, 3, and 4). It implies that the inclusion of real stock prices in the long-run demand for money function to be appropriate in the case of (M1/P). Indeed, this finding is actually similar to those of Choudhry (1996), for the case of Canada and USA, and Thorton (1998), for the Germany case. In our case, with one being the only exception, namely Model 2, the finding suggests that there is a long-run relationship among the variables, that is (M1/P), *IPI*, *TB*12, and *RSP*.

The Causality between Real Stock Prices and Money Demand Function

After identifying the cointegrating test for the variables, what remains to be the focus of the paper is to analyze the causality between these variables as discussed in Section "The Cointegration Test Result." The analysis includes the error correction terms. We note here, because real broad money (M2/P) and 3-month *T*-Bill rates (TB3) are considered as inappropriate long-run determinant of money demand function as evident by Models 1, 3, and 4 in Table 2, these variables are ignored for further analysis.

Table 3 presents the results of error correction estimations. In this table, the sum of coefficients of the lagged differences and some well-known diagnostic tests are reported in the first and second rows for each dependent variable. The *F*-test statistics, indicating the significance level of the sum of coefficients, are reported in the squared parentheses, while the error correction term(s), tested using *t*-test statistics, are reported in the parentheses.

The significance of real stock prices in the money demand function indicates that the real stock prices have significant role in determining demand for money in the Malaysian economy. Among the variables identified above, none of them is recorded has non-causality to each other. The entire variable signs for the model where (M1/P) regressed on *IPI*, *TB*12, and RSP are found to be consistent with our expectation. They are shown in the table

[16] The cointegration tests are also conducted for demand for money specification by excluding real stock prices. The results indicate that the hypothesis of a single cointegrating vector among the variables is rejected in all cases. The results, for reasons of conserving space, are no presented here.

by the positive signs of sum of output (IPI) and real stock prices (RSP) coefficients, and the negative signs of both sum of coefficients of real narrow money (M1/P) and one-year *T*-Bill rates (TB12).

As mentioned previously, the maximum lag-lengths considered are 8 for each error correction model. Based on Final Prediction Error (FPE) criteria, for dependent variables (M1/P) and IPI, the maximum lag-length specification included in the model are 2 and 5 respectively, while 8 lags are required for both models with the dependent variables TB12 and RSP.

Theoretically, negative signs are expected on the error correction term in the real narrow money (M1/P) and one-year *T*-Bill rates (TB12) equations, while positive sign is expected on the error correction term in the real output (IPI) equation.

Table 3. **Error Correction Estimation**

Dependent Variables	ΣΔ(M1/P)	ΣΔIPI	ΣΔTB12	ΣΔRSP	ECT_{t-1}	R^2-Adj
ΣΔ(M1/P)	-0.1758**	0.4273*	-0.0168	0.0132	0.0113	.1516
	[2.3690]	[6.2919]	[0.8049]	[2.5984]	(0.7623)	
	{2}	{2}	{2}	{2}		
	RESET (2)= 0.0819		RESET(3)= 0.1065		RESET(4)= 0.0744	
	JB= 36.4834*	DW= 1.7595	HET(9)= 16.6560		CHOW(10,58)= 2.6336	
ΣΔIPI	0.8535*	-1.3730*	-0.0072***	-0.0088	0.0492*	0.6574
	[2.8808]	[12.6010]	[1.6602]	[0.2552]	(3.9700)	
	{5}	{5}	{5}	{5}		
	RESET (2)= 1.4129		RESET(3)= 2.5087		RESET(4)= 2.2711	
	JB= 0.5116	DW= 1.6915	HET(21)= 25.7190		CHOW(22,31)= 2.2433	
ΣΔTB12	-0.4999**	5.6436	-0.4465*	0.1813**	0.0901**	0.3633
	[1.8188]	[0.6856]	[5.2861]	[2.0127]	(1.8230)	
	{8}	{8}	{8}	{8}		
	RESET (2)= 1.9134		RESET(3)= 17.2450*		RESET(4)= 11.5690*	
	JB= 4.0872	DW= 1.8318	HET(33)= 44.3610		CHOW(34,4)= 1.3495	
ΣΔRSP	-7.8304*	11.6360**	-0.4087**	-0.6551*	0.5246*	0.4585
	[2.8377]	[1.9798]	[2.0973]	[5.1021]	(3.7410)	
	{8}	{8}	{8}	{8}		
	RESET (2)= 2.1771		RESET(3)= 1.3480		RESET(4)= 1.2341	
	JB= 18.0418*	DW= 1.8069	HET(33)= 53.3820		CHOW(34,4)= 0.2618	

Note: *, **, *** denote the levels of significance of 1 percent, 5 percent, and 10 percent.
The numbers in [.] and (.) are the *F*-statistics and *t*-statistics, respectively, are used for testing the null hypothesis that the estimated and coefficients' sum are equal to zero. The {.} is the optimal lag-length included in the model based on the Akaike's (1969) Final Prediction Error (FPE) criteria. *JB* is Jarque-Bera test for normality, *RESET* is Ramsey's test for functional misspecification, *DW* is Durbin-Watson d test for autocorrelation, *CHOW* is Chow test for structural stability, and *HET* is Breusch-Pagan-Godfrey test for heteroskedasticity.

The error correction term in real stock prices (RSP) model can be either positive or negative. These signs are expected because excess supply of money will result in an increase of real output and a decrease in the interest rate. On the other hand, if the purpose of monetary policy is stabilization, then the growth rate of money stock should decrease (Choudhry 1996).

From Table 3, except for the first coefficient estimates for error correction terms in the (M1/P) equation, all the remaining coefficient estimates for error correction terms are positively significance at least at 5 percent significance level. The result from error correction estimations indicates that both goods and financial markets in Malaysia may have adjusted to the disequilibrium in the money market. Except for one-year *T*-Bill rates (TB12), the signs of error term for IPI and RSP equations behaved as expected.

In general, the results suggest a bi-directional causality between the variable. On causality between the real money (M1/P) and real stock prices (RSP), the bi-directional effect is found from both (M1/P) to RSP, and RSP to (M1/P). The directional effects are only found from TB12 to IPI, TB12 to (M1/P) and from RSP to IPI.

Finally, Table 3 also reports some diagnostic tests for model specification. RESET test used for testing specification error indicates that, except for the third model [TB12 regressed on (M1/P), IPI, and RSP] for RESET (3) and RESET (4), other models have no specification error. In addition, Jarque-Bera (JB) test is used for normality test. Unlike Model 2 and 3, both Models 1 and 4 failed to pass the JB normality tests. The Durbin-Watson (DW) *d* test indicates that all models with four different possible dependent variables recorded no autocorrelation. Like DW *d* test, HET test results also imply that all models do not suffer from heteroskedasticity problem. Finally, CHOW test, which is used to test for structural stability, tends to suggest that our chosen model is good enough. Therefore, there is not a need for data break analysis. Overall, the performance of forecasting equations is very satisfactory. They have respectably passed and fitted a series of diagnostic tests.

Conclusion

In examining the role of real stock prices in the long run money demand in the Malaysian market for the period 1977: Q1-1997: Q2, the present paper employs three independent variables, Industrial Production Index (IPI), *T*-Bill Rates (TB) and Real Stock Prices (RSP) with two possibilities of dependent variables, real narrow money (M1/P) and real broad money (M2/P) and also two alternative proxies used for interest rate; 3-month *T*-Bill Rates (TB3) and one-year *T*-Bill Rates (TB12). The (M1/P) is found to be cointegrated with IPI, TB12, and RSP. This implies that there exists a long run co-movement among these variables in the Malaysian market. Only for the cointegrated variables, the paper mainly focuses its analysis by applying the error correction terms in the model. The results indicate that in the long run the real stock prices (RSP) have positive association with demand for real narrow money (M1/P). An increase in stock prices is likely to necessitate an easier monetary policy to prevent a given nominal income or inflation target being undershot. Real narrow money (M1/P) and real stock prices (RSP) recorded to have bi-directional causality. Finally, the chosen model has been proven good and efficient, as it has passed a battery of diagnostic tests.

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