Testing Transmission Mechanisms on Economic Growth in Malaysia

Wai Ching Poon

Abstract
This paper examines various transmission mechanism channels on economic growth in Malaysia over the quarterly period 1980:1-2004:4 using bounds testing approach. The bounds test reveals evidence of cointegration between the real GDP and the real exchange rate and share prices that address the exchange rate and asset price channels as the key transmission mechanisms in the conduct of the monetary policy stance. Nevertheless, the saving interest rate and credit channels are of insignificant.

Keywords: cointegration, transmission mechanism channels, monetary policy stance, bounds test, Malaysia

JEL classification: E52

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I. Introduction
Monetary policy affects the economy and price level through multiple channels of monetary transmission mechanisms. The ultimate goal of monetary policy is to help achieving its maximum sustainable growth, keeping inflation low, stable and predictable. After years of structural change on the exchange rate regime and the fast pace evolving financial liberalization, we now examine if there are other transmission mechanisms contribute to the economic growth in Malaysia apart from the two price indicators of monetary policy stance, i.e., interest rates and the exchange rate. Economic growths have been heavily documented in the mainstream literature both theoretically and empirically.

Customarily, there are a few possible monetary transmission channels identify in the literature by which monetary policy affects aggregate demand (AD) and inflation (see for example, Mishkin, 1995, Monetary Policy Committee, 1999, Kuttner and Mosser, 2002, Arestis, 2002), namely the interest rate channel, the exchange rate channel, the asset prices channel, and credit channel that comprising two sub-channels - the narrow credit channel and the broad credit channel.

Higher interest rate translates into a decline in capital accumulation from both investors and households, which in turn lower total output and depress the economic. Tight monetary policy stance lowers AD and thus reduces inflation and improves balance of payments. Furthermore, the cost associated with high interest rates to stabilize the currency could be overwhelmed if the banking sector was fragile. However, if the corporate sector was heavily exposed to foreign debt, increase interest rates might be the appropriate policy.

On the other hand, stronger currency would affect negative wealth and income effect domestically. However, the substitution effect may become dominant gradually and lead to greater import penetration to the country thereafter. Empirically, Goldfajn and Baig (1998) evaluated the relationship between monetary policy and the exchange rate in five Asian crisis countries, namely Indonesia, Malaysia, the Philippines, Korea and Thailand. Results showed that there was no evidence of overly tight monetary policy in the Asian crisis countries in 1997 and early 1998. In fact, real interest rates in Indonesia, Malaysia, and Philippines were below their pre-crisis levels. Hence, no evidence indicates high interest rates leads to weaker exchange rates. There was also no evidence of large uncovered interest rate differentials in the Asian crisis countries. As such, they suggested the need to reconcile the traditional interest rate-exchange rate trade-off with a corporate balance sheet approach.

The widespread globalization of financial markets and the evolving liberalization of the economy during the 1980s have increased the interest of Central Banks in asset price development. A change in asset prices affects the financial wealth and changes their consumption decisions. When stock prices increase, the net wealth increases, and wealth effect induces a change in household consumption. Empirically, Pichette and Tremblay

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2 See Freedman (1995, 1996) for the presentation of evolution from interest rate channel to exchange rate channel. While, Mishkin (1996) heavily discussed the credit channel, and Hall (2001) for the discussion of narrow and broad credit channels.
Transmission mechanisms in Malaysia

(2003), using vector error correction model, showed a rather weak wealth effect on the stock market fluctuation, but found stronger evidence of housing wealth effect for Canada. Using a backward-looking IS-Phillips curve model, Goodhart and Hofmann (2001) concluded that the index weights for both property and share prices significantly impact the AD for G7 countries; with the effect of housing prices outweigh that of stock prices. However, many studies found that stock returns possess little predictive content for future output (Gauthier et al., 2004; Fama, 1981; Harvey, 1989; Stock and Watson, 1989 & 1999; Estrella and Mishkin, 1998). This has raised the question of re-examine the role of financial macroeconomic variables in the conduct of monetary policy stance.

Narrow credit channel (also referred as bank lending channel), and broad credit channel (also labeled as balance sheet channel) are concerned with how changes in the financial positions of lenders and borrowers affect AD in the economy. The narrow credit channel concentrates on the role of banks as lenders (Bernanke and Blinder, 1988), where banks rely heavily on the demand deposits. Given the significant number of firms and households depend on bank lending; ultimately AD and inflation would be affected. This channel is significant if increases in interest rates lead to a reduction in the supply of bank loans (Hall, 2001). Meanwhile, the broad credit channel relies on the supply of external finance to firms, where the lenders charge borrowers a premium to cover monitoring costs. Policy that increases interest rates will lead the borrower’s premium increase, thus affect the value of asset prices and the real value of consumer wealth, as well as declines the value of the collateral. Because of imperfect information in the credit market, the credit stance of banks has important supplementary effects on monetary policy transmission (Bernanke and Gertler, 1995). From the bank lending perspective, changes in interest rate affect the net worth of firms. A reduction in the statutory reserve ratio would increase the loanable fund to the public. High interest rates tend to lower the net value of companies and reduce borrowing, especially if moral hazard persists (Korhonen, 2002). In contrast, an increase in the supply of credit and a fall in interest rate will lead to a drop in the cost of credit. Hence, improves firms borrowing and increases investment spending generally, and improves the balance sheet of the firms particularly.

According to Duguay (1996, p.1), whichever channel is selected to be the monetary transmission mechanism channel, it supports the notion that easing monetary policy advances the transitory increase in the real output and upsurge the inflation rate, resulting a permanent increase in nominal interest rates. In a nutshell, the channels of monetary transmission are not mutually exclusive, in response to changes in monetary policy incorporates the combined effects of all the channels. Thus, the potential endogenous response of policy to economic conditions is serious impediment to identify the contribution of the individual channels through which the overall impact of monetary policy are transmitted to the economic system and to the inflation rate.

1.1 Malaysia Conditions
Banking crisis took place due to the economic recession in 1985. This recession was caused by weak demand and shortcomings in regulatory and accounting framework as well as inadequate supervision for Deposit Taking Cooperative (DTC) sector. Controls on
interest rates were re-imposed to curtail the problem of recession. In January 1994, capital control was adopted to curb short-term capital inflows, in which residents were prohibited to sell the Malaysian securities to nonresidents³. Meanwhile, following the 1997 Asian financial crisis, the value of ringgit fell. The ringgit was pegged against the USD at US$1.00=RM3.80 on September 2, 1998 to safeguard currency fluctuations and to prevent speculative attacks and cross hedge activity across both foreign exchange and the equity markets. On 21 July 2005, the Malaysian government announced the scrapping of the seven-year ringgit’s peg to the dollar and the ringgit is now operating in a managed float with reference to a basket of currencies with the major trading partners. On the other hand, selective capital controls were instituted on September 1, 1998. The imposition of selective exchange controls disabled the offshore transactions of ringgit and regained monetary independence. However, the adoption of a capital control regime by the Malaysian government was heavily debated for the past years⁴.

The present study aims to estimate the elasticities of the deterministic variables with respect to the real GDP. This paper looks at Autoregressive Distributed Lag (ARDL) bounds testing approach (Pesaran et al., 2001) to determine if cointegration between real GDP, long-term and short-term interest rate, exchange rate, share price, and claims on private sector persist over the quarterly period 1980:1-2004:4. Different transmission mechanisms of monetary policy as well as the time lag involves for each component impacts on output are crucial to be examined.

The remaining of the paper is organized as follows. Section 2 outlines the empirical model framework, variables selection and data sources for the investigation. Section 3 analyzes the empirical results. Section 4 depicts concluding remarks.

2. Methodology: Model specification and data

Abide by the demand pressures approach, following Stevens (1998), the weights of MCI is estimated using two major variables that affect AD function vis-à-vis the single equation based MCI, as shown in equation (1).

\[ y = ar + be + \text{other variables} \] (1)
where $y$ is the natural logarithm of the real GDP [calculated by the ratio of nominal GDP on percent of CPI (2000=100)]$^5$, $r$ is the short-term real interest rate$^6$ [following Batini and Turnbull (2002) and OECD (1996), the ex-ante short-term real interest rate is measured by the difference between the 3-month Treasury bill rate and actual inflation rate], and $e$ is the natural logarithm of real exchange rate [as units of Ringgit Malaysia per unit of US dollar]. The parameter $a$ and $b$ are the coefficients terms on interest and exchange rate in the demand equation. The higher the weight for one variable implies greater role-plays of this variable for adjusting than the other variables when shocks happen.

The details of “other variables” are as follows: 1) Government bonds yield as proxy for long-term interest rate (bond) [5-Year coupon rate on Federal Government Securities minus CPI]; 2) the Kuala Lumpur Composite Index (KLCI) – the real share price ($Sp$) to account for asset prices channel [KLCI deflated by the CPI]; 3) real claims on private sectors ($Cops$) is used as a proxy for credit channel [Cops/CPI%]. All variables with the exception of the interest rate are expressed in logarithms (Guender, 2001; Burger and Knedlik, 2003) and all series are expressed in real terms. $Sp$ is collected from datastream, bond yield is gathered from SEACEN Financial Statistics, while other data are obtained from the International Financial Statistics, which is published by the International Monetary Fund (IMF).

The bounds testing proposed by Pesaran et al. (2001) is applied to examine cointegration relation between output and its determinant variables. Bounds test approach is based on the estimation of an unrestricted error-correction model (UECM). ARDL estimation strategy is valid asymptotic inference that use the ordinary least square (OLS) estimates, provided the values of the maximum lag lengths are appropriately chosen to mitigate any residual serial correlation and the problems of endogenous regressors, irrespective of whether the variables are I(0) or I(1) (Hsiao, 1997; Pesaran et al., 2001). Bounds test involves two asymptotic critical value bounds. If the test statistics exceeds their respective upper critical values, there is evidence of a long-run relationship, and reject the null hypothesis of no cointegration, regardless of the order of integration of the variables. If it falls below the bound, the null hypothesis of no cointegration cannot be rejected, and if it lies between the bounds, inference is inconclusive.

A general function of the real GDP in the equation (1), can be written as: $LRGDP_t = f(r_t, e_t, Rbond_t, Cops_t, Sp_t)$, where $r$ for short-term and $Rbond$ for long-term interest rate channel, $e$ for exchange rate channel, $Sp$ as a proxy for asset price channel, and $Cops$ for credit channel. Meanwhile, the deterministic regressor consist of TIME trend to account for changes in innovations, an intercept, and four dummy$^7$ variables to address 1) Asian financial crisis; 2) pegging exchange rate regime, 1998Q3-2004Q4; 3) capital control on


$^6$ Fisher effect indicates the one-for-one adjustment of the nominal interest rate to the inflation rate. Therefore, the real interest rate is the nominal interest rate adjusted for the effects of inflation.

$^7$ The asymptotic theory developed for bound test procedure is not affected by the inclusion of such “one-off” dummy variables (Pesaran et al., 2001, p.307).
1994Q1; and 4) recession effect, 1985-1987. Hence, the log-linear model is specified in equation (3).

\[
LrGDP_t = \beta_0 + \beta_1 r_t + \beta_2 Le_t + \beta_3 Rbond_t + \beta_4 LrCops_t + \beta_5 LrSp_t + \beta_6 TIME + \beta_7 DUMcrisis + \beta_8 DUMpeg + \beta_9 DUM941 + \beta_{10} DUM8587 + \epsilon_t
\]  

(3)

Positive values are expected for \( \beta_4 \) and \( \beta_5 \), while \( \beta_1, \beta_3 \) and \( \beta_6 \) should be negative. To carry out the bounds test, the equation (3) is converted into UECM form as represented by equation (4) to test for cointegration. Variables up to lag four are included in the estimation of ARDL equation, considering the common practice of using quarterly data for the optimal order of the lags in the ARDL model (Pesaran and Pesaran, 1997:304).

\[
\Delta LrGDP_t = \beta_0 + \beta_1 LrGDP_{t-1} + \beta_2 r_{t-1} + \beta_3 e_{t-1} + \beta_4 Rbond_{t-1} + \beta_5 LrCops_{t-1} + \beta_6 LrSp_{t-1} + \beta_7 TIME + \beta_8 DUMcrisis + \beta_9 DUMpeg + \beta_{10} DUM941 + \beta_{11} DUM8587 + \sum_{i=1}^{p} \beta_{12i} \Delta LrGDP_{t-i} + \sum_{i=0}^{p} \beta_{13i} \Delta r_{t-i} + \sum_{i=0}^{p} \beta_{14i} \Delta Le_{t-i} + \sum_{i=0}^{p} \beta_{15i} \Delta Rbond_{t-i} + \sum_{i=0}^{p} \beta_{16i} \Delta LrCops_{t-i} + \sum_{i=0}^{p} \beta_{17i} \Delta LrSp_{t-i} + \epsilon_t
\]  

(4)

where \( \beta_0 \) is an intercept term, \( \Delta \) is difference operator, \( \epsilon_t \) is the random error terms, and \( p \) is the lag length. The long-run relationship between the real GDP and its determinants is tested by imposing restriction on the jointly significant estimated parameters for cointegrating test (Pesaran et al., 2001), with \( H_0: \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = \beta_6 = 0 \) (no cointegration), and \( H_1: \) at least one \( \beta_i \neq 0, i = 1,2, \ldots, 6. \) (cointegration).

The long-run elasticity can be derived from UECM that is the estimated coefficient of the one-lagged explanatory variables (multiplied with a negative sign) divided by the estimated coefficient of the one lagged dependent variable (Bardsen, 1989). Meanwhile, the estimated coefficient of the first differenced variable in UECM is the short-run elasticity.

Initially, all the explanatory variables are used to account for different transmission mechanism channels; however results show insignificant for \( LrCops, r, \) and \( RBond^8 \). Exclusion test also demonstrates that these variables are statistically insignificant at 5 percent level and should be excluded from the model. Therefore, UECM model is reexamined using general-to-specific modeling strategy to obtain a parsimonious model by dropping the most insignificant first differenced variables sequentially, i.e., \( LrCops \). The process is continued and the final UECM model for Malaysia is parsimonious as:

\[
\Delta LrGDP = \beta_0 + \beta_1 LrGDP_{t-1} + \beta_2 \epsilon_{t-1} + \sum_{i=1}^{q} \beta_{4i} \Delta LrGDP_{t-i} + \sum_{i=0}^{q} \beta_{5i} \Delta r_{t-i} + \sum_{i=0}^{q} \beta_{6i} \Delta LrSp_{t-i} + \epsilon_t
\]  

(5)

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8 Results are available upon request.
3. Empirical results

Augmented Dickey Fuller (ADF), Kwiatkowski, Phillips, Schmidt, and Shin (hereafter KPSS) and Ng and Perron (2001) unit root tests have been employed to examine stationarity of the series. The results of the unit root test are reported in Table 1. Results reveal that all series are I(1) variables.

Table 1: Results of Augmented Dickey Fuller (ADF), Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) and Ng-Perron unit root test

<table>
<thead>
<tr>
<th>Variable</th>
<th>ADF Level</th>
<th>ADF First difference</th>
<th>KPSS Level</th>
<th>KPSS First difference</th>
<th>Ng-Perron Level</th>
<th>Ng-Perron First difference</th>
</tr>
</thead>
<tbody>
<tr>
<td>LrGDP</td>
<td>-2.833(1)</td>
<td>-9.072(0)***</td>
<td>0.238(3)***</td>
<td>0.122(2)</td>
<td>-9.095(1)</td>
<td>-51.431(2)***</td>
</tr>
<tr>
<td>r</td>
<td>-2.277(1)</td>
<td>-14.781(0)***</td>
<td>0.198(8)**</td>
<td>0.151(2)</td>
<td>-6.846(1)</td>
<td>-11.622(1)***</td>
</tr>
<tr>
<td>RBond</td>
<td>-2.603(1)</td>
<td>-17.363(0)***</td>
<td>0.162(8)**</td>
<td>0.177(8)</td>
<td>-7.361(1)</td>
<td>-9.632(1)**</td>
</tr>
<tr>
<td>Le</td>
<td>-2.287(0)</td>
<td>-9.029(0)***</td>
<td>0.160(3)**</td>
<td>0.049(0)</td>
<td>-9.674(1)</td>
<td>-42.334(0)***</td>
</tr>
<tr>
<td>LrSp</td>
<td>-2.705(0)</td>
<td>-10.932(0)***</td>
<td>0.147(8)**</td>
<td>0.046(2)</td>
<td>-12.447(0)</td>
<td>-22.534(2)**</td>
</tr>
<tr>
<td>LrCops</td>
<td>-1.916(1)</td>
<td>-5.791(0)***</td>
<td>0.386(1)***</td>
<td>0.327(1)</td>
<td>-1.558(0)</td>
<td>-17.136(0)***</td>
</tr>
</tbody>
</table>

Critical values:

- 1%: -4.055
- 5%: -3.457
- 1% level of significance.

Notes: LrGDP is the real GDP, r is real 3-month Treasury bill, RBond is the government bond yields, e is the exchange rate, LrSp is the share prices index, and LrCops is the claims on private sectors. All series are in natural logarithm except for r and RBond, and all series are in real term. The number in parentheses represents the optimal lagged length. Asterisks (**, *** ) denote the rejection of the null hypothesis of a unit root at 5%, and 1% level of significance respectively. The source of critical values is based on MacKinnon (1991) for Augmented Dickey Fuller, Kwiatkowski-Philips-Schmidt-Shin (1992, Table 1) for KPSS test and Ng-Perron (2001).

Table 2: Real GDP functions – UECM

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>t-statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>3.8858</td>
<td>7.6326***</td>
</tr>
<tr>
<td>DLrGDPt-1</td>
<td>-0.11156</td>
<td>-1.6853*</td>
</tr>
<tr>
<td>DLrGDPt-4</td>
<td>0.44335</td>
<td>6.7027***</td>
</tr>
<tr>
<td>DLc-1</td>
<td>0.13737</td>
<td>1.9945**</td>
</tr>
<tr>
<td>DLrSp-2</td>
<td>0.02845</td>
<td>1.7655*</td>
</tr>
<tr>
<td>DLrSp-3</td>
<td>0.03341</td>
<td>2.2569**</td>
</tr>
<tr>
<td>LrGDPt-1</td>
<td>-0.41116</td>
<td>-7.5502***</td>
</tr>
<tr>
<td>Le-1</td>
<td>-0.16230</td>
<td>-3.8291***</td>
</tr>
<tr>
<td>LrSp-1</td>
<td>0.055886</td>
<td>3.3805***</td>
</tr>
<tr>
<td>TIME</td>
<td>0.007006</td>
<td>7.5452***</td>
</tr>
<tr>
<td>DUMcrisis: 1 for 1997Q3-1999Q2: 0 for others</td>
<td>0.038992</td>
<td>3.6435***</td>
</tr>
<tr>
<td>DUMpeg: 1 for 1998Q3-2004Q4: 0 for others</td>
<td>0.028944</td>
<td>2.2469**</td>
</tr>
<tr>
<td>DUM941: 1 for 1994Q1: 0 for others</td>
<td>-0.083885</td>
<td>-3.4374***</td>
</tr>
<tr>
<td>DUM8587</td>
<td>-0.030812</td>
<td>-3.1690***</td>
</tr>
</tbody>
</table>

Notes: Sample (adjusted): 95 observations, 1981Q2 to 2004Q4. R-squared: 0.68917; adjusted R bar-squared: 0.63928; Standard Error of Regression: 0.022220; residual sum of squares: 0.039992; F-statistic (P-value): 13.8145 (0.000); Durbin-Watson statistic: 1.7962; Breusch-Godfrey lagrange multiplier LM[1]: 1.0725 (0.300); and autoregressive conditional heteroskedasticity ARCH test [1]: 0.17961 (0.672); Ramsey RESET functional test [1]: 2.3967 (0.122), Jarque Bera normality test= 0.68379 (0.710) . Figure in brackets
The estimated UECM model in equation (5) is adequate when the model satisfies the assumption of the classical linear regression model. Results in Table 2 show parsimonious estimated UECM passes a battery of diagnostic tests. Diagnostic tests confirm the model has the desired properties of OLS such as uncorrelated Breusch-Godfrey serial correlation LM test, homoscedasticity of residuals (ARCH test), correct Ramsey RESET specification test, and Jarque-Bera normality test at 5% level.

The results of bounds test for cointegration analysis are reported in Table 3. Since the computed F-statistic, $F(\text{LrGDP} | e, \text{LrSp})=8.2503$ exceeds the upper critical value $I(1)$ band of 7.52 at 1 percent level, the result of the inference shows the null hypothesis of no cointegration can be rejected, indicating that real GDP, exchange rate and share price are cointegrated.

To reaffirm the cointegration results, Banerjee et al. (1998) have developed the error-correction mechanism test based on the OLS coefficient of the lagged dependent variable in level in the UECM. From Table 2, the computed $t$-statistic for the coefficient on $\text{LrGDP}_{t-1}$ exceeds the critical value at the 1% significance level. The significant and negative magnitude of the ECT of 0.41 reflects the moderate speed of adjustments from short-run disequilibrium towards the long-run equilibrium state, which again confirms the evidence of long-run relationship between the examined variables. Obviously, both the ECM cointegration techniques and bounds test approach have provided empirical evidence of the presence of cointegration during the period under study. The finding of a cointegrating relation indicates that the real GDP equations in this study are specified correctly with their determinants.

The estimated long-and short-run elasticities of the real GDP function are presented in Table 4. Consistent with economic theory, the estimated long-run elasticity is found to be higher than in the short-run. Since there is cointegration, thus we emphasize on the long-run equation. In the long-run, all the series seem to correctly sign as expected. It is expected that the real GDP is negatively associated with the real exchange rate appreciation. The estimated long-run cointegrating equation is as follows:

$$\text{LrGDP} = 9.451 \cdot 0.394 e + 0.136 \text{LrSp} - 0.125 \text{TIME} + 0.095 \text{DUMcrisis}$$
$$+ 0.070 \text{DUMpeg} - 0.204 \text{DUM941} - 0.075 \text{DUM8587}$$  

### Table 3: Results of the bounds test for cointegration analysis

<table>
<thead>
<tr>
<th>Computed F-statistics (Wald test) :</th>
<th>8.2503***</th>
</tr>
</thead>
<tbody>
<tr>
<td>$H_0 : \beta_1 = \beta_2 = \beta_3 = 0$</td>
<td></td>
</tr>
<tr>
<td>Asymptotic critical values bounds</td>
<td></td>
</tr>
<tr>
<td>1% level</td>
<td></td>
</tr>
<tr>
<td>Lower bound, $I(0)$</td>
<td>6.34</td>
</tr>
<tr>
<td>Upper bound, $I(1)$</td>
<td>7.52</td>
</tr>
</tbody>
</table>

According to Perman (1991:20), if we fail to find a cointegrating vector for a given variables sets, a broader set of series is needed to see if they are cointegrated. Hence, cointegration analysis could serve as a misspecification test, or equivalently as a guide to variable selection.
Notes: The reported critical values are from Pesaran et al. (2001) Table CI(v) Case V: Unrestricted intercept and trend (page 301) for F statistic ($k=2$), $k$ is the number of regressors.

Table 4: Estimated short-run and long-run elasticity of real GDP based on equation (4) (Dependent variable is $LrGDP$)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Short-run</th>
<th>Long-run</th>
</tr>
</thead>
<tbody>
<tr>
<td>Le</td>
<td>0.13737**</td>
<td>-0.39474**</td>
</tr>
<tr>
<td>LrSp</td>
<td>0.06186**</td>
<td>0.135923**</td>
</tr>
<tr>
<td>C</td>
<td>3.8858***</td>
<td>9.450822***</td>
</tr>
<tr>
<td>TIME</td>
<td>0.007006***</td>
<td>-0.12536***</td>
</tr>
<tr>
<td>DUMcrisis</td>
<td>0.038992***</td>
<td>0.094834***</td>
</tr>
<tr>
<td>DUMpeg</td>
<td>0.028944**</td>
<td>0.070396**</td>
</tr>
<tr>
<td>DUM941</td>
<td>-0.083885***</td>
<td>-0.20402***</td>
</tr>
<tr>
<td>DUM8587</td>
<td>-0.030812***</td>
<td>-0.07494***</td>
</tr>
</tbody>
</table>

Note: Asterisks (** and ****) denotes statistically significant at 5% and 1% level respectively. The structural form long-run coefficient for $Le$ and $LrSp$ are derived as ($\beta_2/\beta_1$), and ($\beta_3/\beta_1$), respectively from the estimated equation UECM. Sample adjusted: 95 observations.

Results of the exclusion Wald test in Table 5 conclude that all the series are statistically significant at 1% level. Meanwhile, Wu-Hausman test is employed to test for the exogeneity of the regressors. The computation of the Wu-Hausman statistic can be carried out by running OLS regression of the auxiliary exogenous variables in the model, and the residuals are saved as $Re$ and $RLrSp$. The results in Table 6 reveal that the computed $F$-statistics of 0.3997 cannot reject the null hypothesis of exogeneity. Similarly, the $t$-ratio of the exchange rate and share price variables are 0.046 and 0.222 respectively. They are found to be insignificant at 5% level, suggesting that $Le$ and $LrSp$ are exogenous cannot be rejected.

Table 5: Results of Likelihood Ratio Exclusion Test

<table>
<thead>
<tr>
<th>Variables</th>
<th>LrGDP</th>
<th>Le</th>
<th>LrSp</th>
<th>TIME</th>
<th>DUMcrisis</th>
<th>DUMpeg</th>
<th>DUM941</th>
<th>DUM8587</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\chi^2$-statistics</td>
<td>.000***</td>
<td>.000***</td>
<td>.001***</td>
<td>.000***</td>
<td>.000***</td>
<td>.025**</td>
<td>.001***</td>
<td>.002***</td>
</tr>
</tbody>
</table>

Notes: Asterisk (****) denotes statistically significant at 1% level. Figure in parentheses are the p-value.

Table 6: Wu-Hausman statistic for testing the exogeneity of $Le$ and $LrSp$

<table>
<thead>
<tr>
<th>Regressors</th>
<th>Coefficient</th>
<th>t-ratio</th>
</tr>
</thead>
<tbody>
<tr>
<td>$RLe$</td>
<td>-0.0033102</td>
<td>-0.046485 (0.963)</td>
</tr>
<tr>
<td>$RLrSp$</td>
<td>-0.0038462</td>
<td>-0.22269 (0.824)</td>
</tr>
<tr>
<td>$F$-statistic (2, 81)</td>
<td>-</td>
<td>0.025332 (0.975)</td>
</tr>
</tbody>
</table>

Notes: $RLe$ and $RLrSp$ are the residuals from $Le$ and $LrSp$, regressions respectively. Figure in brackets () are p-values.

4. Concluding remarks
This study investigates the long-run relationship of components in the real GDP by considering key transmission mechanism channels using the bounds testing approach proposed by Pesaran et al. (2001) for the quarterly period 1980:1-2004:4. It is evident that the real GDP, real exchange rate, and real share price are cointegrated in the long-run. A possible light of policy implication can be drawn are: First, it is noteworthy that
both the short-run and long-run interest rates show insignificant variables in the model. It is not surprise as Ip (2001) claims that interest rate is not a good indicator of monetary conditions since changes can be caused by factors not related to changes in the saving market. For instance, a lower interest rate and exchange rate compared to the base year may not signal easing monetary conditions because interest rate may have been affected by changes in investment demand, and exchange rate by changes in the terms of trade. Second, asset prices channel shows significant transmission mechanism channel in influencing the conduct of monetary policy and ultimately affecting the real GDP. Observation of the share price would be a crucial source of information in setting monetary policy stance. Third, credit channel does not significantly influence the real GDP during the period study. Fourth, changes in real GDP or the rate of inflation depend on the source of disturbances.
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