AN EMPIRICAL MODELLING ON SAVINGS BEHAVIOUR IN MALAYSIA

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Abstract

The purpose of this study is to re-investigate the savings behaviour in Malaysia through the cointegration and causality analyses. This study covered the annual data from 1970 to 2004. The results of bounds test confirmed a long run equilibrium relationship between savings and its determinants, namely real income, dependency ratio and real interest rates. The study found that the major causes of savings in Malaysia are real income and dependency ratio. On the other hand, the effect of real interest rate on savings in Malaysia is less. Therefore, monetary policy is not an effective macroeconomic policy instrument to encourage savings in Malaysia although savings and interest rates showed bilateral causality. In addition, savings Granger caused economic growth in Malaysia and the study found that households in Malaysia are very protective and the savings behaviour is driven by the precautionary savings.

1. Introduction

Malaysia is one of the fastest growing economies in the Southeast Asia in conjunction with high growth rate and low inflation rate. The flourishing economic growth in Malaysia is always linked to the rapid expansion of intraregional trades and flows of foreign direct investment (FDI). However, high savings rate may be an alternative source for the rapid growth of the Malaysian economy. The savings rate in Malaysia was the second highest among the Association of Southeast Asian

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Nations (ASEAN) countries and the third highest among the Newly Industrialising Economies (NIEs). According to Solow's (1956) and Romer's (1986) growth theories, high savings rate will lead to higher economic growth through its impact on capital formation. However, Lin (1992) added that economic growth can be sustained only if the resources such as savings are mobilised efficiently and translated effectively into the productive activities. Therefore, savings is Granger causes economic growth only if the resources are optimally mobilised.

Figure 1
Domestic Savings Trend in Malaysia (1970 to 2004)
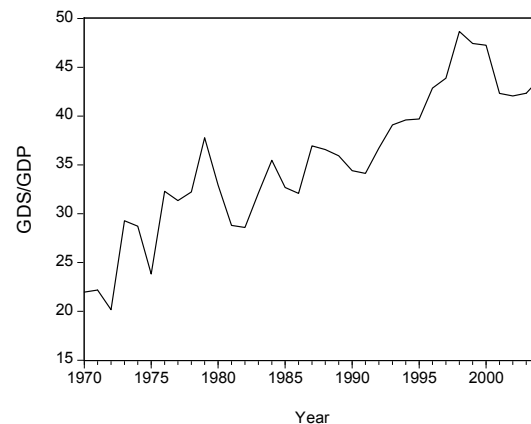
Historically, Malaysia had managed to sustain a high level of savings rate. Overall, the country saved an average of 35.3 per cent of GDP a year from the past three decades. The savings rate in Malaysia from 1970 to 2004 is depicted in Figure 1. From the visual inspection, the savings rate in Malaysia has been oscillating from 32.9 per cent in 1980 to 34.4 per cent in 1990. In 1998, the savings rate level peaked at 48.7 per cent but falls sharply to 42.1 per cent in 2001. This is from the impact of Asian financial crisis that hit Thailand in the middle of 1997. The crisis has spread to Malaysia in 1997 and triggered a massive outflow of foreign capital in the late 1998 due to serious depreciation of Ringgit Malaysia. Consequently, savings rates in Malaysia declined tremendously.

In the past literatures, study on savings behaviour in Malaysia is relatively few and their finding failed to reach unanimous result. Kuznets (1960) was the first study that empirically examines the relationship between savings rate and disposable income. The study found that savings rate and income is positively correlated. However, he concluded that high savings rate is less important for the low income

per capita countries. In view of the relation between savings and real interest rate, Fry (1980) estimated a savings function for seven Asian developing countries to examine the validity of the McKinnon and Shaw model on financial repression. The sample period covered the annual data from 1962 to 1972 (70 observations). The study found that the real interest rate exerts a positive effect on the savings rate in Asian countries. On the other hand, Giovannini (1983) revisited the Fry's equations using the same set of countries but over a different sample period. The study found that real interest rate elasticity on savings rate is significantly negative. This is contradicted with the finding of Fry (1980). The author argued that the interest rate elasticity of savings is significantly positive in Fry's (1980) study is mainly due to the omission of relevant variables such as tax and others potential variables. This magnitude is clearly demonstrated by the significant of the lagged dependent variable.

Life-Cycle Hypothesis (LCH) and Leff (1969) noted that dependency ratio is negatively related to savings rate. Gupta (1971) argued that dependency ratio does not appear to play any role in the low income per capita countries. However, the aggregate savings in these countries is determined by other variables such as income per capita and growth rate of income per capita. In contrast, Rossi (1989) found that dependency ratio has a significant negative effect on saving rates in developing countries. This implied that the decrease of labour participation rate will subsequently decrease the savings rate in developing countries. Therefore, demography is a vital factor to explain savings behaviour in developing countries. In the context of Malaysia, Hamid and Kanbur (1993) conducted a study to investigate the savings behaviour over the period of 1970 to 1990. The study found that gross national saving is determined by real disposable income, dependency ratio, current account (as a proxy for foreign savings) and income growth rate. However, the authors concluded that the real interest rate of the commercial banks is not an effective policy instrument to increase the savings rate in Malaysia. Therefore, the result did not support the financial repressionist theory which stressed the importance of the real interest rate in raising the savings rate. Furthermore, they found that dependency ratio is negatively related to gross national saving in Malaysia. This implied that the savings rate in Malaysia will decrease when the non-productive population size is increase. This result supports the LCH argument and Leff's (1969) study.

Recently, Agrawal (2001) has investigated the savings function for seven Asian countries by using Stock and Watson (1993) Dynamic Ordinary Least Squares (Dynamic OLS) procedure and vector error-correction modelling (VECM) approach. The study found that savings rate and its determinants for Malaysia, Korea, Taiwan and India are stationary at different orders. Thus, conventional cointegration test cannot be

adopted to examine the presence of long run equilibrium relationship. Consequently, the author assumed that savings rate and its determinants for these countries are not coalescing in the long run. However, the study employed the Granger causality test based on vector autoregressive (VAR) modelling approach to investigate the causal link between savings and economic growth. The study found that savings rate Granger cause economic growth in Malaysia, but there is no evidence of reverse causality. For Korea and Taiwan, the author found that savings rate and economic growth is neutral, but economic growth Granger cause savings rate in India. On the other hand, Baharumshah *et al.* (2003) conducted a study to investigate the savings behaviour in five Asian countries through the Johansen and Juselius (1990) multivariate cointegration test. The sample periods covered the annual data from 1970 to 1998. The study found that savings and its determinants, namely income, interest rate, dependency ratio and current account are cointegrated in all the selected Asian countries. With this finding, they proceed to determine the causality direction through the VECM approach. The study showed that the causal link between savings and economic growth is neutral in all the selected Asian countries, except Singapore. For Singapore, the authors showed that savings Granger cause economic growth, but not vice versa. More recently, Mohan (2006) investigated the relationship between income growth and savings for various economies with different class of income (e.g. low-income, lower-middle income, upper-middle income, and high-income). The author found that the order of integration between gross domestic product (GDP) and gross domestic saving (GDS) for Egypt, Malaysia and the United States are not uniform. Therefore, these three countries are excluded from the study because the Johansen-Juselius cointegration test cannot be used. Moreover, the use of Granger causality test to determine the causality direction for these countries would lead to misspecification problem. The finding of the study suggests that savings and income growth are cointegrated for most of the economies. Furthermore, the author found that the causality evidence for the low-income economies is mixture. However, the causality direction for the lower-middle income and the high-income economies are running from income growth to savings. For the upper-middle income economies, the causality direction for savings and income growth is bilateral.

From the past literature, there are some empirical issues on savings behaviour that we have to consider in this study. The first issue is related to the LCH, which suggests that savings depend on population age alone. According to LCH, only the young and the old or retired are dissave because they are the non-productive population. So that the higher the dependency ratio the lower the savings rate. The efforts of the earlier empirical works have focused on the effect of dependency ratio on domestic savings due to its ambiguity effect. In this study, we argue

that the computational method of dependency ratio is not correct because it does not take into consideration the unemployed labour force which is also a dis-savings population in an economy. This may one of the valuable exploration and also contribution to the existing literature on savings. The second issue is whether interest rates play role on the savings rate. The literature indicates the interest rate is positively related to savings if the positive substitution effect dominates the negative income effect on savings. Otherwise, the interest rate is negatively related to savings.

The final issue is on the econometric techniques, the existing studies used the conventional cointegration (Engle and Granger, 1987; Johansen and Juselius, 1990) and Granger causality tests. Thus, pre-testing of unit root of the interested variables is required to ensure that the order of integration is uniform. Mah (2000) noted that the conventional cointegration tests are not appropriate in finite sample study. In addition to that, Gonzalo and Lee's (1998) Monte Carlo evidence shows that Johansen and Juselius (1990) cointegration test tend to find spurious cointegration with probability approaching to one when the order of integration is not purely $I(1)$ process. Besides that, the Granger causality test based on VECM approach as in Baharumshah *et al.* (2003) is considered low power because it sensitive to the value of the nuisance parameter (Zapata and Rambaldi, 1997). If the first differenced VAR is used such as Agrawal's (2001) study the Granger causality test may be bias due to loss of long run causal information (see Granger, 1988). Thus, the results produced by the earlier studies should be accepted with caution.

In view of these, it will be useful for this study to re-investigate the savings behaviour in Malaysia through the relatively new bounds testing procedure for cointegration (Pesaran *et al.,* 2001) and modified Wald test (Toda and Yamamoto, 1995). The advantage of using these tests is that it is applicable irrespective of whether the underlying regressors are purely $I(0)$ or purely $I(1)$ process. In addition, Mah (2000) documented that the bounds test approach has superior statistical properties in small sample.[1] Thus, finding with these approaches is relatively more reliable than conventional tests.

The remainder of this study is organised as follows. Section 2 will briefly discuss the model specification, econometric techniques and data used

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[1] The author employed bounds test with 18 annual observations to investigate the presence of long run equilibrium relationship between import demand and its determinants in Korea. However, none of the Monte Carlo study is conducted to verify the small sample properties of bounds test approach. To the best of our knowledge, there are some published studies supporting the use of bounds test in small sample and among them are Pattichis (1999), Mah (2000) and Tang (2001).

in this study. The empirical results will be reported in Section 3. Finally, the conclusion and some policy implication are presented in Section 4.

2. Model Specification, Econometric Techniques and Data

2.1 *Model Specification*

In the past years a large amount of empirical work has appeared to examine the savings behaviour in developed and developing countries. The most common explanatory variables are income, interest rates and dependency ratio. The economic theory suggested that income is positively related to savings. Thus, the higher the income growth rate the higher is the savings rate. However, the relationship between interest rates and savings is mixture due to the well known positive substitution effect and negative income effect on savings.[2] In line with this, Olson and Bailey (1981) found that the positive substitution effect on savings dominates the negative income effect on savings. Furthermore, Baharumshah *et al.* (2003) found that interest rate is positively related to savings in all the selected Asian countries, except Thailand.

A numbers of studies have been performed to examine the effect of dependency ratio on savings. Among them are Rossi (1989), Edwards (1996), Faruqee and Husian (1998), Agrawal (2001) and Baharumshah *et al.* (2003). However, the impact of dependency ratio on savings is mixture. According to LCH, savings is negative when individuals are young (below 15 years old) and old or retired (after 65 years old). While savings is positive when individual is productive (between 16 to 64 years old). Thus, the LCH suggests that age distribution of the population will have an inverse effect on aggregate savings. On the other hand, Faruqee and Husian (1998) found that the dependency ratio is positively related to savings in developing economies.

This study argues that the computational method of dependency ratio may be incorrect because it does not take into consideration the unemployed labour force which is also a dis-savings population. Broadly speaking, there are three generations involved in an economy, namely the young, working and old or retired generations. However, there is a group of people from the working generation who are currently unemployed and reliant on savings especially during economic slowdowns or recessions. The relevant of unemployment to savings can be derived from the simple Keynesians' consumption functions as below:

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[2] Government intervention may also cause the mixture effect of interest rate on savings.

$$C_i = \alpha + \beta Y_i, \tag{1}$$

where *C* is the consumption expenditure and *Y* is the income. α represents the autonomous expenditure and β is the marginal propensity to consume (MPC). Equation (1) indicates that consumption expenditure is dependent on income and the effect of income on consumption expenditure is measured by the parameter β or $MPC = \Delta Y / \Delta C$. According to Keynesians' theory, the autonomous expenditure (α) is always positively related to consumption. Assuming that an individual is unemployed, therefore their income will be zero and consumption function may be expressed as in the Equation (2):

$$C_i = \alpha + \beta Y_i$$
$$C_i = \alpha + \beta(0)$$
$$C_i = \alpha . \tag{2}$$

By solving Equation (2) the consumption function above clearly shows that when income is zero (unemployed), individual will still consume by using their savings. That's why α is always positively related to consumption expenditure. The above explanation clearly shows that unemployment is highly correlated with savings and hence the dependency ratio used by the earlier studies may not have captured the actual scenario. This may be the plausible reason of non-consistency results produced by the earlier studies. In this circumstance, we propose to use modified version of dependency ratio (MDR). The MDR is a re-computes series by Equation (3):

$$MDR = \frac{N - L + U}{N}, \tag{3}$$

where N is the total population, L is the employed labour force[3] and U is the unemployed labour force. As a result, we specify the savings function for this study as follows:

$$\ln S_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln MDR_t + \beta_3 R_t + \varepsilon_t \tag{4}$$

where $\ln$ denotes the natural logarithm. S_t is the real gross domestic savings, Y_t is the real gross domestic products, MDR_t is the modified version of dependency ratio and R_t is the real interest rate. The

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[3] Employed labour force includes employee, employer, self-employed and unpaid family workers.

disturbance term ε_t is used to capture the unobservable effect and is assumed to be white noise.

2.2   *Cointegration Analysis*

In this study, we employ the bounds testing for cointegration procedure to examine the potential long run equilibrium relationship within the autoregressive distributed lag (ADRL) framework. The bounds test procedure was developed by Pesaran *et al.* (2001) and this is a multivariate non-residuals based cointegration test. Pattichis (1999) stated that the bounds test is likely to have better statistical properties because it does not push the short run dynamics into the residual terms as in the case of the Engle-Granger two-step approach. To implement bounds test, we estimate the following ARDL model.

$$\Delta \ln S_t = a_1 + a_2 \ln S_{t-1} + a_3 \ln Y_{t-1} + a_4 \ln MDR_{t-1} + a_5 R_{t-1} + \sum_{i=1}^{p} b_{1i} \Delta \ln S_{t-i}$$

$$+ \sum_{i=0}^{p} b_{2i} \Delta \ln Y_{t-i} + \sum_{i=0}^{p} b_{3i} \Delta \ln MDR_{t-i} + \sum_{i=0}^{p} b_{4i} \Delta R_{t-i} + \mu_t , \qquad (5)$$

where Δ is the first difference operator and the residuals μ_t are assumed to be white noise. The bounds test for cointegration is based on the standard Wald or *F*-statistics and the presence of long run equilibrium relationship is tested by restricting the lagged levels variables, $\ln S_{t-1}$, $\ln Y_{t-1}$, $\ln MDR_{t-1}$ and R_{t-1} in the Equation (5). Therefore, it is a joint significance *F*-test for the null hypothesis of no cointegrating relation $(H_0 : a_2 = a_3 = a_4 = a_5 = 0)$ against the alternative hypothesis of a cointegrating relation $(H_1 : a_2 \neq a_3 \neq a_4 \neq a_5 \neq 0)$.

We are aware of the fact that the critical values provided by Pesaran *et al.* (2001) is not suitable for our small sample size and hence we derive the appropriate critical values from the response surface procedure developed by Turner (2006). If the computed *F*-statistic exceeds the upper bounds critical values then the variables are cointegrated. Otherwise the variables are not cointegrated.

2.3   *Causality Analysis*

Toda and Yamamoto (1995) have proposed a relatively new procedure to determine the causality direction between a set of variables through the augmented $VAR(k + d_{max})$ model, where k is the optimal lag length in the VAR system and d_{max} refers to maximal order of integration suspected to

occur in the system.[4] This procedure is known as modified Wald (MWALD) test. The advantage of using this procedure is that it does not depend on whether the underlying variables are stationary at $I(1)$, or $I(2)$, or on whether the variables are cointegrated or not. Another advantage of using MWALD test is because of its simplicity. He and Maekawa's (2001) Monte Carlo experiment evidence shows that the F-statistic for Granger causality test often lead to spurious causality result when one or both of the variables are non-stationary. Furthermore, Granger's (1988) study indicates that causality test with first different VAR may be bias due to loss of long run causal information. Therefore, the MWALD test is preferred in this study. The study will estimate the augmented VAR system with the Seemingly Unrelated Regression (SUR) method because the former is more efficient than the standard VAR procedure (see Rambaldi and Doran, 1996).   To ascertain the causality direction among savings and its determinants, we estimate the following augmented VAR system in the SUR form.

$$
\begin{bmatrix} \ln S_t \\ \ln Y_t \\ \ln MDR_t \\ R_t \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \\ \alpha_4 \end{bmatrix} + \begin{bmatrix} A_{11,k} & A_{12,k} & A_{13,k} & A_{14,k} \\ A_{21,k} & A_{22,k} & A_{23,k} & A_{24,k} \\ A_{31,k} & A_{32,k} & A_{33,k} & A_{34,k} \\ A_{41,k} & A_{42,k} & A_{43,k} & A_{44,k} \end{bmatrix} \times \begin{bmatrix} \ln S_{t-k} \\ \ln Y_{t-k} \\ \ln MDR_{t-k} \\ R_{t-k} \end{bmatrix}
$$
$$
+ \begin{bmatrix} A_{11,p} & A_{12,p} & A_{13,p} & A_{14,p} \\ A_{21,p} & A_{22,p} & A_{23,p} & A_{24,p} \\ A_{31,p} & A_{32,p} & A_{33,p} & A_{34,p} \\ A_{41,p} & A_{42,p} & A_{43,p} & A_{44,p} \end{bmatrix} \times \begin{bmatrix} \ln S_{t-p} \\ \ln Y_{t-p} \\ \ln MDR_{t-p} \\ R_{t-p} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \\ \varepsilon_{4t} \end{bmatrix}, \qquad (6)
$$

where α_i is the constant terms. The residuals ε_{it} are assumed to be spherical distribution and white noise. The lag length of p represents $(k + d_{max})$. The optimal lag length k is selected based on Akaike's Information Criterion (AIC).[5]

Finally, the MWALD test is carried out by computing the χ^2- statistics on the lagged k explanatory variables. However, it should be pointed out here that the extra lagged explanatory variables, d_{max} is unrestricted because the inclusion of extra lag is to ensure that the asymptotic χ^2 distribution critical values can be applied when the test for causality between the integrated variables are conducted.

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[4] Dolado and Lütkepohl's (1996) simulation study shows that $d_{max} = 1$ is an ideal maximal order of integration for MWALD test.
[5] In finite sample study ($T < 60$), AIC and Final Prediction Error (FPE) are superior to other information criterion such as SBC (see Lütkepohl, 1991; Liew, 2004).

2.4 *Data*

The data used in this study are annual data and covered the sample period from 1970 to 2004. The gross domestic saving, gross domestic products and interest rate (fixed deposits rate) for Malaysia are extracted from World Bank, *World Development Indicators* (WDI), International Monetary Funds (IMF), *International Financial Statistics* and Bank Negara, *Monthly Statistical Bulletin.* The Consumer Price Index (CPI=2000) is used to derive the real term. The modified version of dependency ratio is derived from Equation (3). The annual data is used in this study is to avoid the seasonal bias problem. Furthermore, Hakkio and Rush (1991) noted that cointegration is a long run phenomenon, thus using long span of data is rather than merely high data frequency and short span of data. All variables are in natural logarithm (except for real interest rates).

As usual we start with examine the time series properties of each series. Although the Pesaran's *et al.* (2001) bounds testing procedure and the Toda and Yamamoto's (1995) MWALD test do not require the pre-testing unit root, but it is necessary to test the order of integration of the examined series to avoid the inclusion of *I*(2) explanatory variables into the bounds testing Equation.[6] Another advantage of using unit root test is to ensure that the order of integration of the examined series in the augmented VAR testing Equation is uniform. This study employed the augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwaitkowski *et al.* (1992) – KPSS stationarity tests.

According to the results presented in Table 1, all the variables are integrated of order *I*(1) process. However, it is interesting to report here that both ADF and PP stationarity tests for real interest rates (R_t) rejected the null hypothesis of unit root at level. This implies that real interest rate is stationary at level, *I*(0) process. However, it is well documented that ADF and PP tests failed to distinguish between a unit root and a near unit root stationary process, thus these tests is considered low power (Campbell and Perron, 1991; DeJong *et al.,* 1992). Therefore, the KPSS test is employed to affirm the order of integration of each series. The KPSS results indicate that all the variables included real interest rates rejected the null hypothesis of stationary at level, but cannot reject the hypothesis of stationarity when the variables are after first differencing. This finding is consistent with the notion that most of the macroeconomic series are non-stationary at level, but become stationary after first differencing (Nelson and Plosser, 1982). With this finding, we can proceed with the bounds testing procedure for cointegration and also the MWALD test.

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[6] Pesaran *et al.* (2001) noted that bounds testing procedure is applicable merely for either *I*(0) or *I*(1) process independent variables.

Table 1
The Results of Unit Root and Stationarity Tests

Variables	ADF	PP	KPSS
$\ln S_t$	−2.210 (3)	−3.176 (2)	0.691 (5)**
$\Delta \ln S_t$	−7.125 (1)*	−7.897 (4)*	0.074 (1)
$\ln Y_t$	−2.580 (3)	−2.553 (1)	0.694 (3)**
$\Delta \ln Y_t$	−5.395 (0)*	−5.390 (3)*	0.078 (3)
$\ln MDR_t$	−2.819 (1)	−2.269 (2)	0.169 (4)**
$\Delta \ln MDR_t$	−5.638 (4)*	−2.471 (1)	0.067 (3)
R_t	−4.474 (0)*	−4.422 (2)*	0.122 (0)***
ΔR_t	−7.220 (1)*	−15.392 (20)*	0.169 (12)

Note: The asterisks *, ** and *** denotes the significance level at 1, 5 and 10 per cent. ADF, PP and KPSS refer to Augmented Dickey-Fuller, Phillips-Perron and Kwiatkowski *et al.* (1992) unit root tests. The optimal lag length for ADF test is selected using the AIC while the bandwidth for PP and KPSS tests are selected using the Newey-West Bartlett kernel. Figure in parentheses denotes the optimal lag length and bandwidth. The critical values for ADF and PP tests are obtained from MacKinnon (1996) while the asymptotic critical values for KPSS test are obtained from Kwiatkowski *et al.* (1992). The null hypothesis for ADF and PP tests is H_0: a unit root, while the null hypothesis for KPSS test is H_0: trend stationary.

3. Empirical Results

3.1 *Bounds Testing for Cointegration Results*

To implement bounds testing procedure we start with determine the optimal lag order for the ARDL model. In annual data analysis a maximum lag orders of three to four years are sufficiently long to capture the system's dynamics'. The Schwarz Bayesian Criterion (SBC) indicates that ARDL(0, 0, 0, 1) is the appropriate lag structure. The results of bounds test for cointegration, together with the response surface critical values for T = 30, are presented in Table 2. A numbers of diagnostic tests were performed on the final ARDL model. The Ramsey RESET test indicates that the final ARDL model is free from the specification error problem. Both the Breusch-Godfrey LM test and Ljung Q-statistics for serial correlation failed to reject the null hypothesis of no serial correlation at 10 per cent significance level. This implied the estimated residuals are not serially correlated. The plots of CUSUM and CUSUM of Squares tests in Figure 2 show the estimated coefficients are stable over the sample period.

Table 2: The Estimated ARDL Equation

Dependent Variable: $\Delta \ln S_t$

Method: Ordinary Least Squares (OLS)

Independent Variable	Coefficient	*t*-Statistic
Constant	−4.646	−4.092*
$\ln S_{t-1}$	−0.520	−3.201*
$\ln Y_{t-1}$	0.904	3.985*
$\ln MDR_{t-1}$	1.491	3.100*
R_{t-1}	0.005	1.579
$\Delta \ln Y_t$	1.469	6.893*
$\Delta \ln MDR_t$	−2.145	−1.272
ΔR_t	−0.006	−2.134**
ΔR_{t-1}	−0.003	−2.097**

Bounds Test:

F-Statistics	5.832**	
# Adjusted Critical Values Bounds (F-test):	**Lower *I*(0)**	**Upper *I*(1)**
1%	6.328	7.993
5%	4.260	5.538
10%	3.416	4.532

Note: *, **, *** denote significance at 1, 5 and 10 per cent level, respectively. # Unrestricted intercept and no trend (k = 3 and T = 30) critical values are derived from response surface procedure proposed by Turner (2006). *R*-squared: 0.864; Adjusted *R*-squared: 0.918; *F*-Statistic: 42.733 (0.000); Jarque-Bera: 1.585 (0.453); Ramsey RESET [1]: 0.269 (0.604), [2]: 2.161 (0.339); Breusch-Godfrey LM test [1]: 0.072 (0.788), [2]: 1.151 (0.563); Ljung-Box Q-statistic test [1]: 0.046 (0.831), [2]: 0.816 (0.665); ARCH test [1]: 0.467 (0.494). [ ] refer to the order of diagnostic tests. ( ) refer to *p*-value.

In order to ascertain the presence of long run equilibrium relationship between savings and its determinants, a joint significance *F*-test for $(H_0 : a_2 = a_3 = a_4 = a_5 = 0)$ was conducted. The computed *F*-statistic for bounds test is 5.832 which is greater than the 5 per cent upper bounds critical values derived from Turner's (2006) response surface procedure. This implied that savings and its determinants, real income, real interest rate, dependency ratio are cointegrated.

Figure 2
The Plots of CUSUM and CUSUM of Squares Statistics
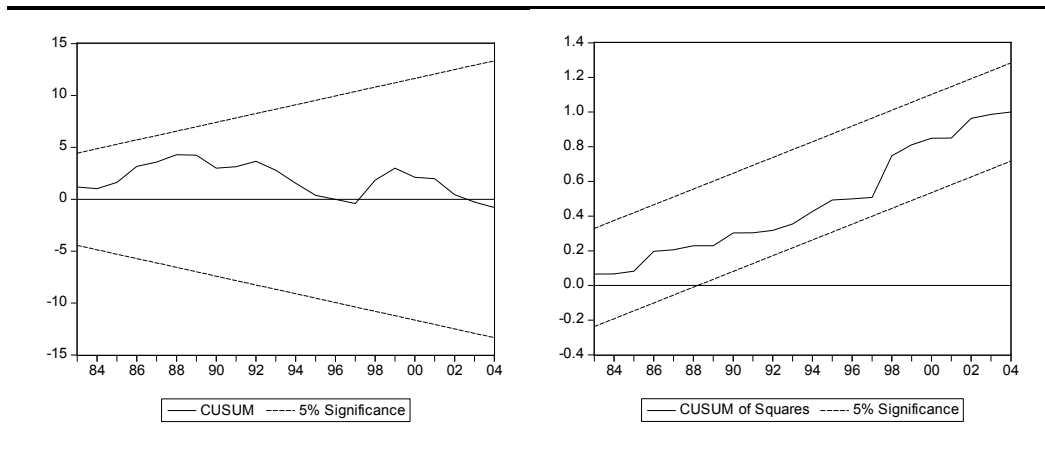
As the variables are cointegrated, the short and long run coefficients including the asymptotic significance level are derived from the ARDL model (Pesaran and Pesaran, 2002). Table 3 reported the short and long run coefficients derived from the ARDL model. In overall, the sign of the short and long run coefficients are inconsistent, except real income is positively related to savings. The positive sign of real interest rates coefficient is not significant in the long run, but the real interest rate carried a negative sign and is statistically significant in the short run. This indicates that the negative income effect on savings dominates the positive substitution effect on savings in Malaysia. This is consistent with the study of Giovannini (1983) but contrary with Fry (1978), Hamid and Kanbur (1993), Agrawal (2001) and Baharumshah *et al.* (2003). Therefore, tightening monetary policy such as increase of real interest rates may have an inverse effect on savings in Malaysia.

Table 3 shows that the dependency ratio is positively related to savings in the long run and the variable is statistically significance at 5 per cent level. However, this variable is not significant in the short run. As a result, we suggest that demographic factors are likely to explain the long run savings behaviour in Malaysia and the savings rate in Malaysia will increases when the non-productive population size is larger. The significant positive effect of dependency ratio has shed some light on the existence of precautionary savings behaviour in Malaysia. This is not consistent with the LCH argument and also Hamid and Kanbur's (1993) and Agrawal's (2001) studies for the case of Malaysia. One of the plausible explanations for this result is that the earlier study does not take into account the effect of unemployed labour force. Therefore, the used of modified version of dependency ratio is relatively more informative. Finally, the one period lagged error-correction term ECT_{t-1} has a negative sign and statistically significant at 1 per cent level. This

implies that the variable is not overshooting and hence the long run equilibrium relationship is attainable.

Table 3
The Estimated Short and Long Run Coefficients
Using the ADRL Approach

Dependent Variable: $\ln S_t$

Long Run Coefficients		
Independent Variable	**Coefficient**	**t-Statistic**
Constant	−9.112	−4.111*
$\ln Y_t$	1.755	7.872*
$\ln MDR_t$	2.963	2.209**
R_t	0.011	1.520

Dependent Variable: $\Delta \ln S_t$

Short Run Coefficients		
Independent Variable	**Coefficient**	**t-Statistic**
Constant	−4.440	−4.474*
$\Delta \ln Y_t$	1.480	7.127*
$\Delta \ln MDR_t$	−2.046	−1.250
ΔR_t	−0.006	−2.137**
ECT_{t-1}	−0.487	−3.533*

Notes: The asterisks *, **, and *** denote significance at 1, 5 and 10 per cent level, respectively. The lags order used in ARDL model ($\ln S_t$, $\ln Y_t$, $\ln MDR_t$, R_t) and are selected by SBC.

3.2 *Modified Wald Test Results*

In the earlier section, we found that the variables are cointegrated through the bounds testing procedure. However, the presence of cointegration does not imply causation and policy effectiveness. Therefore, MWALD test is employed in this study to determine the causality direction between savings and its determinants.

Table 4 reports the results of MWALD test. The results reveal that the null hypothesis of income growth does not Granger cause savings can be rejected at the 5 per cent significance level. The null hypothesis of savings does not Granger cause income growth also can be rejected at 5 per cent significance level. These results accord with the Solow's (1956) neoclassical growth theory and also the Romer's (1986) endogenous growth model views that savings is a prominent source for a country's

economic growth through its impact on capital formation. Therefore, policies that encourage savings should be implemented to stimulate Malaysia's economic growth. This finding is not consistent with the work did by Agrawal (2001) and Baharumshah *et al.* (2003).

Table 4
The Results of Modified Wald Test (χ^2-Statistics)

Dependent Variable	$\ln S_t$	$\ln Y_t$	R_t	$\ln MDR_t$
$\ln S_t$	–	8.324**	5.426***	5.983***
$\ln Y_t$	8.782**	–	4.407	6.051**
R_t	20.844*	6.574**	–	13.997*
$\ln MDR_t$	2.232	1.441	0.656	–

Notes: The asterisks *, ** and *** denotes significance at 1, 5 and 10 per cent level respectively. The optimal lag length for the VAR system is selected by AIC.

Table 4 shows that the causal link between savings and interest rates is bilateral. This indicates that interest rates can be used as a macroeconomic management instrument to encourage savings which eventually stimulates Malaysia's economic growth.[7] Furthermore, the MWALD test results show that the dependency ratio Granger cause savings, but not vice versa. This implies that change of non-productive population size including unemployed labour force significantly affects the savings behaviour in Malaysia.

4. Conclusion and Policy Implication

Malaysia is one of the rapid grow nations in the Southeast Asia. Many studies have been devoted to examine the story behind the rapid grow of Malaysia's economy from the past few decades. One of the potential sources that lead to economic growth in Malaysia is the high savings rate. This study was intended to re-investigate the factors that could explain the savings behaviour and high savings rate in Malaysia through the relatively new bounds testing procedure for cointegration and MWALD test. Using the annual data from 1970 to 2004, this study found that the explanatory variables, real income, real interest rate and dependency ratio are moving together with savings to achieve their long

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[7] If interest rate is a choice of policies in Malaysia, reducing interest rate will lead to economic growth because interest rate is also a cost of borrowing. Thus, reducing the interest rate will encourage capital investment that will directly contribute to economic growth. This is consistent with the significant negative effect of real interest rate on savings as presented in Table 3.

run equilibrium relationship, although there may be a deviations in the short run. Furthermore, this study found that the major causes of savings in Malaysia are real income and dependency ratio and that the relationship is elastic. On the other hand, the significance negative effect of real interest rate on savings in Malaysia is inelastic. As a result, monetary policy may not an effective macroeconomic policy instrument to encourage savings in Malaysia although the evidence of MWALD test shows that savings and real interest rate is bilateral causality. Furthermore, the presence of bilateral causal relationship between savings and income growth in Malaysia appears to support the conventional wisdom of savings leads economic growth through its impact of capital formation. This reveals that the savings in Malaysia are mobilised and financed into the productive activities. On the other hand, the uni-directional causal relationship running from dependency ratio to savings in conjunction with high positive elasticity in the long run implied that households in Malaysia are very protective and the high savings rate in Malaysia is leads by the precautionary savings.

In a policy context, this finding may shed some light to the policymaker that savings is a prominent source for Malaysia's economic growth. Therefore, policies that encourage savings should implement to foster economic growth in Malaysia. In addition, the comprehensive development of financial system in Malaysia should be accelerated in order to further mobilise savings and translate it into viable capital formation that would ultimately contribute to economic growth in Malaysia.

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