FINANCIAL DISINTERMEDIATION IN THE 1990s: IMPLICATIONS ON MONETARY POLICY IN MALAYSIA[*]

ANTHONY C.K. TAN

Faculty of Economics & Administration, University of Malaya
Kuala Lumpur 50603, Malaysia
anthonytan_ck@yahoo.com

AND

KIM-LENG GOH[**]

Faculty of Economics & Administration, University of Malaya
Kuala Lumpur 50603, Malaysia
klgoh@um.edu.my

Abstract

The increased financial disintermediation that characterizes the Malaysia's financial system since the early 1990s has contributed towards changes in the dynamics of monetary transmission mechanism. Using quarterly data from 1980:1 to 2005:4, we found a greater effectiveness of monetary policy during the pre-1990:3 period, but the post-1990:3 period poses much difficulty for the conduct of monetary policy. Innovations in the financial market appeared to have led to lower output variability. Further, when the real interest rate is made a function of financial disintermediation, the real interest rate appeared to have lost its significance in influencing real variables in the post-1990:3 period. This study did not, however, find evidence in support of the significance of the real interest rate in affecting real variables through the direct financing channel via the capital market.

Keywords: bank lending channel, capital market, cointegration, VAR.
JEL classification: E44, E52

I. *Introduction*

Modern empirical research in monetary economics places emphasis on the ability of policy to stabilize the macroeconomy (Friedman and Schwartz (1963)). Changes in policy are

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[**] Corresponding author.

important insofar as they affect aggregate real activities in an economy. Essentially, this requires the existence of some form of nominal rigidities which would then allow monetary policy action to be translated into changes in real variables. Accordingly, a well-developed financial system is both crucial and critical for the conduct of monetary policy.

Financial developments over the past three decades, including capital market developments (as an alternative to bank intermediation), appear to have had altered the conduct of monetary policy both in developed and developing economies. Accordingly, it has been argued specifically by several authors that monetary policy appears to have less of an impact on real activity than it once had (see Hussein (1992), Azali (1998), Ghazali (1998), Boivin and Giannoni (2002), Kuttner and Mosser (2002)). In spite of this general observation, the cause of this perceived change however, remains an open question. Structural changes affecting the monetary transmission mechanism and the conduct of monetary policy may in fact be non-financial in nature as it is often posited.

The Malaysian financial system is becoming more market-based, as both public and private sectors are increasingly relying on direct (or market) financing, as compared to indirect (or intermediated) financing. Indeed, the amount of funds raised from the Malaysian capital market has increased significantly from 1980 to 2005. The annual growth rate of new listings on Bursa Malaysia over the 26-year period is on average, approximately 20.7%. Following positive developments in the Malaysian capital market[1], there is now greater access to debt and equity markets for large classes of borrowers and investors. The process of financial disintermediation — defined here as the process where deficit financial units, in meeting their financing needs, bypass financial institutions in favour of the capital markets has become more apparent.

On one hand, firms (notably large corporations) are no longer constrained by the banking system for funds. On the other hand, it is acknowledged that access to the capital market for funds, even with better information dissemination may not be readily available, particularly to the small and medium-size enterprises (SMEs). It is more likely that these SMEs would still rely heavily on the banking system for their financing needs. In Malaysia, the banking system is an important conduit for the propagation of monetary transmission mechanism. A study by Mansor (2002) confirms that bank deposits and bank loans served as important channels of monetary transmission mechanism during the pre-Asian crisis for Malaysia. Unfortunately, it is not readily apparent from this study whether the increasing trend of financial disintermediation implies a weakening of the bank lending channel of monetary transmission mechanism.

Apart from financial disintermediation, another recent trend that could be observed in Malaysia is the significant growth of mortgage and consumption credit extended by the banking system to the household sector — defined as individual persons and non-incorporated businesses. In response to the increased financial disintermediation, banking institutions have shifted their composition of lending from large corporations to SMEs, including the more aggressive marketing of consumer and mortgage loans. The rising households' medium- to long-term debts which are financed by the banking system suggest that the strength of the bank lending channel in Malaysia may not likely be compromised.

According to Sellon, Jr (2002), the above finding is consistent with that of the US economy, where the fall in direct corporate borrowings brought about by the process of financial disintermediation has likely been offset by an increase in household borrowings. This

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[1] See Attila (2000) for an overview of the Malaysian capital market as well as recent developments.

implies that the optimal design and conduct of monetary policy since the 1990s could no longer be undertaken without taking into account capital market developments, as well as greater household access to medium- to long-term financing.

The effect of changes in the real interest rate elasticity of the output gap offers an indicative measure of the effectiveness of monetary policy. From the New-Keynesian interest rate point of view, the effect of such changes is ambiguous. On one hand, increased financial disintermediation may have removed or limited the impact of an exogenous monetary policy shocks on real activity. In this respect, Bean, Larsen and Nikolov (2002) argued that when firms and/or households are unconstrained in their access to credit, output (and inflation) tends to be stabilized because firms and households would become relatively less sensitive to current and/or expected changes in economic conditions. Following rational expectations as well as the optimizing behaviour of firms and households, innovations in the policy interest rate will tend to affect output (and inflation) thereby lowering the persistence of moves in the output gap. In view that real interest rates are an important monetary policy tool in the transmission mechanism, Bean, Larsen and Nikolov (2002) further argued that a reduction in financial friction associated with financial deepening and/or disintermediation would tend to lower the persistent movements in the output gap, leading to a corresponding reduction in the real interest rate elasticity of the output gap (referred hereafter as real interest rate elasticity). Accordingly, this argument suggests a lower amplification of an exogenous monetary policy shock. On the other hand, the greater degree of access to interest-sensitive financial assets and liabilities may have exposed firms and households directly to interest rate fluctuations. To the extent that a larger segment of borrowers and lenders may now be directly affected by interest rate variations, the output gap will appear to be relatively more interest-sensitive.

A monetary policy tightening is likely to cause credit-constrained firms (whom are mostly bank-dependant) to face difficulty in securing financing for their investments even from the banking system since the problem of adverse selection is likely to be more pronounced. And since 'state verification' is costly, lenders would therefore demand an external finance premium to compensate them for this 'state verification' cost. Empirical evidence in the US suggests that the estimated premium on external finance was found to be very low during expansionary period in 1997 to 1999, but rose sharply in 2000 especially for higher-geared firms (Levin, Natalucci and Zakrajsek (2004)). In general equilibrium, such a setting/mechanism has the potential to provide amplification and propagation, causing persistent movements in firms' average cost of capital and therefore investment (and ultimately, output).

Essentially, the presence of financial frictions in the form of credit market imperfections could potentially cause greater output persistence[2] since the costly external finance premium would discourage firms from investing and/or caused their investments to be limited to the availability of their retained profits. In this respect, the presence of financial market frictions would likely produce an IS curve with relatively greater real interest rate elasticity since the output gap tends to be larger as compared to a model that is free from financial frictions.

Although the argument that greater financial disintermediation may imply the reduction of

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[2] Of course, as argued by Clarida, Gali and Gertler (1999), there may be a number of features other than financial frictions which can generate endogenous output persistence. For example, investment adjustment costs and habit persistence in consumption can both be used to motivate output persistence. But even so, the presence of financial frictions should generate greater output persistence than in their absence.

financial frictions, leading to a likely reduction in real interest rate elasticity, it is also plausible that the concept of 'relationship-banking' renders bank lending less interest rate sensitive as compared to direct financing (which is relatively more price-sensitive). Bank lending contracts are implicit in nature, allowing the possibility of re-negotiation and risk-sharing amongst the parties involved — a feature that may not be reflected in market interest rates. Consequently, greater financial disintermediation would tend to increase the real interest rate elasticity following an exogenous monetary policy shock. Given that there is no theoretical consensus on the effect on real interest rate elasticity, the net impact of these changes could only be assessed empirically.

Studies on monetary transmission mechanism in Malaysia either concentrated on the traditional interest rate, monetary aggregate and/or following Bernanke and Blinder (1988) and Bernanke and Gertler (1995)[3], to include credit aggregates as channels of monetary transmission mechanism (see Tan and Cheng (1995), Masih and Masih (1996), Azali and Matthews (1999), Shanmugam, Nair and Ong (2003)). An analysis of monetary policy, and therefore its effectiveness in influencing target variables could benefit from a more explicit consideration of the evolving role of the Malaysian capital market.

Studies on monetary transmission in Malaysia have focused solely on variables such as bank loans (Mansor (2002)) or commercial banks' claim on the private sector (Azali and Matthews (1999)), without drawing specific reference to variables characterizing capital market developments. Further, there does not seem to appear studies that explicitly consider the implication of monetary policy on the output gap following the increasingly important role played by the Malaysian capital market. Although Ghazali (1998) concluded the case for a reduced effectiveness of monetary policy in light of financial liberalization and innovation for Malaysia, this particular study evaluated the response of banks' portfolio allocation (notably bank loans) to changes in Bank Negara Malaysia's (BNM, or the Central Bank of Malaysia) policy interest rate. The scope of study has been rather limited.

The effect of changes in BNM's policy interest rate, specifically the degree of interest rate pass-through to the commercial banks' average lending rate has also been documented for Malaysia. The recent study by Tee (2001), however, requires further examination. The degree of interest rate pass-through, despite rapid in Malaysia, could be extended to account for the implications on the output gap. Of primary importance is whether the decline in output following a positive monetary policy shock is often followed by greater or lesser degree of output persistence.

This paper seeks to examine the effect of financial disintermediation — the greater use of market-based financing by the non-financial corporate sector (NFCS) and the change of composition of lending by the banking institutions — on the conduct of monetary policy. It is of particular interest to ascertain whether the increased financial disintermediation since the early 1990s could have had altered the propagation of monetary transmission mechanism. We wish to investigate whether (1) the dynamics of monetary transmission mechanism have

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[3] In examining monetary transmission mechanism, Bernanke and Gertler (1995) argued specifically that it is difficult to explain the magnitude, timing and composition of the economy's response to monetary policy solely in terms of the traditional interest rate channel of monetary transmission mechanism. Consequently, the broad credit channel (namely bank lending and balance-sheet channels) was proposed by Bernanke and Blinder (1988) to fill the gap in the traditional story.

changed following structural changes that took place in the early 1990s, and (2) whether the increased financial disintermediation that took place during the corresponding period is amongst one of the causes of this change.

The remaining sections are organized as follows. Section II introduces the data and data sources, model specifications, as well as the methodology of analysis. This is followed by Section III where we report the results of our findings. Finally, in the concluding section, we present a discussion as well as evaluation of the results obtained, including some policy recommendations.

II.   *Methodology of Study*

In order to test our propositions, quarterly data from 1980:1 to 2005:4 are analyzed. In view of the apparent increased in capital market activities and rising household mortgage and consumption credit since the early 1990s, two sub-periods are distinguished. The first sub-period is from 1980:1 to 1990:3, while the second sub-period is from 1990:4 to 2005:4. In determining these two sub-periods, the breakpoint — 1990:3 has been determined exogenously by examining the patterns revealed by the indicators of financial disintermediation (see discussions below). The variables and the specification of the variables, as well as the sources of data used in our study are set out in Tables 1 and 2 respectively.

The methodology employed in our study involves two stages. In Stage One, our objective is to ascertain whether there has been a change in the dynamics of monetary transmission mechanism since the early 1990s. In Stage Two, we introduce a simple structural IS model in order to examine whether increased financial disintermediation has contributed to changes in the dynamics of monetary transmission mechanism in Malaysia.

1.   Stage One Analysis

Firstly, we examine whether there has been a change in the characterization of monetary transmission mechanism following increased capital market activities, beginning in the early 1990s. For this purpose, we employ a VAR to track such changes.

Before proceeding with the VAR estimation, unit root tests of the time series variables are first conducted in order to ascertain their stationarity properties. We subject the said variables to a unit root test following Dickey and Fuller (1979), and Phillips and Perron (1987). In these widely applied stationarity tests, it is acknowledged that these conventional unit root tests do not allow for the existence of structural break in the time series. Accordingly, an adjusted augmented Dickey-Fuller (ADF)-type unit root test as proposed by Perron (1989) is applied to the full sample period and the second sub-period in order to account for the possible break in trend following Malaysia's adoption of capital control and pegged exchange rate regime on 1 September 1998.

Possible changes in both the intercept and slope of the trend function are therefore considered, and the regression to test the null hypothesis for the presence of a unit root is of the form,

$$y_t = \mu + \beta t + \theta_1 DU_t + \theta_2 DT_t + \gamma y_{t-1} + \sum_{i=1}^{m} \phi_i \Delta y_{t-i} + \varepsilon_t$$

TABLE 1.   DESCRIPTION OF VARIABLES

No.	Variables	Proxy variables
1.	Output	Nominal GDP deflated by the aggregate price level, y_t (in logarithmic) Due to the non-availability of GDP data from 1980:1 to 1990:4, it was necessary to interpolate the annual series on real GDP to obtain quarterly series based on Goldstein and Khan (1976).
2.	Output gap	Difference between actual output and potential output where potential output is a Hodrick-Prescott filtered version of the actual output series, $ygap_t$ (in percentage) The Hodrick-Prescott filter (HP-filter) is used as a detrending method. Supposing the original series y_t is composed of a trend component, g_t and a cyclical component, c_t, i.e. $y_t = g_t + c_t$. The HP-filter isolates the cyclical component by solving the following minimization problem. $$\sum_{t=1}^{T}(y_t - g_t)^2 + \sum_{t=2}^{T-1}[(g_{t+1} - g_t) - (g_t - g_{t-1})]^2$$ The output gap is computed as $ygap_t = (y_t - g_t)/y_t$
3.	Expected output gap	Difference between forecasted output and the potential output, $ygapf_t$ (in percentage) Forecasted output is computed based on the estimated output equation from the extended VAR (see discussion below). Expected output gap is then computed in the manner similar to the output gap, i.e. as the percentage difference in the forecasted output series from the detrended output series.
4.	Aggregate price level	Consumer Price Index (CPI) (1980 = 100), P_t (in logarithmic)
5.	Nominal interest rate	Average BNM's 3-month Inter-bank Rate/Overnight Policy Rate (OPR), i_t
6.	Real interest rate	Ex-post/actual real interest rate, r_t $r_t = i_t = E_t(\pi_{t+1})$, where expected inflation, $E_t(\pi_{t+1}) = \pi_t + \mu_t$, $E(\mu_t) = 0$
7.	Nominal exchange rate	Average bilateral exchange rate between US and Malaysia, s_t (in logarithmic) (expressed as units of domestic currency per unit of foreign currency)
8.	Loans to the NFCS (in real terms)	Loans disbursed by the banking system to the NFCS, deflated by the aggregate price level, BL_t
9.	Household mortgage credit (in real terms)	Outstanding household mortgage extended by the banking system to the household sector, deflated by the aggregate price level, M_t (in logarithmic)
10.	Household consumption credit (in real terms)	Outstanding household consumption credit extended by the banking system to the household sector, deflated by the aggregate price level, C_t (in logarithmic)
11.	Financial disintermediation (i) Measure of financial disinter-mediation activities (ii) Measure of both financial dis-intermediation activities and growth in household mortgage and consumption credit	· Ratio of NFCS' direct to indirect financing[i], $F2_t$ · Ratio of NFCS' direct financing to total loans disbursed by the banking system[ii], $F3_t$
12.	Measure of broad external finance premium	Difference between the average lending rates extended by commercial banks to the NFCS and the average discount rate of the 3-month Treasury Bill, D_t
13.	Time	Quarterly, t

Notes:
(i)   Direct financing refers to gross equity & private debt securities raised by the NFCS from the capital market, while indirect financing refers to total loans disbursed by the banking system to the NFCS.
(ii) Loans disbursed by the banking system[a] primarily comprise loans to the (1) agriculture, hunting, forestry and fishing sector, (2) mining and quarrying sector, (3) manufacturing sector, (4) utility sector (electricity, gas, water), (5) wholesale and retail trade, (6) broad property sector (excluding purchase of residential property), and (7) finance, insurance and business services sector.
    [a]   *Loans disbursed to the government are included, however these loans are insignificant, representing less than one percent of the total loans disbursed by the banking system.*

TABLE 2.   SOURCES OF DATA

No.	Data	Sources of data
1.	Nominal GDP	Department of Statistic's Monthly and Quarterly Bulletins (various issues); and IMF's International Financial Statistics
2.	CPI	Department of Statistic's Monthly and Quarterly Bulletins (various issues)
3.	3-month Inter-bank Rate	BNM's Monthly Statistical Bulletins and Quarterly Economic Bulletins (various issues)
4.	Discount rate of 3-month Treasury Bill	
5.	Commercial banks' lending rate	
6.	Nominal exchange rate (units of domestic currency per unit of foreign currency)	
7.	Loans disbursed to the NFCS	
8.	Household mortgage credit	
9.	Household consumption credit	
10.	NFCS' direct financing; NFCS' indirect financing; and Total loans disbursed by the banking system to the NFCS	BNM's Monthly Statistical Bulletins and Quarterly Economic Bulletins (various issues), and Bursa Malaysia

where the dummy variables, DU_t and DT_t are defined as:

$$DU_t = \begin{cases} 1 \ if \ t > T_B \\ 0 \ otherwise \end{cases}, \ DT_T = \begin{cases} t \ if \ t > T_B \\ 0 \ otherwise \end{cases}$$

and the breakpoint, T_B is fixed at 1998:2.

The asymptotic distribution of the t-statistic for testing $H_0: \gamma = 1$ is given by Perron (1989). The percentage points for the test are dependent on the value of the break function, T_B/n. The optimal lag length for the unit root testing is determined by the Schwarz Information Criterion (BIC). Once the stationarity properties of the variables are established, it is necessary to decide the particular specification of the VAR. Estimate of VARs in levels run the risk of being spurious if the variables are integrated series or non-stationary. On the other hand, VARs specified in differences when the variables are non-stationary will generate efficient estimates, but at the cost of ignoring potential long-run relationship(s) that is/are of importance.

Engle and Granger (1987a, b) demonstrated that if the variables under consideration are cointegrated, i.e. long-run equilibrium relationship exists amongst the variables, the dynamic relationship between those variables could be more appropriately represented using VECM. VECM is essentially a restricted VAR model which imposes long-run constraints among levels of the variables as implied by their cointegration. They also demonstrated that the aforementioned long-run constraints are also satisfied asymptotically in an unrestricted VAR. Provided that the variables under consideration are cointegrated, both approaches of restricted and unrestricted VAR are appropriate for modeling the dynamic interaction among the time series variables. In view that monetary transmission mechanism is a short to medium-run phenomenon, it is therefore not surprising that estimating the unrestricted VAR for the cointegrated variables seemed to be the normal route taken by many researchers in the

literature.[4]

Likewise, for the purpose of our study, we approach the issue of non-stationarity by pre-testing the variables in our study using the Johansen (1988) procedure for multivariate cointegration for both the baseline and the extended models (see further explanation below). In the event cointegration is found, we shall then proceed to estimate an unrestricted VAR (hereafter referred to as VAR) in levels as our preferred VAR specification. VAR models are commonly used since this technique allows the researcher to address the simultaneity problem associated with the effects of monetary policy. The fact that monetary authority would usually loosen policy when the economy weakens and tightens when the economy strengthens shows that such endogenous response of policy to economic conditions is one reason why it is difficult to identify the effects of policy.

The evolution of the vector X_t which contains the macroeconomic variables whose behaviour we seek to understand, depends both on unexpected disturbances, u_t and on a systematic component, $A_0 + A_1 X_{t-1} + A_2 X_{t-2} + \ldots + A_p X_{t-p}$, that determines how shocks are propagated to the rest of the economy. The estimates of A_0, A_1, A_2, ..., A_p are obtained by applying ordinary least squares (OLS), and the estimate of $E(u_t u'_t)$ is given by the sample variance-covariance matrix of the OLS residuals. In this study, we consider a baseline VAR containing four variables. The vector of X_t is given by,

$$X_t = [\; i_t,\; s_t,\; y_t,\; P_t]'$$

The above baseline VAR contains the minimum set of variables that are crucial for any discussion of monetary policy. The model is then extended with a block of financial variable to reflect the prominence of the credit channel in the literature of monetary transmission mechanism. These variables are namely (1) quantity variables, BL_t (capturing the role/existence of alternative sources of financing for the NFCS), M_t and C_t (capturing increased households access to medium- to long-term financing), and (2) price variable, D_t — capturing the role of broad external finance premium in the monetary transmission mechanism. Our augmented vector $X_t^{augmented}$ is therefore,

$$X_t^{augmented} = [i_t,\; s_t,\; D_t,\; BL_t,\; C_t,\; M_t,\; y_t,\; P_t]'$$

To determine whether there is statistical evidence in support of significant changes in the characterization of monetary transmission mechanism beginning in the early 1990s, we subject the estimated individual equations in both the baseline and extended VARs to a Chow's breakpoint test following Chow (1960). The F-statistic has an exact finite sample F-distribution under the null hypothesis of no structural change assuming if the errors are independent and identically distributed normal random variables. In the event the estimated individual equations corresponding to each of the sub-period is found to be significantly different (in accordance with our expectations), the VARs for both sub-periods are estimated. Estimation of VAR for the individual sub-periods is necessary to distinguish the dynamics of output during the pre- and post-1990:3 period.

Our main analysis focuses on how output responds to interest rate shocks using the

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[4] VARs in levels can be reparameterized in the form of an error correction model if the variables in the system are cointegrated. Empirical studies on monetary transmission mechanism that adopted the unrestricted VARs in levels are amongst others, Bernanke and Blinder (1992), Dale and Haldane (1995), and Ramaswamy and Sloek (1998).

impulse response function and variance decomposition derived from the VARs.[5] The propagation of monetary policy is examined through changes in interest rate because the monetary policy operational framework of the Central Bank has evolved from targeting monetary aggregates to targeting short-term interest rates, amidst the progressive transition towards a more market-based monetary policy implementation procedure. The shift in the monetary policy strategy since the early 1990s towards interest rate targeting was motivated by the fact that liberalization of interest rates since 1978, coupled with deregulation and liberalization measures resulted in the disassociation between monetary aggregates and the Central Bank's ultimate objective of maintaining price stability. Furthermore, there was clear structural change in the financing pattern of the economy, whereby the financing pattern shifted from interest-inelastic market (notably, the Malaysian Government Securities (MGS) market) towards a relatively more interest-sensitive market such as bank credit and capital market instruments (see Bank Negara Malaysia, 1999).

Since VARs are essentially in reduced-form, it is not possible to ascertain the interest rate elasticities. As a result, specific tests involving structural framework are required, thereby motivating the stage two analysis.

2. Stage Two Analysis

In the New Keynesian tradition, models of monetary policy are grounded in dynamic general equilibrium theory and capture the forward-looking behaviour of optimizing firms and consumers (for example, see Clarida, Gali and Gertler (1999)). A baseline New Keynesian IS equation[6] that abstracts from investment and capital accumulation is of the following form:

$$ygap_t = E_t(ygap_{t+1}) - \varphi[i_t - E_t(\pi_{t+1})] + \delta_t \qquad (1)$$

where $ygap_t$ is the percentage deviation from a long-run trend.

According to Roldos (2006), the empirical performance of the New Keynesian-type models are not satisfactory and backward elements have to be added to achieve a reasonable fit. Accordingly, we modify Equation (1) above to include lagged endogenous persistence, as well as replacing the current real interest rate with the lagged real interest rate. The specification of the baseline IS equation therefore takes the following form:

$$ygap_t = \alpha_0 + \alpha_1 E_t(ygap_{t+1}) + \alpha_2 ygap_{t-1} + \alpha_3 r_{t-1} + \varepsilon_t \qquad (2)$$

In view that we are studying the effects of increased financial disintermediation beginning

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[5] Some recent works (for example, Gorden and Leeper (1994), Christiano, Eichenbaum and Evans (1996) and Sims and Zha (1998)), imposed restrictions on the structural component of the error terms in the VAR system to identify monetary policy shocks. This approach, known as structural VAR or SVAR, is able to produce dynamic impacts of monetary policy that are consistent with traditional analyses. The identification of shocks has another advantage in isolating the unpredictable component from systematic changes in the policy. As the aim of this study is to examine the effect of financial disintermediation, a comparative analysis through the VAR baseline and extended models is necessary. Identification of the sources of shocks is of less importance as our focus is mainly on interest rate changes given that short-term interest rate targeting is the key policy instrument of the Central Bank. Further, as is reported in the following section, the results are largely consistent with the traditional analyses.

[6] This equation is often called the intertemporal IS equation, which lies at the core of many recent macroeconomic models. It is similar to the traditional Keynesian-type IS equation, in the sense that it relates output negatively to the real interest rate. However, output is also affected by expected future output, as consumers tend to smooth their

in the early 1990s, we extend the baseline IS equation to account for these recent trends. To accomplish this, we make the coefficient, α_3 a function of the variables, *F2* and *F3,* following the approach of Roldos (2006) i.e.

$$ygap_t = \alpha_0 + \alpha_1 E_t(ygap_{t+1}) + \alpha_2 ygap_{t-1} + (\alpha_{31} + \alpha_{32} Fi_t)r_{t-1} + \varepsilon_t \qquad (3)$$

where $i = 2, 3$. The coefficient α_{31} indicates the real interest rate elasticity (or more accurately, the elasticity of intertemporal substitution in consumption), and this measure will serve as a guide for examining the effectiveness of monetary policy in Malaysia in recent years.

In order to examine the dynamism of longer lag lengths in our model, we extend the IS equation by increasing the number of lag lengths to two. The specification of Equations (2) and (3) now takes the following form:

$$ygap_t = \beta_0 + \beta_1 E_t(ygap_{t+1}) + \sum_{j=1}^{2} \beta_{2j} ygap_{t-j} + \sum_{j=1}^{2} \beta_{3j} r_{t-j} + \varepsilon_t \qquad (4)$$

$$ygap_t = \beta_0 + \beta_1 E_t(ygap_{t+1}) + \sum_{j=1}^{2} \beta_{2j} ygap_{t-j} + \sum_{j=1}^{2} (\beta_{3j,i} + \beta_{4j,i} Fi_t)r_{t-j} + \varepsilon_t \qquad (5)$$

where $i = 2, 3$. The real interest rate elasticities in Equations (4) and (5) are now given by $\beta_{31} + \beta_{32}$ and $\beta_{31,i} + \beta_{32,i}$, respectively. The re-estimated coefficient capturing the interaction effect of the real interest rate and measures of financial disintermediation in the extended IS Equation (5) is given by $(\beta_{41,i} + \beta_{42,i})$.

We propose to estimate the IS equations by applying OLS and employ the coefficient restriction test in order to determine whether the real interest rate elasticities, and the coefficient capturing the interaction effect of real interest rate and measures of financial disintermediation are statistically significant. However, tbe baseline model as given in Equation (2) has the specification of a typical hybrid IS curve as discussed by Christiano, Eichenbaum and Evans (1999), Erceg, Henderson and Levin (2000) and Smets and Wouters (2002), and thereby suffers from severe autocorrelation.[7] To circumvent the problem, these equations are also estimated using GMM. To obtain robust estimates, the moment conditions are weighted by heteroskedasticity and autocorrelation robust variance covariance matrix. The variance covariance matrix is obtained using the Bartlett kernel weights with Newey and West's fixed bandwidth selection criterion.[8]

III. *Results and Data Analysis*

Figure 1 shows the trend of financial disintermediation measured using two indicator ratios

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consumption over time. An equation of this form can be obtained as a log-linear approximation to a consumption Euler equation in a fairly large variety of models. The coefficient φ is more appropriately known as the elasticity of intertemporal substitution in consumption, and δ represents unforecastable demand shocks.

[7] In a typical hybrid IS curve, $\varepsilon_t = 1/(1+h)E_t(\varepsilon_{c,t} - \varepsilon_{c,t+1})$ where h is the coefficient of consumption habit formation and $\varepsilon_{c,t}$ represents the preference shocks. We thank the referee for alerting us on this point and for the suggestion on the solution.

[8] Eviews Version 5.1 is used for estimations. We kept the variables predetermined at period t and measures of financial disintermediation as instrumental variables. We included $E_{t-2}(ygap_{t-1})$ as an additional instrumental variable to remedy the problem of $cov(\varepsilon_t, E_t(ygap_{t+1})) \neq 0$.

FIG 1.   INDICATORS OF FINANCIAL DISINTERMEDIATION IN MALAYSIA, 1980:1 — 2005:4
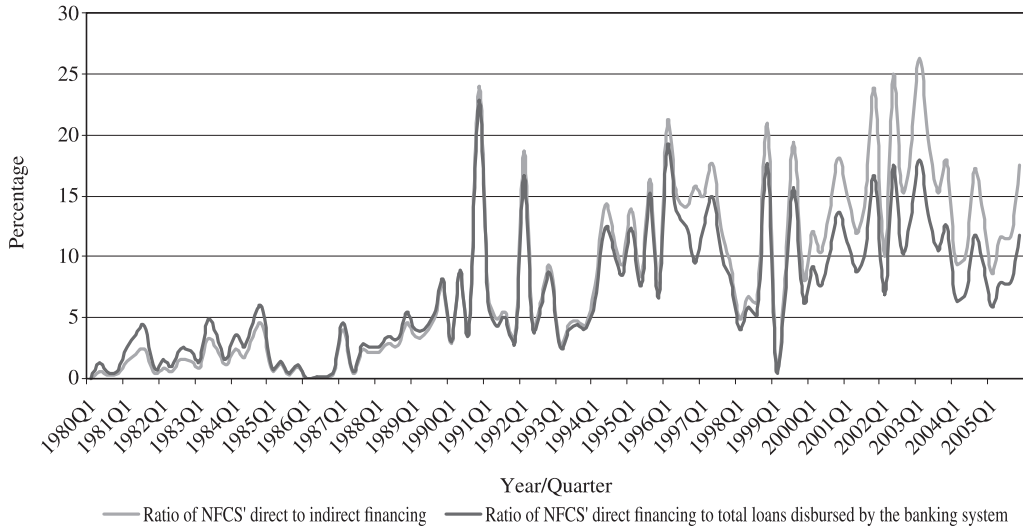
Source : Bank Negara Malaysia and Bursa Malaysia

for the period 1980 to 2005 (see Item 11 of Table 1 for further explanation of these indicators). It is evident from Figure 1 that the trend of financial disintermediation is on the rise. The ratios are larger in magnitude since 1990:3, an indication of a likely reduction in financial frictions beginning in the 1990s. Thus, the post-1990: 3 period is characterized by increased capital market activities.[9]

1.   Stage One Results

Table 3 reports the results for the augmented Dickey-Fuller (ADF) as well as the Phillips and Perron (PP) unit root tests when both the constant and trend component are present in the test regression. For the full sample period, both ADF and PP tests do not provide evidence against the presence of unit roots in the level variables. However, the test on the first differences indicates strong rejection of the null hypothesis at 1% level, implying that these variables are integrated of order one, $I(1)$. These time-series properties remain the same for both the sub-periods. According to Choi (1992), the PP test is relatively more powerful than the ADF test in finite samples. Indeed, in our study, the PP test is able to reject the null hypothesis in several cases where the results of the ADF test are found to be insignificant (see for example, $ln\ BL_t$, $ln\ M_t$ and $ln\ C_t$). For further analysis of the variables of interest, we shall henceforth rely on the results of the PP test in our study.

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[9] The process of financial disintermediation has become more apparent especially with the establishment of Rating Agency Malaysia Berhad (RAM) in November 1990, the Securities Commission (SC) in March 1993 and the Malaysian Rating Corporation Berhad (MARC) in October 1995.

TABLE 3.   SUMMARY OF THE ADF AND PP UNIT ROOT TEST STATISTICS

(a) Full sample

Variables	Level		First difference	
	ADF[(a)]	PP	ADF[(a)]	PP
$ln\ y_t$	-3.5768	-2.3801	-5.8749 **	-10.9857 **
$ln\ s_t$	-1.8011	-2.1999	-6.6054 **	-6.7678 **
i_t	-3.0571	-3.5770 *	-6.2713 **	-11.1369 **
$ln\ P_t$	-1.8589	-1.6225	-5.0421 **	-8.1416 **
r_t	-2.0510	-3.3681	-8.6290 **	-13.0332 **

(b) 1st and 2nd sub-period

Variables	1st sub-period				2nd sub-period			
	Level		First difference		Level		First difference	
	ADF	PP	ADF	PP	ADF[(b)]	PP	ADF[(b)]	PP
$ln\ y_t$	-2.3493	-1.7467	-2.7751 ^	-2.8654 ^	-2.9099	-2.5107	-5.6585 **	-9.6555 **
$ln\ s_t$	-2.7926	-2.6401	-5.5196 **	-6.8925 **	-2.1386	-2.1031	-5.9997 **	-4.7852 **
i_t	-3.0641	-3.4082	-7.5149 **	-8.4153 **	-4.1316	-2.2947	-5.3398 **	-4.9099 **
$ln\ P_t$	-3.0734	-2.6702	-5.3483 **	-5.6483 **	-3.0571	-1.5368	-7.0869 **	-5.7687 **
r_t	-2.3354	-2.3152	-6.9362 **	-7.4725 **	-3.4032	-2.1658	-4.9718 **	-6.4560 **
D_t	-0.4736	-0.6354	-5.3212 **	-5.2695 **	-2.9242	-2.8411	-3.8776	-9.6046 **
$ln\ BL_t$	-2.0371	-2.4972	-2.6443	-5.8125 **	-1.9230	-0.8090	-2.9691	-4.7517 **
$ln\ M_t$	-1.6403	-1.2693	-10.2821 **	-10.5444 **	-2.8254	-2.0807	-4.7034	-13.9640 **
$ln\ C_t$	-1.7057	-1.1199	-4.7906 **	-5.0959 **	-2.7356	-2.0917	-3.6587	-3.8642 *

Notes:

^ (*) (**)   Denotes significance at the 10% (5%) and (1%) level.

\#   The ADF test applied to all the variables are assumed to follow a process containing a drift and a deterministic time trend.

(a)   The 5% and 1% critical values (for $\lambda \approx 0.3$) for the adjusted ADF-type unit root test are -4.17 and -4.78 respectively (Perron, 1989).

(b)   The 5% and 1% critical values (for $\lambda \approx 0.5$) are -4.24 and -4.90 respectively (Perron, 1989).

TABLE 4.   SUMMARY OF THE JOHANSEN COINTEGRATION TEST: FULL SAMPLE

No. of cointegrating relations	Trace statistics	
	Baseline model	Extended model
None	62.4371 **	271.0997 **
At most 1	29.0037	199.9442 **
At most 2	7.3040	137.9139 **
At most 3	1.7313	77.5240 **
At most 4	-	39.8331

Notes:

*(**) Denotes significance at 5% (1%) level.

Following the above, the Johansen (1988) framework is adopted to investigate if long-run linkages exist empirically amongst the variables in our study. The Johansen's method is achieved by testing the null hypothesis of at least r cointegrating vectors against a general hypothesis of more than r cointegrating vectors. The lag length that minimizes the BIC is once again used for this purpose. The result of the Johansen cointegration test is reported in Table 4.

TABLE 5.   SUMMARY OF THE FULL SAMPLE VAR MODEL

(a) Baseline model

Independent variable	Dependent variable			
	i_t	$ln\ s_t$	$ln\ y_t$	$ln\ P_t$
Constant	9.8659 *	-0.0390	-0.0824	0.1441 **
	(1.6823)	(-0.2288)	(-0.4768)	(4.4824)
i_{t-1}	0.7729 **	0.0007	-0.0041 **	0.0005
	(13.4544)	(0.3930)	(-2.3990)	(1.6127)
$ln\ s_{t-1}$	-1.5674	0.9586 **	-0.0696 *	0.0172 **
	(-1.1992)	(25.2593)	(-1.8068)	(2.3974)
$ln\ y_{t-1}$	3.6014 **	0.0636	0.9368 **	0.0398 **
	(2.5633)	(1.5588)	(22.6200)	(5.1614)
$ln\ P_{t-1}$	-8.7622 *	-0.1137	0.1690	0.8870 **
	(-2.3080)	(-1.0311)	(1.5101)	(42.6113)

(b) Extended model

Independent variable	Dependent variable							
	i_t	$ln\ s_t$	D_t	$ln\ BL_t$	$ln\ C_t$	$ln\ M_t$	$ln\ y_t$	$ln\ P_t$
Constant	42.7418 **	0.0608	-10.4969	-0.2682	-2.4648 **	1.1807	0.2872	0.2178
	(2.6851)	(0.1209)	(-1.1417)	(-0.7322)	(-3.8777)	(1.1039)	(0.5713)	(2.3270)
i_{t-1}	0.6056 **	0.0001	0.0569	0.0011	0.0032	-0.0121 **	-0.0072 **	0.0000
	(7.5906)	(0.0470)	(1.3509)	(0.6668)	(1.1029)	(-2.4642)	(-3.1290)	(0.1416)
$ln\ s_{t-1}$	-1.4529	0.9916 **	-1.3091 *	-0.0911 **	-0.0207	0.0778	-0.1129 **	0.0236 **
	(-1.0435)	(24.6407)	(-1.7803)	(-3.1082)	(-0.4072)	(0.9098)	(-2.8064)	(3.1537)
D_{t-1}	0.1847	0.0096 **	0.8577 **	-0.0034	-0.0243 **	0.0021	-0.0103 **	0.0010 *
	(1.4064)	(2.5217)	(12.3658)	(-1.2479)	(-5.0733)	(0.2617)	(-2.7183)	(1.4710)
$ln\ BL_{t-1}$	-1.0665	0.0258	0.0430	0.9693 **	0.0426	-0.1416 *	0.0007	-0.0178 *
	(-0.9973)	(0.8343)	(0.0761)	(43.0792)	(1.0912)	(-2.1559)	(0.0237)	(-3.0931)
$ln\ C_{t-1}$	2.9053 **	0.0006	-0.3531	-0.0123	0.7953 **	0.1570 **	0.0281	0.0126
	(2.6778)	(0.0201)	(-0.6161)	(-0.5383)	(20.0744)	(2.3556)	(0.8963)	(2.1544)
$ln\ M_{t-1}$	-2.0212 **	-0.0069 *	0.0533	-0.0150	0.0437	0.7847 **	-0.0255	-0.0140 *
	(-2.5112)	(-0.2987)	(0.1253)	(-0.8884)	(1.4874)	(15.8687)	(-1.0950)	(-3.2390)
$ln\ y_{t-1}$	0.5503	0.1798 **	-0.7982	0.1305 **	0.3577 **	-0.2726 *	0.7319 **	0.0398 **
	(0.2221)	(2.5100)	(-0.6099)	(2.5030)	(3.9533)	(-1.7908)	(10.2270)	(2.9850)
$ln\ P_{t-1}$	-8.9800	-0.4318 **	4.5064	-0.0732	-0.0186	0.7516 *	0.5346 **	0.9124 **
	(-1.5403)	(-2.5628)	(1.4636)	(-0.5968)	(-0.0874)	(2.0985)	(3.1752)	(29.1032)

Notes:
*(**) Denotes significance at 5%(1%) level.
The numbers in parentheses are the *t*-statistics.

The result indicates that at 1% level of significance, one cointegrating vector is found for the baseline model, while three cointegrating vectors are found for the extended model. In view of the existence of long-run empirical relationship amongst the variables in both models, the dynamic interactions are henceforth analyzed more appropriately using VAR in levels. Both the baseline and extended VARs are estimated and the BIC consistently choose lag length of order one for both the models. The results are summarized in Table 5. At least one of the variables lagged one period is significant in explaining the dependent variable. In the output equation,

TABLE 6.   SUMMARY OF THE CHOW'S BREAKPOINT TEST

Equation	F-Statistics (Break point — 1990:3)	
	Baseline VAR	Extended VAR
i_t	3.6329 **	3.2557 **
$ln\ s_t$	2.9165 *	4.7534 **
D_t	-	2.3706 *
$ln\ BL_t$	-	1.5856
$ln\ C_t$	-	4.3850 **
$ln\ M_t$	-	4.7890 **
$ln\ y_t$	7.8250 **	5.7104 **
$ln\ P_t$	4.8010 **	3.7506 **

Notes:
^ (*) (**)   Denotes significance at 10% (5%) and (1%) level.

TABLE 7.   SUMMARY OF THE BASELINE VAR ESTIMATES

(a) 1st sub-period

Independent variable	Dependent variable			
	i_t	$ln\ s_t$	$ln\ y_t$	$ln\ P_t$
Constant	-21.2514	-0.2772	0.3810	0.4133 **
	(-0.8807)	(-0.8328)	(1.1886)	(3.6128)
i_{t-1}	0.5764 **	0.0003	-0.0045 **	0.0000
	(4.4971)	(0.1693)	(-2.6236)	(0.0443)
$ln\ s_{t-1}$	-12.0675 **	0.9023 **	0.0063	0.0210
	(-2.4290)	(13.1672)	(0.0949)	(0.8897)
$ln\ y_{t-1}$	5.1334	0.0833 *	0.9979 **	0.0259
	(1.4673)	(1.7258)	(21.4766)	(1.5619)
$ln\ P_{t-1}$	-3.2051	-0.0925	-0.0658	0.8591 **
	(-0.4484)	(-0.9382)	(-0.6928)	(25.3511)

(b) 2nd sub-period

Independent variable	Dependent variable			
	i_t	$ln\ s_t$	$ln\ y_t$	$ln\ P_t$
Constant	10.8592 *	-0.6009	-0.3094	0.1258 **
	(1.8630)	(-1.5610)	(-0.8861)	(3.0057)
i_{t-1}	0.8921 **	0.0044	-0.0042	0.0007 *
	(20.5651)	(1.5245)	(-1.5994)	(2.2436)
$ln\ s_{t-1}$	1.5055 *	1.0149 **	-0.2435 **	0.0254 **
	(1.6565)	(16.9102)	(-4.4731)	(3.8882)
$ln\ y_{t-1}$	7.9335 **	0.2935 **	0.4759 **	0.0530 **
	(5.0255)	(2.8150)	(5.0313)	(4.6746)
$ln\ P_{t-1}$	-18.7300 **	-0.4976 *	1.2054 **	0.8612 **
	(-4.8330)	(-1.9443)	(5.1915)	(30.9606)

Notes:
*(**) Denotes significance at the 5%(1%) level.
The numbers in parentheses are the t-statistics.

TABLE 8. SUMMARY OF THE EXTENDED VAR ESTIMATES

(a) 1st sub-period

Independent variable	Dependent variable							
	i_t	$\ln s_t$	D_t	$\ln BL_t$	$\ln C_t$	$\ln M_t$	$\ln y_t$	$\ln P_t$
Constant	37.5691	0.6480	5.8037	0.2296	-1.7088	0.7587	-0.3949*	0.0952
	(0.9797)	(1.3488)	(0.4995)	(0.3941)	(-1.6023)	(0.7323)	(-0.8291)	(0.5796)
i_{t-1}	0.3737**	0.0025	0.0718	0.0030	0.0139**	-0.0173**	-0.0032*	-0.0005
	(2.5642)	(1.3428)	(1.6257)	(1.3470)	(3.4340)	(-4.3883)	(-1.7483)	(-0.8031)
$\ln s_{t-1}$	-18.1995**	0.6897**	-2.2831	-0.2080*	-0.2855	0.0816	0.0657	0.0914**
	(-2.5480)	(7.7083)	(-1.0549)	(-1.9164)	(-1.4372)	(0.4231)	(0.7406)	(2.9896)
D_{t-1}	-0.0357	-0.0000	0.7948**	-0.0108*	-0.0377**	0.0272**	-0.0142**	0.0006
	(-0.0911)	(-0.0135)	(6.6893)	(-1.8137)	(-3.4597)	(2.5685)	(-2.9244)	(0.3351)
$\ln BL_{t-1}$	-3.0167	-0.2063*	1.0320	0.5878**	-0.2843	0.4038*	0.0373	0.1117**
	(-0.3767)	(-2.0568)	(0.4253)	(4.8310)	(-1.2769)	(1.8665)	(0.3753)	(3.2587)
$\ln C_{t-1}$	6.0671*	0.0667*	-0.1135	0.0848*	0.8886**	0.1689*	-0.0920**	-0.0341*
	(1.9122)	(1.6782)	(-0.1181)	(1.7587)	(10.0712)	(1.9702)	(-2.3359)	(-2.5121)
$\ln M_{t-1}$	-1.1174	0.1811**	0.2987	0.1640**	0.3024**	0.4652**	0.0218	-0.0689**
	(-0.2769)	(3.5824)	(0.2442)	(2.6741)	(2.6947)	(4.2670)	(0.4340)	(-3.9869)
$\ln y_{t-1}$	-2.7860	0.1106	-2.9303*	0.0288	0.3770*	-0.0825	0.8873**	0.0349
	(-0.4358)	(1.3808)	(-1.5127)	(0.2965)	(2.1205)	(-0.4775)	(11.1744)	(1.2739)
$\ln P_{t-1}$	0.9625	-0.2965	2.7933	0.3936	-0.0807	-0.1759	0.3546*	0.8344**
	(0.0563)	(-1.3832)	(0.5388)	(1.5141)	(-0.1696)	(-0.3806)	(1.6686)	(11.3895)

(b) 2nd sub-period

Independent variable	Dependent variable							
	i_t	$\ln s_t$	D_t	$\ln BL_t$	$\ln C_t$	$\ln M_t$	$\ln y_t$	$\ln P_t$
Constant	62.8806**	2.0935**	-8.0539	0.6257	0.7072	-2.4134	1.7475*	0.1806
	(4.5699)	(2.3523)	(-0.4804)	(1.0304)	(1.0205)	(-1.2622)	(2.0322)	(1.5517)
i_{t-1}	0.6704**	-0.0113**	-0.0578	-0.0029	-0.0141**	-0.0059	-0.0076*	0.0007
	(10.5109)	(-2.7336)	(-0.7435)	(-1.0125)	(-4.3812)	(-0.6680)	(-1.9153)	(1.2127)
$\ln s_{t-1}$	0.6958	1.0124**	0.8488	-0.1515**	-0.2190**	0.4052**	-0.3050**	0.0223**
	(0.7225)	(16.2522)	(0.7233)	(-3.5635)	(-4.5154)	(3.0279)	(-5.0671)	(2.7418)
D_{t-1}	0.1023	-0.0014	0.6107**	-0.0012	-0.0125**	-0.0376**	0.0033	0.0008
	(1.0984)	(-0.2313)	(5.3834)	(-0.2909)	(-2.6733)	(-2.9032)	(0.5659)	(0.9534)
$\ln BL_{t-1}$	-0.3476	-0.0520	1.6641	0.9788**	0.0522*	-0.8602**	-0.1690*	0.0064
	(-0.2583)	(-0.5971)	(1.0146)	(16.4766)	(0.7700)	(-4.5988)	(-2.0093)	(0.5600)
$\ln C_{t-1}$	4.1973**	0.2583**	-0.7414	0.0733	1.0613**	0.3592**	0.2406**	-0.0010
	(3.8327)	(3.6472)	(-0.5556)	(1.5157)	(19.2424)	(2.3601)	(3.5158)	(-0.1065)
$\ln M_{t-1}$	-1.7233	-0.1812**	-0.3395	-0.0323	-0.0214	0.1726	-0.1051	0.0065
	(-1.6011)	(-2.6024)	(-0.2589)	(-0.6794)	(-0.3943)	(1.1540)	(-1.5629)	(0.7138)
$\ln y_{t-1}$	3.5582*	0.1024	3.8804*	-0.0011	-0.1231	0.3382	0.2232*	0.0482**
	(1.8286)	(0.8139)	(1.6366)	(-0.0125)	(-1.2558)	(1.2508)	(1.8357)	(2.9260)
$\ln P_{t-1}$	-23.4573**	-0.6457*	-7.8859	-0.1111	0.0053	2.4801**	1.4518**	0.8363**
	(-4.6333)	(-1.9719)	(-1.2783)	(-0.4974)	(0.0209)	(3.5253)	(4.5888)	(19.5248)

Notes:
*(**) Denotes significance at 5%(1%) level. The numbers in parentheses are the *t*-statistics.

TABLE 9.   VARIANCE DECOMPOSITION (IN PERCENTAGE) OF OUTPUT TO
MONETARY POLICY SHOCK

Quarters	Baseline VAR		
	Full sample	1st sub-period	2nd sub-period
2	6.46	27.07	1.10
4	12.84	42.04	2.14
6	17.40	48.04	2.65
8	20.13	49.86	2.61
10	21.56	49.67	2.50
12	22.17	48.60	2.50

Quarters	Extended VAR		
	Full sample	1st sub-period	2nd sub-period
2	12.90	13.62	0.73
4	19.00	18.97	1.02
6	19.90	17.67	4.70
8	19.87	14.91	7.81
10	19.82	12.08	8.31
12	19.88	9.72	7.82

interest rate is significantly and negatively related to output.

In Table 6, the F-statistics of the break point test suggest that (with the exception of ln BL_t), the single-equation parameters are significantly different from the parameters of the separate equations for each sub-period. The F-statistic for output, $ln\ y_t$ is particularly of interest. At 1% level of significance, the result suggests that the structural change in output coincides with changes in the financial system beginning in the 1990s.

We proceed to estimate the VARs for each sub-period in view of the significance of the Chow's breakpoint test. These results are reported in Tables 7 and 8. The estimated VARs are then used to compute the variance decomposition (VDC) of 12 quarters ahead to evaluate the contribution of the interest rate shock to the variance of output (see Table 9). The result shows that the contribution of interest rate shock to the variance of output fluctuations seems to have decreased dramatically in the second sub-period as compared to the first sub-period under both VAR models.

The characterization of monetary transmission mechanism appeared to have deteriorated in the post-1990:3 period when credit/financial variables are included in the VAR analysis. This is particularly true for the first sub-period. Under the baseline VAR, innovations in the policy interest rate have accounted for up to 48.60% of the output variations in the first sub-period within a three-year time horizon. The same output variation that is attributable to innovations in the interest rate fell to under 10% when we consider the extended VAR. Relative to the baseline VAR, this is indeed a significant reduction, suggesting that monetary policy shocks have accounted for very little variability in output during the first sub-period when credit/financial variables are included in the VAR analysis.

Figure 2 shows the generalized impulse responses of output to interest rate shocks. The impulse response functions in the full sample period indicate a significant reduction in output following a generalized one standard deviation innovation in the policy interest rate within the first to sixth quarter after the shock. This result is consistent in both baseline and extended VAR. However, output appears to fall rapidly within two quarters in the extended VAR

FIG 2.   RESPONSE OF OUTPUT TO A GENERALIZED ONE STANDARD DEVIATION
INNOVATION IN THE POLICY INTEREST RATE

(a)   Baseline VAR

(b)   Extended VAR
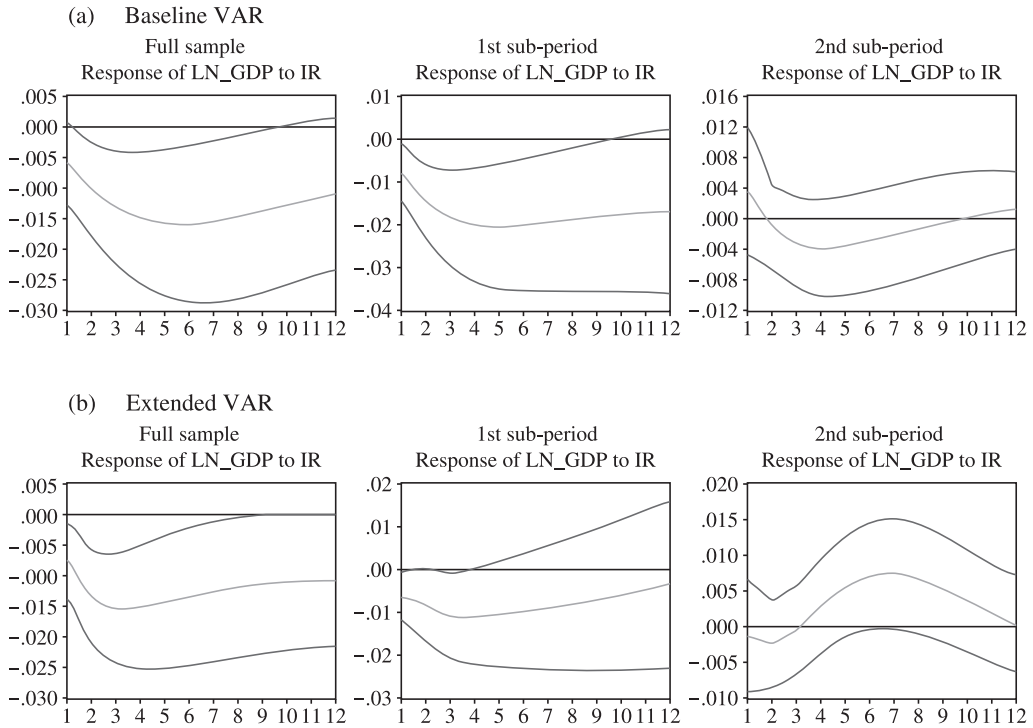
compared to the baseline VAR. The impact of the shocks generated from the extended model seems more short-lived.

When we consider the results for individual sub-periods, output tends to fall more gradually and persistently in the first sub-period under the baseline VAR. When we include the credit/financial variables in the extended VAR, the decline in output appears to reach its trough within a shorter period of time. Further, the decline in output is statistically significant for the baseline model (the two standard error band is consistently below zero), but not the extended VAR. The responses in the second sub-period are not statistically significant for both the baseline and extended models. The findings again suggest a reduction in the impact of interest rate changes on output in the post-1990:3 period or when credit/financial variables are included in the analysis.

Our findings that the effect of monetary policy shocks on output appears to have reduced in the post-1990:3 period (in both the baseline or the extended VAR) are consistent with the findings of several researchers such as Leeper, Sims and Zha (1996) and Boivin and Giannoni (2002) in the US, and Roldos (2006) in Canada. We should also take note that the standard deviation of interest rate shocks has reduced from the first sub-period to the second sub-period (see Table 10).

TABLE 10.   STANDARD DEVIATION OF INTEREST RATE SHOCKS

VAR category	Full sample	1st sub-period	2nd sub-period
Baseline VAR	1.12	1.27	0.54
Extended VAR	1.06	1.34	0.60

Notes:
The residuals from the interest rate equation in the VAR model are used for computation.

TABLE 11.   FULL SAMPLE ANALYSIS OF THE IS EQUATION

(a) OLS estimates

(I) IS equation estimates

Coefficient	Baseline model	Model with $F2$	Model with $F3$
β_0	0.9287 * (1.9025)	0.5926 (1.2035)	0.5788 (1.1621)
β_1	0.2788 * (2.2426)	0.3100 ** (2.3412)	0.3219 ** (2.4756)
β_{21}	0.7140 ** (5.7285)	0.7008 ** (5.5191)	0.6951 ** (5.5873)
β_{22}	-0.2063 ** (-3.2267)	-0.2045 ** (-3.2372)	-0.2045 ** (-3.2105)
β_{31}	-0.0830 (-0.4427)	-	-
$\beta_{31,i}$	-	0.0622 ** (0.3396)	0.1023 (0.5957)
$\beta_{41,i}$	-	-0.0433 (-1.1146)	-0.0546 (-1.2664)
β_{32}	-0.2537 * (-2.1191)	-	-
$\beta_{32,i}$	-	-0.4493 ** (-3.9024)	-0.5337 ** (-4.1452)
$\beta_{42,i}$	-	0.0674 * (1.7842)	0.0904 * (2.0376)

(II) Diagnostic tests

Test	Test statistics		
	Baseline model	Model with $F2$	Model with $F3$
Heteroscedasticity White's test (without cross terms)	1.4120 [0.1876]	1.1786 [0.3061]	1.1206 [0.3520]
ARCH	0.3241 [0.5704]	0.0117 [0.9142]	0.0903 [0.7645]
Autocorrelation Breusch-Godfrey LM test	4.6501 [0.0119]	6.7174 [0.0111]	5.1739 [0.0252]

2.   Stage Two Results

The objective of the Stage Two analysis is to derive the real interest rate elasticities for the structural IS equation for both baseline model and models incorporating measures of financial disintermediation. The elasticities provide an indication on the degree of amplification of an exogenous monetary policy shock on output. The IS equations (2) and (3) are estimated for both the full sample and individual sub-periods.[10] We subject our estimated results to several

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[10] The results are not reported, but available on request.

(b) GMM estimates

(I) IS equation estimates

Coefficient	Baseline model	Model with $F2$	Model with $F3$
β_0	0.8518 ^ (1.8337)	0.5369 (1.1853)	0.5286 (1.1705)
β_1	0.1910 (1.3598)	0.2108 (1.4592)	0.2279 (1.5661)
β_{21}	0.7491 ** (5.6430)	0.7406 ** (5.9840)	0.7330 ** (5.9847)
β_{22}	-0.1962 ** (-3.2376)	-0.1928 ** (-3.2773)	-0.1935 ** (-3.2511)
β_{31}	-0.0467 (-0.2382)	-	-
$\beta_{31,i}$	-	0.0885 (0.4853)	0.1242 (0.7050)
$\beta_{41,i}$	-	-0.0397 (-1.0429)	-0.0506 (-1.1911)
β_{32}	-0.2550 * (-1.9740)	-	-
$\beta_{32,i}$	-	-0.4311 ** (-3.5673)	-0.5097 ** (-3.7432)
$\beta_{42,i}$	-	0.0614 ^ (1.6383)	0.0829 ^ (1.8816)

Notes:

^ (*) (**) Denotes significance at 10% (5%) (1%) level.

The numbers in parentheses and brackets are the *t*-statistics and p-values respectively.

TABLE 12.   SUB-PERIOD ANALYSIS OF THE IS EQUATION

(a) OLS estimates

(I) IS equation estimates

Coefficient	Baseline model		Model with $F2$		Model with $F3$	
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
β_0	0.2770 (1.1721)	-0.8068 (-0.8309)	0.1039 (0.4271)	-0.8447 (-0.8039)	0.0870 (0.3569)	-0.7965 (-0.7712)
β_1	0.0774 (0.9576)	0.4406 * (2.4674)	0.1250 (1.5269)	0.4245 * (2.2666)	0.1100 (1.3454)	0.4241 * (2.2923)
β_{21}	1.5301 ** (16.3765)	0.3866 ** (4.1990)	1.4294 ** (15.9237)	0.3882 ** (3.9685)	1.4178 ** (15.7896)	0.3924 ** (4.0211)
β_{22}	-0.7124 ** (-8.1313)	-0.3917 ** (-5.2737)	-0.6609 ** (-7.3131)	-0.3730 ** (-4.8568)	-0.6470 ** (-7.0068)	-0.3714 ** (-4.8365)
β_{31}	0.0825 (0.6905)	0.7764 (1.4179)	-	-	-	-
$\beta_{31,i}$	-	-	0.1280 (1.1332)	2.0014 ** (2.9162)	0.0886 (0.7913)	1.9726 ** (2.7345)
$\beta_{41,i}$	-	-	0.0154 (0.2072)	-0.1141 (-1.4057)	0.0392 (0.6284)	-0.1377 * (-1.3659)
β_{32}	-0.1814 (-1.3370)	-0.4259 (-0.6495)	-	-	-	-
$\beta_{32,i}$	-	-	-0.3687 ** (-3.1963)	-1.7603 ** (-2.6174)	-0.3387 (-1.1218)	-1.7854 ** (-2.5550)
$\beta_{42,i}$	-	-	0.0772 (1.4264)	0.1267 * (1.8491)	0.0478 (1.1564)	0.1572 * (1.7827)

(II) Diagnostic tests

| Test | Test statistics | | | | | |
| | Baseline model | | Model with $F2$ | | Model with $F3$ | |
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
Heteroscedasticity White's test (without cross terms)	1.0512 [0.4276]	0.9464 (0.5199)	1.1386 [0.3737]	0.8265 (0.6377)	1.2243 [0.3166]	0.8431 (0.6210)
ARCH	0.0090 [0.9250]	0.7975 [0.3757]	0.0187 [0.8918]	0.1001 [0.7529]	0.0304 [0.8625]	0.1877 [0.6665]
Autocorrelation Breusch-Godfrey LM test	4.6432 [0.0167]	1.7128 [0.1966]	1.3146 [0.2601]	3.4360 [0.0697]	0.9003 [0.3498]	3.1496 [0.0820]

(b) GMM estimates

(I) IS equation estimates

| Coefficient | Baseline model | | Model with $F2$ | | Model with $F3$ | |
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
β_0	0.1560 (0.6382)	-0.6495 (-0.8496)	0.0525 (0.2456)	-0.6978 (-0.7958)	0.0437 (0.2002)	-0.6445 (-0.7607)
β_1	0.0123 (0.1773)	0.3558 ^ (1.7964)	0.0967 (1.1852)	0.3293 ^ (1.6560)	0.0877 (1.0471)	0.3354 ^ (1.6885)
β_{21}	1.5661 ** (16.8478)	0.4141 ** (4.0370)	1.4459 ** (15.9745)	0.4176 ** (4.2293)	1.4304 ** (15.5307)	0.4203 ** (4.2246)
β_{22}	-0.7100 ** (-7.6429)	-0.3826 ** (-5.3554)	-0.6613 ** (-7.5191)	-0.3635 ** (-4.9823)	-0.6471 ** (-7.2710)	-0.3622 ** (-4.9888)
β_{31}	0.1150 (0.8876)	0.8750 (1.6139)	-	-	-	-
$\beta_{31,i}$	-	-	0.1331 (1.2499)	2.0807 ** (3.0614)	0.0929 (0.8670)	2.0652 ** (2.8190)
$\beta_{41,i}$	-	-	0.0200 (0.3093)	-0.1110 (-1.4483)	0.0423 (0.7990)	-0.1365 (-1.3919)
β_{32}	-0.1632 (-1.2852)	-0.5845 (-0.9620)	-	-	-	-
$\beta_{32,i}$	-	-	-0.3491 ** (-3.4082)	-1.9312 ** (-2.8847)	-0.3239 ** (-3.3555)	-1.9579 ** (-2.8069)
$\beta_{42,i}$	-	-	0.0711 (1.4535)	0.1268 ^ (1.9281)	0.0441 (1.2197)	0.1585 ^ (1.8280)

Notes:

^ (*) (**)   Denotes significance at 10% (5%) (1%) level.

The numbers in parentheses and brackets are the *t*-statistics and p-values respectively.

diagnostic tests. These tests include the White's heteroscedasticity test, the autoregressive conditional heteroscedasticity (ARCH) test of one lag, and the Breusch-Godfrey serial correlation LM test of one lag. The test results suggest the presence of serial correlation in models (2) and (3).

The lag lengths are increased to two in order to eliminate the presence of serial correlation, as well as to examine the dynamics of higher-order lag lengths in our model specification. The results of the estimated Equations (4) and (5) are reported in Tables 11 (full sample analysis) and 12 (sub-period analysis). With the increase in the number of lags, the extent of serial correlation has been reduced substantially. At 1% level, there is no evidence of serial correlation, and the results are consistent across both the full sample as well as individual

TABLE 13. ESTIMATES OF REAL INTEREST RATE ELASTICITY

(a) OLS estimates

(I) Full sample

	Baseline model	Model with $F2$	Model with $F3$
$(\beta_{31} + \beta_{32})$	-0.3368 * (4.5596)	-	-
$(\beta_{31,i} + \beta_{32,i})$	-	-0.3871 * (5.2166)	-0.4315 * (6.1337)

(II) Sub-period analysis

	Baseline model		Model with F2		Model with F3	
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
$(\beta_{31} + \beta_{32})$	-0.0988 (1.5134)	0.3505 (0.9609)	-	-	-	-
$(\beta_{31,i} + \beta_{32,i})$	-	-	-0.2407 ** (9.4068)	0.2411 (0.4412)	-0.2501 ** (9.7778)	0.1873 (0.2549)

(b) GMM estimates

(I) Full sample

	Baseline model	Model with $F2$	Model with $F3$
$(\beta_{31} + \beta_{32})$	-0.3017^ (3.1767)	-	-
$(\beta_{31,i} + \beta_{32,i})$	-	-0.3426 * (4.2285)	-0.3855 * (4.9396)

(II) Sub-period analysis

	Baseline model		Model with F2		Model with F3	
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
$(\beta_{31} + \beta_{32})$	-0.0482 (0.3424)	0.2905 (0.9210)	-	-	-	-
$(\beta_{31,i} + \beta_{32,i})$	-	-	-0.2160 * (4.5930)	0.1494 (0.2414)	-0.2309 * (4.9541)	0.1072 (0.1164)

Notes:
^ (*) (**) Denotes significance at 10% (5%) (1%) level.
The numbers in parentheses are *F*-statistics in (a) and Wald-statistics in (b).

sub-periods. In addition, the results of the OLS estimates are broadly similar to those provided by the GMM estimates, hence providing support to the robustness of the findings.

In order to test the significance of the real interest rate elasticity, we apply the coefficient restriction test on both the IS equations estimated by OLS and GMM, and the results are summarized in Table 13. Both set of results are consistent in suggesting that the real interest rate elasticity remains significantly negative in the full sample period for both the baseline IS as well as the IS models incorporating measures of financial disintermediation. Individual sub-period analysis suggests that the real interest rate elasticity is not significant under the baseline IS for both sub-periods. The real interest rate elasticity remains significantly negative but only in the first sub-period for the IS models incorporating the variables *F2* and *F3*.

The significance test results for the interaction effect between interest rate and measures of financial disintermediation are given in Table 14. The findings based on the OLS and GMM estimates are similar. The coefficients capturing the interaction effect appear to be significant

Table 14. Interaction Effect of Real Interest Rate and Measures of Financial Disintermediation

(a) OLS estimates

(I) Full sample

	Model with F2	Model with F3
$(\beta_{41,i} + \beta_{42,i})$	0.0241 (1.9925)	0.0357 ^ (3.3051)

(II) Sub-period analysis

	Model with F2		Model with F3	
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
$(\beta_{41,i} + \beta_{42,i})$	0.0925 * (6.9272)	0.0125 (0.1404)	0.0870 ** (8.1453)	0.0195 (0.2238)

(b) GMM estimates

(I) Full sample

	Model with F2	Model with F3
$(\beta_{41,i} + \beta_{42,i})$	0.0216 (1.5925)	0.0323 ^ (2.6367)

(II) Sub-period analysis

	Model with F2		Model with F3	
	1st sub-period	2nd sub-period	1st sub-period	2nd sub-period
$(\beta_{41,i} + \beta_{42,i})$	0.0911 ** (7.4282)	0.0158 (0.2207)	0.0864 ** (9.3942)	0.0220 (0.2842)

Notes:

^ (*) (**)   Denotes significance at 10% (5%) (1%) level.

The numbers in parentheses are F-statistics in (a) and Wald-statistics in (b).

but we note that significance is once again, only for the first sub-period. The positive sign of the coefficients indicates that increased financial disintermediation would lead to a reduction in the dampening impact of an increase in interest rate on output.

3. Discussion of Results

Following innovations in the policy interest rate, evidence from VDC suggests that the inclusion of credit/financial variables in the VAR analysis has accounted for lesser variability in output in the second sub-period. In other words, the inclusion of credit/financial variables appears to have absorbed some of the variability of output.

The full sample analysis indicates that the real interest rate elasticity has increased significantly when we include measures of financial disintermediation. When individual sub-periods are examined, our results indicate larger and significant real interest rate elasticity for both models incorporating measures of financial disintermediation in the first sub-period. Further, the interaction effect between the real interest rate and *F2/F3* has been found to be significant. On the basis of this finding, innovations in the policy interest rate prior to the 1990:3 have accounted for larger variations in output in the presence of financial disintermedia-

tion, compared to the second sub-period.

The real interest rate elasticity is not statistically different from zero in the second sub-period under the extended IS model. The dynamics of monetary transmission mechanism appeared to have changed, resulting in lower output variability following an exogenous monetary policy shock. Our results therefore suggest that increased financial disintermediation since the early 1990s has contributed towards changes in the propagation of an exogenous monetary policy shock.

The contribution of the interest rate shock to the variability of output tends to fall when the broad credit channel of monetary transmission mechanism is included, alongside with the traditional interest rate channel beginning in the early 1990s. We could therefore make an assertion that monetary policy appears to be more effective in influencing real variables prior to the 1990:3 period in Malaysia. From the above assertion, it would appear that increased financial disintermediation in the early 1990s may have caused real interest rate to lose their significance in influencing real variables. Monetary policy in this respect is therefore less effective in influencing real variables during this period. Output tends to be stabilized because as long as firms and households are unconstrained in their access to credit, they would become less responsive to the current and/or changing economic condition. The implication is the lowering of the persistence of movement in the output gap.

IV.   *Conclusion*

The primary objective of this study is to examine the relative effectiveness of monetary policy in Malaysia in recent years with increased capital market activities following structural changes that took place in the financial system beginning in the early 1990s. Our data suggest that when we take into account the credit/financial variables into the analysis, the variability of output fell dramatically. Our data also suggest that the dynamics of monetary transmission mechanism have changed significantly from the pre-1990:3 period (first sub-period) to the post-1990:3 period (second sub-period). Specifically, output variability that is attributable to interest rate innovations in the post-1990:3 period appeared to have fallen substantially as compared to the pre-1990:3 period. Following reduced output variability in the post-1990:3 period, monetary policy appears to be less effective in influencing real variables.

There is strong evidence suggesting greater effectiveness of monetary policy in influencing real variables during the pre-1990:3 period when the process of financial disintermediation is taken into account. The interaction effect of the real interest rate and measures of financial disintermediation has been found to be significant in the first sub-period. The evidence suggests that the subdued capital market activities may have caused monetary policy to be relatively more potent during the pre-1990:3 period. There is evidence indicating that the increased financial disintermediation activities in the post-1990:3 period has been associated with reduced effectiveness of monetary policy in influencing real variables.

The empirical evidence that we have reported suggests that the lower variability of output in the post-1990 period is consistent with the results obtained by Boivin and Giannoni (2002) in the US and Roldos (2006) in Canada. A given exogenous change in the policy interest rate has led to smaller response of output. Although our results supported the findings of several researchers mentioned above, one aspect of our result is strikingly different from that of Roldos

(2006). In his study of financial disintermediation in Canada, Roldos (2006) argued that increased capital market activities have been associated with greater real interest rate elasticity, implying that the output gap in Canada is relatively more responsive to changes in real interest rate in recent years.

The findings of lower variability of output as well as the greater real interest rate elasticity in Canada imply that the systematic component of monetary policy (one that is characterized by monetary policy reaction function) has likely to have become more important with greater use of market-based financing. The situation in Malaysia appears to be different in the sense that although the effectiveness of an exogenous monetary policy shock appears to have diminished, the real interest rate elasticity has also appeared to have lost its significance in influencing the output gap in the post-1990 period. In this regard, the increased financial disintermediation since the early 1990s appears to have made the conduct of monetary policy more difficult. Our findings confirm the argument made by Chong and Goh (2005) whom suggest that in recent years, the use of interest rate alone may be limited in its effectiveness for affecting real economic activities.

Our study involves the examination of real interest rate elasticity as a function of financial disintermediation. An examination of the extent of market-based financing of the NFCS in Malaysia shows that gross nominal funds raised in the capital market accounted for, on average, only 12.16% (for $F2$) and 9.59% (for $F3$) for the period 1990:3 to 2005:4. As at the end of 2005, bank loans accounted for more than 85% of total NFCS' total financing. In industrialized economies such as Canada, this percentage is as low as 40%. It is thus not surprising that the capital market (direct financing) channel of monetary policy transmission mechanism appears to be overshadowed by the interest rate channel, resulting in the insignificance of the interaction effect of the real interest rate and measures of financial disintermediation in influencing real variables in the post-1990:3 period.

Nonetheless, we ought to acknowledge that the reduced variability in output in recent years may in actual fact be non-financial in nature (Kuttner and Mosser (2002)). Other factors such as changing behaviour of firms and consumers, re-organization of markets, as well as more effective management of inventories could have contributed to the apparent stability in output in recent years. These are legitimate arguments that warrants further consideration.

This paper focuses on the interest rate channel for monetary policy transmission mechanism. Clearly, future research should take other policy transmission channels into account to investigate the dynamics of the transmission process. The asset price channel would be particularly challenging given the changing financial landscape in Malaysia where the corporate sector is increasingly relying on market financing and the growth of mortgage loans extended to the household sector is on the increase. These developments have important implications on the asset price channel through the wealth effects on consumption and investment.[11]

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[11] We owe this point to the referee.

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