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AN APPLICATION OF AUGMENTED MONETARY CONDITIONS INDEX (MCI) IN THAILAND

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ABSTRACT

This paper estimates the augmented monetary conditions index over the quarterly period 1980:1-2004:4 using bounds testing approach for cointegration analysis as proposed by Pesaran et al., (2001). The bounds test confirms a long-run equilibrium relationship between the real output (GDP) and its determinants, i.e., the short-term interest rate (r), exchange rate (e), claims on private sector (Cops) and share prices (Sp) that account for different transmission mechanisms in the conduct of monetary policy, namely the interest rate, exchange rate, credit and asset price channels. Results reveal evidence of long-run cointegration between these variables. This has verified the stability of Thai output demand function that is needed to construct the MCI. Nevertheless, credit channel is insignificant in the model. The monetary policy stance that Bank of Thailand responses is quite corresponded with the movement in the augmented MCI during the period under study.

Keywords: *monetary policy stance, bounds test, cointegration, Thailand*

I. INTRODUCTION

During the 1980s, there was shrinking from monetary targeting and provoking an inflation targeting regime as the focal point due to the volatile financial innovations and the intensified movement of international capital transaction. Among the central banks involved are Central Bank of New Zealand, Canada, Sweden, England, Finland, Australia, Spain and Israel (see for example, Green, 1996; Svensson, 1997). Many central banks and institutions are confined to the computation of Monetary Conditions Index (hereafter MCI). The MCI model characterizes the aggregate demand (AD) (or output gap) of the economy as a function of, among other things, the real interest rates and real exchange rate. Theoretically, they are viewed as important determinants for changes in AD relative to the trend level of real output. There are a few possible monetary transmission channels by which monetary policy affects AD and inflation (see for example, Mishkin, 1995, Kuttner and Mosser, 2002, Arestis, 2002), namely the interest rate channel, the exchange rate channel, the asset prices channel, the wealth effect channel, and credit channel. In principle, as claim by Benoit (2000), all variables play a role in the monetary policy transmission process should be taken into account in the computation of the augmented MCI.

Short-term interest rates (r) have been the primary instrument of domestic monetary conditions for many countries, where changes in the stance of monetary policy affect the r and make saving-investment decisions vary. Given a certain price level, Keynes argues that an increase in money supply constitutes to a decrease in *r*, lowers the cost of capital, and simulates investment. Implicitly, marginal cost of borrowing is sensitive to spending. Changes in *r* will change the deposit rate that affects the income and cash flow of borrowers as well as lenders that ultimately influences the AD and inflation. Meanwhile, exchange rate (*e*) has direct effect on inflation and output via consumer prices of imported goods and on the prices exporters charge on the other. That is to say when foreign exchange market reacts to variations in *r*, it changes the currency value that will be transmitted through various channels and this will impact the *e* reaction that is then factored into the MCI. Meanwhile, asset prices may alter the value of collateral to facilitate lending ability. Intuitively, a fall in *r* reduces the cost of capital and increases the demand for real assets and equities (Modigliani, 1971; Gerlach and Smets, 2000; Smets, 1997; Selody and Wilkins, 2004). However, some against the incorporation of asset prices in monetary policy feedback rules (for example, Fuhrer and Moore, 1992). A credit channel of the monetary policy transmission exists when a rise in asset prices increases the borrowing capacity of individuals and firms by expanding the value of the collateral (Gauthier *et al*., 2004). Given the availability of credit, excessive growth of money supply leads to high rates of spending on domestic or foreign goods.

Since the early 1990s, the Thai economy has been claimed to be more open (Hataiseree, 1998) with respect to financial reform, which can be seen during the official acceptance of IMF's Article VIII in 1990, and the elimination of all interest rate ceilings for financial institutions in 1992, among others. Given her heavily reliance on bank lending, it constitutes a key element of transmission mechanism. However, under a liberalized financial system, private sector has greater accessed to foreign financing, which enabled them to bypass domestic financial intermediary; corollary erodes the effectiveness of monetary policy conduct through the credit channel.

After years of structural change on the exchange rate regime and the fast pace evolving financial liberalization, we now examine if there are other transmission mechanisms contribute to the calculation of the augmented MCI apart from the two benchmark indicators, i.e., *r* and *e*. The size of the elasticites to be attached to the different components of the AD, and the weights of MCI would also differ with the incorporation of other transmission mechanisms into the benchmark model. The application of an augmented MCI index in the case of Thailand is therefore an important issue to be studied to comprehensively measure domestic monetary conditions.

The present study aims to examine the possible transmission mechanisms of the conduct of monetary policy that eligible for the estimation of the augmented MCI index. The weights of MCIs are estimated according to the relationship among the variables using a single equation based econometric approach (Ozer and Mutluer, 2005). The time lag involved for each component to affect the real GDP is also examined. This paper looks at Autoregressive Distributed Lag (ARDL) bounds testing approach (Pesaran *et al*., 2001) to determine the long-run cointegration between real GDP, and its determinants over the quarterly period 1980:1-2004: 4.

The remaining of the paper is organized as follows. Section II outlines the empirical model framework, variables selection and data sources for investigation. Section III analyzes the empirical results. Section IV depicts the estimated MCIs under consideration. Finally, Section V offers conclusions and policy implications.

II. METHODOLOGY: MODEL SPECIFICATION AND DATA

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

$$rgdp = ar + be + other\ variables \tag{1}$$

where *rgdp* is the natural logarithm of the real GDP [calculated by the ratio of nominal GDP on percent of CPI (2000=100)], *r* is the short-term real interest rate [following Batini and Turnbull (2002) and OECD (1996), the *ex-ante r* is measured by the difference between call money market rate and actual inflation rate], and *e* is the natural logarithm of real exchange rate [as units of Thai currency per unit of US dollar] (where a rise in e_t represents appreciation). The parameter *a* and *b* are the coefficients terms on interest and exchange rate in the demand equation that determine the weights in the augmented MCI. The MCI ratio of *b/a* is the average coefficient estimates for the real exchange rate, reflecting the relative impact of *r* and *e* on a policy goal. By construction, *a*% point rise in *r* has the same effect on that goal of a *b*% real appreciation in the domestic currency. The higher the weight for one variable implies greater role-plays of this variable for adjusting than the other variables when shocks happen. "Other variables" include the follow: 1) Government bonds yield as proxy for long-term interest rate (*rbond*) [following Peng and Leung (2005), it is calculated by one-year rate minus CPI]; 2) real stock price index (*Sp*) to account for asset prices channel [following Mayes and Viren (2001), the real stock price is calculated by the Bangkok SET deflated by the CPI]; 3) real claims on private sectors (*Cops*) is used as a proxy for credit channel [calculated as the Cops/ CPI%]. All variables with the exception of the interest rate are expressed in logarithms (Guender, 2001; Burger and Knedlik, 2003) and all series are expressed in real terms. Bangkok SET is collected from *datastream*, while other data are obtained from the International Financial Statistics, which is published by the International Monetary Fund (IMF). Based on the outline theoretical arguments discuss above, a general function of the real GDP in the equation (1) can be written as: $rgdp_t = \mathrm{f}(r, e, Rbond, Cops, Sp)$ to account for possible channels in the transmission mechanisms. Dummy and TIME are included to capture the lag effect of Asian financial crisis, and time trend for innovations respectively. Hence, the log-linear model is specified as below:

$$Lrgdp_t = \beta_0 + \beta_1 r_t + \beta_2 Le_t + \beta_3 Rbond_t + \beta_4 LCops_t + \beta_5 LSp_t + \beta_6 TIME + \beta_7 Dummy + \varepsilon_t \tag{3}$$

Positive values are expected for β_4 and β_5, while β_1, β_2, and β_3 should be negative.

According to Eika *et al.* (1996), Steven (1998), Ericsson *et al.* (1998), Batini and Turnbull (2002), and de Wet (2002), the simple transmission process of monetary policy can be depicted in the MCI at time *t* as:

$$\mathrm{MCI}_t \equiv \alpha(r_t - r_b) + \beta(e_t - e_b) \tag{2}$$

Subscript *t* is a time index, the subscript *b* is the base period, r_t and e_t are the real interest rate and the exchange rate at current period respectively. r_b and e_b are the real domestic interest rate and the natural logarithm of the real exchange rate at base period (2000 = 100) respectively. Meanwhile, α and β are denoted respective weights, where $\alpha = a/(a+b)$, $\beta = b/(a+b)$. Intuitively, an increase in MCI reflects tightening, while a decrease in MCI signifies easing of monetary conditions.

The “bounds” testing is applied to examine the cointegration relation between output and its determinant variables. This approach is based on the estimation of an unrestricted error-correction model (UECM), with two main advantages over the conventional cointegration methods (Mah, 2000). First, ARDL estimation strategy is valid asymptotic inference that use the ordinary least square (OLS) estimates, provided the values of the maximum lag lengths are appropriately chosen to mitigate any residual serial correlation and the problems of endogenous regressors, irrespective of whether the variables are I(0) or I(1) (Hsiao, 1997; Pesaran *et al.*, 2001). Secondly, the ARDL technique has the advantage of not acquiring a precise identification of the order of the underlying data. If the “bounds” test statistics exceeds the upper critical values, there is evidence of a long-run relationship, and reject the null hypothesis of no cointegration, regardless of the order of integration of the variables. If it falls below the bound, the null hypothesis of no cointegration cannot be rejected, and if it lies between the bounds, inference is inconclusive.

To carry out the bounds test, the equation (3) is converted into UECM form as represented by equation (4) to test for the cointegration. A set of UECMs (the ARDL equation) has been estimated with four lags ($p = 4$) imposed on each first differenced term, considering the common practice of using quarterly data for the maximum order of the lags in the ARDL model (Pesaran and Pesaran, 1997: 304). According to Pattichis (1999), a parsimonious UECM is desirable to ascertain the regressor to be included in the model. The general UECM is tested downwards sequentially to arrive at a parsimonious equation using general-to-specific approach by dropping those insignificant first differenced variables sequentially.

$$\begin{aligned}\Delta Lrgdp_t &= \beta_0 + \beta_1 Lrgdp_{t-1} + \beta_2 r_{t-1} + \beta_3 Le_{t-1} + \beta_4 Rbond_{t-1} + \beta_5 LCops_{t-1} + \beta_6 LSp_{t-1} + \beta_7 TIME \\ &+ \beta_8 Dummy + \sum_{i=1}^{p} \beta_{9i} \Delta Lrgdp_{t-i} + \sum_{i=0}^{p} \beta_{10i} \Delta r_{t-i} + \sum_{i=0}^{p} \beta_{11i} \Delta Le_{t-i} + \sum_{i=0}^{p} \beta_{12i} \Delta Rbond_{t-i} + \sum_{i=0}^{p} \beta_{13i} \Delta LCops_{t-i} \\ &+ \sum_{i=0}^{p} \beta_{14i} \Delta LSp_{t-i} + \varepsilon_{1t}\end{aligned} \tag{4}$$

where β_0 is an intercept term, Δ is difference operator, ε_1 is the random error terms, and p is the lag length. The long-run relationship between the real GDP and its determinants is tested by imposing restriction on the jointly significant of estimated parameters for cointegrating test (Pesaran *et al.*, 2001), with $H_0 : \beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = \beta_6 = = H_0$: (no cointegration), and

$$H_1 : H_1 : \beta_1 \neq 0,\ \beta_2 \neq 0,\ \beta_3 \neq 0,\ \beta_4 \neq 0,\ \beta_5 \neq 0,\ \beta_6 \neq 0 \text{ (cointegration).}$$

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

III. EMPIRICAL RESULTS

The results of the exclusion test have suggested excluding *LCops* and *Rbond* from the model since they are statistically insignificant at 5 percent level. Hence, the same UECM model are re-tested and general-to-specific approach are employed, with the omission of these variables. Results in Table 1 show the estimated parameter of the final parsimonious UECM appears to trace the data well. It has the desired properties of OLS such as uncorrelated Breusch-Godfrey serial correlation LM test, constant variance or homoscedasticity of residuals–ARCH test, correct Ramsey RESET specification test, normality (Jarque-Bera test) at 5% level, as well as stable CUSUM test over the sample period.

Table 1
Real GDP Functions–UECM

Regressor	*Coefficient*	*t-statistic*
Dependent variable: $\Delta LRGDP_t$		
Constant	0.13551	1.5006
$\Delta LRGDP_{t-1}$	-0.19621	-2.1352**
$\Delta LRGDP_{t-2}$	-0.49205	-5.3263***
$\Delta LRGDP_{t-3}$	-0.24660	-2.8020***
$\Delta LRGDP_{t-4}$	0.23246	2.6083**
Δe_{t-2}	-0.08068	-2.8573***
Δe_{t-3}	-0.12672	-4.3116***
Δr_{t-1}	0.00149	1.6682*
Δr_{t-2}	0.00172	2.1430**
Δr_{t-3}	0.00156	2.0569**
$\Delta LCops_{t-4}$	0.10378	2.0813**
$LRGDP_{t-1}$	-0.05483	-2.0798**
Le_{t-1}	0.05402	3.1687***
r_{t-1}	-0.00376	-4.8874***
LSp_{t-1}	0.02584	4.6876***
$LCops_{t-1}$	-0.00805	-0.5953

Notes: Sample (adjusted): 1981Q2 to 2004Q4. R-squared: 0.79143; adjusted R bar-squared: 0.75183; Standard Error of Regression: 0.01377; residual sum of squares: 0. 01498; F-statistic (P-value): 19.9849 (0.000); Durbin-Watson statistic: 2.0650; Breusch-Godfrey lagrange multiplier test [1]: 0.42128 (0.516); LM[2]: 3.1667 (0.205); and autoregressive conditional heteroskedasticity ARCH test [1]: 0.031652 (0.859); Ramsey RESET functional test [1]: 1.3747 (0.241), Jarque Bera normality test [2]=0.031075 (0.985) . Figure in brackets (.) is p-values. Asterisk (*,** and ***) denotes statistically significant at 10%, 5% and 1% level respectively.

Table 2
Results of the Bound Tests for Cointegration Analysis

Computed F-statistics (Wald test) ($H_0 : \beta_1 = \beta_2 = ... = \beta_5 = 0$)		*9.1392****
Asymptotic critical values bounds	Lower bound, *I(0)*	Upper bound, *I(1)*
1% level	3.74	5.06
5% level	2.86	4.01
10% level	2.45	3.52

Notes: The reported critical values are from Pesaran *et al*. (2001) Table CI (iii) Case III: Unrestricted intercept and no trend (page 300) for F statistic (*k*=4).

The results of bounds test for cointegration analysis are reported in Table 2. Since the computed F-statistic, F(*Lrgdp|r, e, LCops, LSp*) = 9.139 exceeds the upper critical value *I(1)* band of 5.06 at 1 per cent level, the result of the inference shows the null hypothesis of no cointegration can be rejected, indicating *rgdp*, *r*, *e*, *LCops*, and *Lsp* are cointegrated. Next, the estimated long-and short-run elasticities of the real GDP function are presented in Table 3. The estimated long-run elasticity is found to be higher than in the short-run. This finding is consistent with economic theory and not misleading. The estimated short-run cointegrating equation is: *Lrgdp* = 0.004*r* -0.08*e* + 0.1*LCops* + 0.13; while the estimated long-run cointegrating equation is written as: *Lrgdp* = –0.06*r* + 0.98*e* - 0.14*LCops* + 0.47*LSp* + 2.47. Since there is cointegration, the long-run equation is employed for long-run policy implication purposes. It is expected that the real GDP is negatively associated with real interest rate and exchange rate appreciation, and is positively related to share price. In the long-run, all the series seem to correctly sign as expected, with the exception for *Le*. The estimated coefficient of *Le* is in positive sign which run contrary to the expected sign implying appreciation in Thailand responds in favor to trade, and induces exports at the margin in the short-run especially when income effect exceeds substitution effect.

Table 3
Estimated Short-Run and Long-Run Elasticity of Real GDP IN Thailand

Variable	*Short-run*	*Long-run*
r	0.00478**	-0.06868***
e	-0.08068***	0.98524***
LCops	0.10378**	-0.14689
LSp	-	0.47138***
C	0.13551	2.47118

Note: Asterisks (*, **, and ***) denote significant at 10%, 5% and 1% level respectively. Sample adjusted: 95 observations used for estimation from 1981Q2 to 2004Q4

Table 4
Results of Exclusion Test

Variables	*LRGDP*	*r*	*e*	*LCops*	*LSp*
X^2-statistics	4.3257	23.8871	10.0404	0.35440	21.9740
	(0.038)**	(0.000)***	(0.002)***	(0.552)	(0.000)***

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

To test for the statistical significant of the variables included in the model, Wald tests are carried out. The results of the exclusion test in Table 4 conclude that all the variables, except for *LCops*, are statistically significant at 5% level and cannot be excluded from the model.

IV. ESTIMATED MCIS

The estimated *a* and *b* from equation (1) are used to derive the MCI weight. The coefficients relating to the real short-term interest rate and the real exchange rate are -0.06 and 0.98 respectively. The associated weights of MCI index, α and β, are thus -0.065 and 1.065 respectively. The MCIs derived from the estimated AD function is

$$MCI_t = -0.065\ (r_t - r_b) + 1.065\ (e_t - e_b)$$

The weight of the interest rate over the weight of the exchange rate is -0.061. The results suggest that a one-percentage point rise in the *r* has about the same effect on GDP to 0.061-

percentage depreciation in *e*. The results give a much lower weight to the *r* than to the *e*, reflecting a one percent point change in the real interest rate has less effect on output than one percent change in the real exchange rate. As claim by Peng (2000), the estimates for small open Asian economies indicate generally smaller ratios of real interest rate against real exchange rate. Comparatively, this study shows that the absolute MCIs ratio for Thailand of 0.061 is much lower than the Thailand MCI ratio of 3.3:1 in Hataiseree (1998)'s study. This could be probably due to the coefficient is estimated with respect to the use of import price index. According to the justification argued by Grande (1997), the relative weight of the exchange rate is expected to be higher because of its direct effect on prices through imported goods.

Do the constructed MCIs match with the monetary policies implemented by the Bank of Thailand? Figure 1 plots the real GDP growth and the MCI from quarterly data. Inspect virtually, MCI tracks the inverse movements of real GDP growth reasonably well, but discrepancies notable after the onset of Asian financial crisis at the end of 1997, in which the real GDP growth appears to fluctuate. Tight stance was pursued from mid 1980-end 1981. Easing monetary policy was implemented thereafter before tight policy was practiced for a year in 1983. Easing monetary conditions was executed almost monotonically during the period of 1984 until mid-1987. Tight policy was followed thereafter, reaching its peak in mid-1990. Next, monetary conditions were loosen, reaching its trough at 1993Q3 before tightening stance was imposed thereafter for 5-quarter. Moderation of monetary conditions was pursued until the onset of Asian financial crisis in 1997Q3. Drastic changes of monetary conditions from easing conditions (1997Q2-Q3) to tighten conditions (1997Q3-Q4) were executed within three quarters consecutively. Ease monetary conditions was then followed.

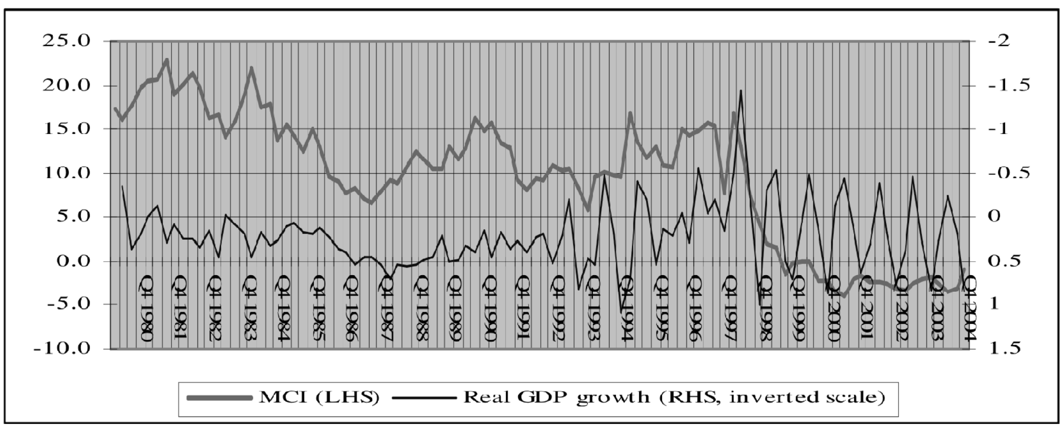

Figure 1: The Real Monetary Conditions Index and Real GDP Growth

Monetary conditions during the period under study are found to be reflected in the Bank of Thailand (BOT) reaction to the prevailing economic situation, but discrepancies are notable in a few episodes. The tight money condition was a consequence of the accelerated credit throughout 1983. However, the increasing value of the dollar against other major currencies neutralized the monetary measures to improve the economy's external position. To achieve long-term economic equilibrium, the government resorted to adjust the exchange rate system and unpegged the baht from the US dollar. This exchange rate arrangement together with speculation on exchange rate induced the capital inflows. On the whole, the government continued to pursue an accommodating monetary policy in 1987 by injecting money into the economy system. In 1991, liquidity remained generally high because of the deceleration of credit demand. Meanwhile, money market rates were generally higher than in 1989. The rise in commercial banks' deposit rates reflected increased competition among commercial banks in mobilizing deposits following the abolition of the interest rate ceiling on deposits. Since 1993, however, the Thai economy experienced inflationary pressure and persistent high current account deficit. The authorities therefore adopted a cautious monetary stance to curtail expenditure to maintain exchange rate stability. However, this policy slowed down the export and capital inflows and decelerated economic growth. In the mean time, high interest rates amidst the bad debt problem, adversely affected credit expansion. During 1994, credit to deposit ratio continued to rise, and tightened liquidity. Nevertheless, tight liquidity condition quickly eased year end following increase in bank deposits as a result of the sharp rise in net capital inflows. Beginning of 1995, Mexican financial crisis eroded international investor confidence, and caused withdrawal of funds from Thailand.

Thailand experienced currency crisis in 1997. In response to currency speculation in the early and intensified in May 1997, the authorities intervened in the foreign exchange market to stabilize the exchange rate. To defend the basket-peg exchange rate regime against the currency attacks and speculation, the authorities sold substantial amount of foreign exchange in the spot market. The efforts were successful in containing speculative pressure in offshore foreign exchange markets, however, widespread rumors continued to weaken confidence in the baht's value. Thailand exchange rate system has been changed from a basket of currencies to managed float on July 2, 1997, where the baht is determined by market forces to reflect economic fundamentals. The change to a managed float has repercussions on the economy, in which the exchange rate was extremely volatile and adversely affected the domestic price in the initial stage. Under the IMF prescription, Thai has committed pursuing tight monetary policy and financial sector restructuring. However, the contagion effect further depreciated the baht than anticipated; and led to substantial capital outflows which tightened liquidity in the money market and further aggravated the instability of financial system. Thus, the authorities had to maintain high interest rates to stabilize the exchange rate and controlling inflation as agreed under the IMF program, caused the interest rates to more than double in 1997 as compared to the year before. All these have resulted in larger-than-expected contraction in the economy and the capital market and stock market were extremely sluggish. Liquidity tightened again in September 2001 following the terrorist attacks in the US. The rate reductions constituted a signal from the BOT of the continued relaxed monetary policy in line with declining money market interest rate. During the first half of 2003 baht appreciated substantially, reaching the strongest level for the year at 39.21 baht per US dollar in October, the appreciation of the baht was driven by the weak sentiments of the US dollar emanating from US-Iraqi war uncertainties. While in 2004, monetary base expanded in line with economic growth and the baht appreciation *vis-à-vis* the US dollar were weak sentiment of the US dollar over fiscal deficit and US current account, news about devaluation of Chinese renminbi, and increase in the Thai's stock market. However, there were episodes during the baht weakened as a results of outbreak of avian flu, rising oil prices, and temporary appreciation of the US dollar.

V. CONCLUSIONS AND POLICY IMPLICATIONS

This study investigates the long-run relationship of components in the real GDP to estimate the MCIs index using the bounds testing approach proposed by Pesaran *et al*. (2001) for the quarterly period 1980:1-2004:4. It is evident that the real GDP, real exchange rate, money market rate, claims on private sector, and real share price are cointegrated in the long-run. This provides evidence that the real GDP function for Thailand is stable, and indicates stimulation of the augmented monetary conditions in Thailand is linked to the real GDP. However, credit channel is insignificant in the long-run in the construction of the augmented MCI. We find that the policy implemented by the authority corresponds reasonably well to the MCIs. A possible light of policy implication can be drawn are: First, inflation targeting is still an appropriate policy framework to be adopted by the monetary authority, as stable relationship exists among the real GDP and its determinant, which ultimately affects the inflation rate. Second, credit channel does not significantly determine the real GDP during the period study. Third, long-term interest rate is insignificant in the model. This entails that liquidity in the primary market is mature, but the liquidity conditions is less satisfactory and inactive in the secondary market. Fourth, share price appears to be significant in the model, reflecting asset price channel is matter in influencing the real GDP and ultimately the monetary conditions. In sum, the MCI weight of 0.06 implies that the conduct of monetary policy through the exchange rate channel tended to be more effective than interest rate channel in controlling inflation in Thailand during the study period. The limitations of this study exist. The MCIs are formed as a linear combination of real interest rate and real exchange rate, with weights reflecting the relative effects of the two financial variables on AD and ultimately inflation. However, the relationship may not be linear, thus cannot accurately capture nonlinearity and asymmetry, if any. It is partly for this reason that most central bank has been willing to accept quite a marked divergence between actual and desired conditions. As an approximate guideline, policy makers expect actual monetary conditions to be within a range of plus or minus 50 MCI points from desired in weeks immediately following a comprehensive inflation projection (Brash, 1997:6).

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