Changing Macroeconomic Relationships in Thailand: Effect of Financial Crisis

Somchai Amornthum*
University of Hawai‘i at Mānoa

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Abstract

This paper studies how the long-run relationships between Thai macroeconomic variables change after the 1997 financial crisis. We use Johansen's system-based cointegration method under an open-economy framework. Three cointegrating vectors are found during both pre-crisis and post-crisis periods. They represent different theoretical relations in the two periods. In the pre-crisis period, we identify the cointegrating vectors as the IS relation, the comovement of domestic and world interest rates and the Fisher relation. In the post-crisis period, one of the equilibrium relation remains to be the IS relation. The other two vectors are identified as the weak form of uncovered interest parity and the Phillips relation.

Keywords: Cointegration, Financial Crisis, Small Open Economy, Structural Change, Thailand.

JEL classification: C32, E20, E58, F41

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Contact information: amornthu@hawaii.edu.

1 Introduction

On July 2, 1997, Thailand announced the floating of its currency, baht, which marked the onset of the Asian financial crisis. Prior to the event, the baht had been pegged to a basket of currencies, mainly dominated by the U.S. dollar, since 1984. Although it was devalued from time to time, the exchange rate remained stable over a long period of time. The sudden change of exchange rate regime caused immediate and widespread effects on Thai economy. The exchange rate jumped from 25 to 50 baht per dollar within 6 months. Despite of the country's impressive growth in the first half of 1990's averaging slightly more than 8% a year, the growth rate of real GDP was –1.6% in 1997Q3 and bottomed out in a year later at –13.9%. Headline inflation increased to 10% in the middle of 1998, while core inflation peaked at 8% in the same month. Overnight interbank lending rate remained above 15% for a year following the crisis.

There have been many changes in Thai economy, especially in the monetary policy, since the crisis. Following the floating of baht, Thailand received financial assistance from the IMF. The assistance came with many guidelines, one of which was for the Bank of Thailand (BOT) to target the growth of money supply. The monetary targeting regime was adopted principally to stop the outflow of funds. Thailand switched to inflation targeting regime in May 2000. The rationale was that the money-output relation became less stable, and the monetary targeting regime was not effective. Under the new regime, the main objective is to maintain price stability and sustainable economic growth.

Given the changes, it is interesting to investigate how the economic structure has changed after the crisis. The objective of this paper is to identify the relationships between macroeconomic variables in Thailand before and after the crisis and study how they have changed. We believe that now is a good time to revisit this issue as we have learned more about the post-crisis economic structure and gained more data. Early studies often suffer from the amount of post-crisis data. They often need to combine data from pre-crisis and post-crisis periods (Fung, 2002; Disyatat and Vongsinsirikul, 2003). If the economic structure has indeed changed since the crisis, such method would yield biased estimates. Patrawimolpon (2001) fits separate models for the pre-crisis and post-crisis samples, but his post-crisis sample covers a very short period from 1997 to 2000. Reliability of his estimates is in question.

Later studies such as Atchariyachanvanich (2004) and Hesse (2005) have taken advantage of richer post-crisis data. Atchariyachanvanich employs the structural vector autoregression (SVAR) approach to analyze the monetary transmission mechanism in Thailand after the crisis. Her focus is on the post-crisis period only. As a result, her study does not give us the answer about changes occurred after the crisis. Moreover, although the SVAR provides a good description of how economic shocks transmit through the economy, it does not provide information about the long-run relationships between variables. Another approach that better fits our objective is the cointegrated vector autoregression (CVAR) approach originally proposed by Johansen (1988) and Johansen and Juselius (1990). This method can identify the long-run (and short-run) relationships between variables. Hesse (2005) adopts the CVAR approach and finds two stable relationships before the crisis—the IS relation and the inverse of money velocity—and three equilibrium relations after the crisis—the money demand, the inflation adjustment equation and the interest rate policy.

The theoretical framework employed by Hesse however is of a closed economy one. We argue that the closed-economy model is not appropriate for Thailand. Since 1990, international trade (exports plus imports) has accounted for more than 80% of Thailand's GDP, and it passed 100% in 2000. High volume of international trade clearly points to the open-economy model to describe Thai economy. We specifically focus on the *small* open-economy model. Thailand's GDP in 2004 was only \$161 billion, very small compared to \$11,712 billion of the U.S. or \$4,607 billion of Japan. The theoretical model used in this paper is similar to the Mundell-Fleming-Dornbusch model. However, we do not consider it as a strict priori. Rather, we use it as a guideline for identifying the relationships between variables.

One major drawback of the traditional CVAR approach is that it depends on asymptotic inferences. Test statistics usually have complicated finite-sample distributions, and the critical values are obtained under the asymptotic assumption. Li and Maddala (1997) prove that results based on asymptotic critical values are poor approximations of finite-sample results. Johansen (2006) suggests that either simulation or small sample corrections is needed to check the reliability of the asymptotic results. This paper conducts bootstrap simulations to estimate small-sample critical values. The bootstrap approach has a slight advantage over the small sample correction as it can be applied to all tests. Most of the results in this study are based on the bootstrap simulation.

We find three equilibrium relationships in both pre-crisis and post-crisis periods. We identify them as the IS relation, the comovement of domestic and foreign interest rates and the Fisher relation in the pre-crisis period. In the post-crisis period, one of the equilibrium relation is still the IS relation. The other two are identified as the weak form of uncovered interest parity and the Phillips relation. Clearly, the underlying relationships between macroeconomic variables have changed after the crisis. Adjustments toward equilibrium also become slower after the crisis. Further, we find that the exchange rate was weakly exogenous before the crisis, but it becomes endogenous after the crisis. Our result is consistent with the exchange rate regimes adopted at different time.

The organization of the paper is as follows. Section 2 discusses the basic theoretical framework for a small open economy. Data issues are discussed in Section 3. Section 4 outlines the multivariate estimation methodology. Empirical results are presented in Section 5. Conclusions and suggestions in Section 6 then complete the paper.

2 Theoretical Framework

This section introduces a simple small open-economy model. The model is of the Mundell-Fleming-Dornbusch type (Obstfeld and Rogoff, 1996, Ch.9). The country in focus is small in the world economy and takes the world demand for domestic output y_t^* and the world interest rate i_t^* as given. The model is static in the sense that it involves no expectation. Prices are assumed to be sticky. Money plays only the role of a unit of account and therefore does not enter the model.[1] Government does not exist in this model.

The equilibrium relation in the output market is given by the IS function

$$y_t = -\phi_{11} i_t + \phi_{12} y_t^* + \phi_{13} s_t + \phi_{14} t \tag{1}$$

where y_t is the log of domestic output, i_t is the domestic interest rate, and s_t is the log of real exchange rate. The real exchange rate s_t is defined as $e_t + p_t^* - p_t$ where e_t is the log of spot exchange rate of domestic currency to foreign currency, p_t is the log of domestic price, and p_t^* is the log of world price. An increase in the interest rate leads households to increase their savings and reduce current consumption and output. Increasing world demand boosts domestic exports and therefore the domestic output. The sign of ϕ_{13} is ambiguous since the real exchange rate affects imports and exports differently.[2] A time trend is added to capture other factors that may influence the level of output.

The Phillips relationship, which links excess demand and inflation rate, may exist when prices do not adjust instantaneously to clear markets. The level of excess demand or output gap is not observed, and several measures have been used to approximate the output gap such as detrended output, HP-filtered output or unit labor cost. This paper assumes the potential output to follow a linear trend (i.e. we assume the potential growth to be constant). This assumption is possible because our sample covers a short time span. The potential growth usually reflects the country's rates of capital accumulation and technology improvement, which are unlikely

[1]Although Dornbusch (1976) explicitly models domestic money market, the recent literature often excludes money from the utility function and the budget constraint (Gali and Monacelli, 2005). This modeling strategy assumes that money does not yield direct utility, nor is it used to transfer resources intertemporally. As a result, the money can be dropped from the model. The monetary policy is then specified in terms of an interest rate rule instead of money supply growth.

[2]It is often assumed to be positive (Obstfeld and Rogoff, 1996, Ch.9), but such assumption depends crucially on many conditions. We therefore leave the sign of ϕ_{13} to be determined empirically.

to change dramatically in a short period of time. We therefore consider the following open-economy Phillips curve (OPC).

$$\pi_t = \phi_{21}(y_t - \phi_{22}t) + \phi_{23}\Delta e_t + \phi_{24}(y_t^* - \phi_{25}t) \tag{2}$$

where π_t is the inflation. The terms $(y_t - \phi_{22}t)$ and $(y_t^* - \phi_{25}t)$ represent the domestic and world output gap. The parameters ϕ_{22} and ϕ_{25} are the domestic and world potential growth, respectively. All parameters in (2) should be positive. Excess domestic and world demand causes firms to increase their prices faster. Changes of exchange rate affect inflation through the prices of imported goods. We also consider a closed-economy version of Phillips curve (CPC), which drops the exchange rate and world output gap from (2). Bhanthumnavin (2002) has estimated the Phillips curve in Thailand during the pre-crisis and post-crisis periods. She finds that the Phillips curve exists only after the crisis based on the single-step single-equation cointegration test (Banerjee, Dolado, Hendry and Smith, 1986). It is interesting to see if the result holds in our system of equations.

The monetary authority is assumed to follow a policy rule. Our "rule" however needs not be interpreted literally in the sense that the central bank must follow it. Rather, the "rule" here can simply be a reaction function of the BOT. The most well-known policy rule is Taylor (1993) rule:

$$i_t = \phi_{31}\pi_t + \phi_{32}(y_t - \phi_{33}t). \tag{3}$$

This rule indicates that central banks should raise the interest rate in response to inflation and output gap. All parameters in (3) should be positive.

Other relationships that may exist in the open-economy framework includes the uncovered interest parity (UIP), the Fisher hypothesis and the purchasing power parity (PPP). First, if financial markets are complete, arbitrage in the markets will lead the domestic interest rate to be equal to the world rate plus the expected depreciation of the domestic currency.

This condition is known as the UIP. Using the current depreciation as a proxy of the expected depreciation, we can write the condition as

$$i_t = \phi_{41} i_t^* + \phi_{42} \Delta e_t. \tag{4}$$

The strong form of uncovered interest parity (SUIP) restricts all parameters to be unity. However, many empirical evidences does not support to the SUIP. The weak form of uncovered interest parity (WUIP) instead allows ϕ_{41} and ϕ_{42} to differ from unity. Second, the Fisher hypothesis states that the nominal interest adjusts to the expected inflation so that the *ex ante* real interest rate is not stable. Using the current inflation as a proxy for the expected inflation, we should observe that i_t and π_t moves together with one-to-one relationship (i.e. $i_t = \pi_t$). Lastly, the PPP condition holds when the price levels of two countries adjusted for the exchange rate are equal; that is, $p_t = p_t^* + e_t$. Allowing for deviations from PPP in the short run, we can examine the long-run PPP condition by testing the stationarity of the real exchange rate s_t.

3 Data

As in many studies on developing countries, we have difficulties finding high-frequency data with long time span. Although Shiller and Perron (1985)'s main suggestion is to increase the data span in order to increase the test powers, they also show that increasing frequency in a small sample can increase the power. We therefore focus on the monthly data in this study. Data on Thai economic variables are obtained from the Bank of Thailand (BOT). Overall sample covers the period from January 1990 to October 2006. The starting period is restricted by the availability of the core CPI. Our full sample is divided into two sub-samples—the pre-crisis period (January 1990 to June 1997) and the post-crisis period (July 1998 to October 2006). Table 1 summarizes all variables used in this study, and plots of these variables are shown in Figure 1 and 2.

TABLE 1: DATA DESCRIPTIONS

Variables	Descriptions	Source
y_t	Log manufacturing production index	BOT
i_t	Interbank overnight lending rate	BOT
π_t	Core inflation rate	BOT
	(Annualized month-to-month growth of core CPI)	
s_t	Log of real exchange rate	BOT, BLS
	(Log spot exchange rate + log U.S. CPI - log Thai CPI)	
De_t	Annualized growth of spot exchange rate	BOT
	(First difference of log exchange rate $\times$ 1200)	
y_t^*	U.S. real personal consumption expenditure	BEA
i_t^*	U.S. Fed funds rate	FRED

Notes: BOT = Bank of Thailand, BLS = Bureau of Labor Statistics, BEA = Bureau of Economic Analysis, and FRED = Federal Reserve Economic Data, Federal Reserve Bank of St. Louis

We use the manufacturing production index (MPI) for the output y_t.[3] We seasonally adjust the MPI using X12-Arima procedure with annual mean correction. For the interest rate i_t, we use the monthly interbank overnight lending rate, which is the average of daily rates. Although the BOT explicitly uses the 14-day repurchase rate (RP14) as its instrument, we choose the overnight rate over the RP14 mainly because of the length of data. The inflation π_t is the core inflation, calculated as the annualized month-to-month growth rate of the core CPI.[4] We choose the core inflation because (i) the BOT targets it under the current inflation-targeting regime, and (ii) the "non-core" part of headline inflation is not determined within our model but by other exogenous factors such as OPEC's decisions or weather. Next, we define the real exchange rate as $s_t \equiv e_t + p_t^* - p_t$. We use the monthly average of daily effective exchange rate of baht per U.S. dollar as e_t. The world price level p_t^* is approximated by the U.S. CPI, while Thai headline CPI is used as the domestic price level p_t. Lastly, we use the annualized month-to-month growth rate of spot exchange rate to represent the expected depreciation of

[3] Despite the fact that the manufacturing sector accounts only for one-third of the total GDP, the MPI is still the best available proxy for monthly output in terms of data availability. Other proxies such as the private consumption index have shorter time span than the MPI.

[4] Given the size of our sample, we do not want to lose one year of data by calculating year-over-year growth of CPI. We use the annualized rate so that the inflation rate is easily comparable with the interest rate.

Thai baht. To avoid confusions that may arise from the annualized rate, we switch the notation from Δe_t to De_t to denote the rate of currency depreciation.

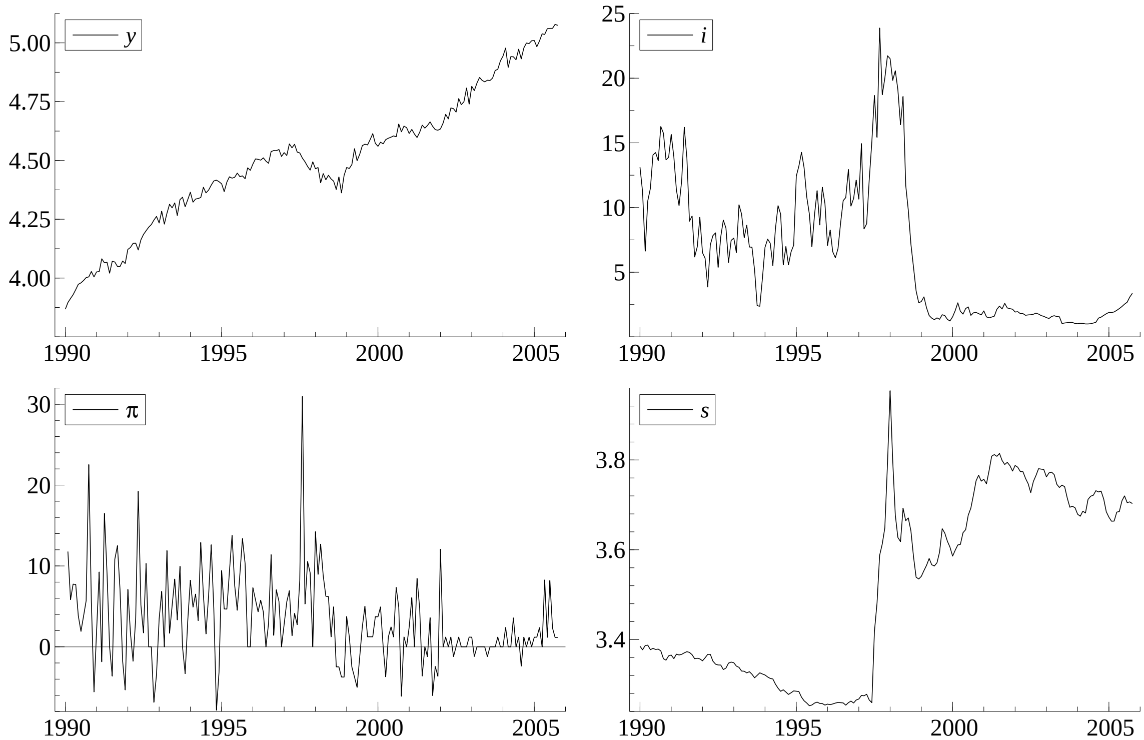

FIGURE 1: VARIABLES IN MODEL

From Figure 1 and 2, it is obvious that each Thai variable undergoes a structural change at the time of crisis. The types of changes may be different for different series. The MPI, for example, shows a drop in level during the crisis, but its growth remains at approximately 0.7% per month during both pre-crisis and post-crisis periods. The interest rate and inflation rate drop significantly and stay at low levels throughout the post-crisis period. The real exchange rate shows a slight declining trend before the crisis, but it increases substantially during the crisis. Since then, s_t have lingered around a new mean. The currency depreciation rate, though has the same mean both before and after the crisis, shows a greater fluctuation after the crisis.

In this paper, we use the log of U.S. real consumption expenditure and the fed funds rate

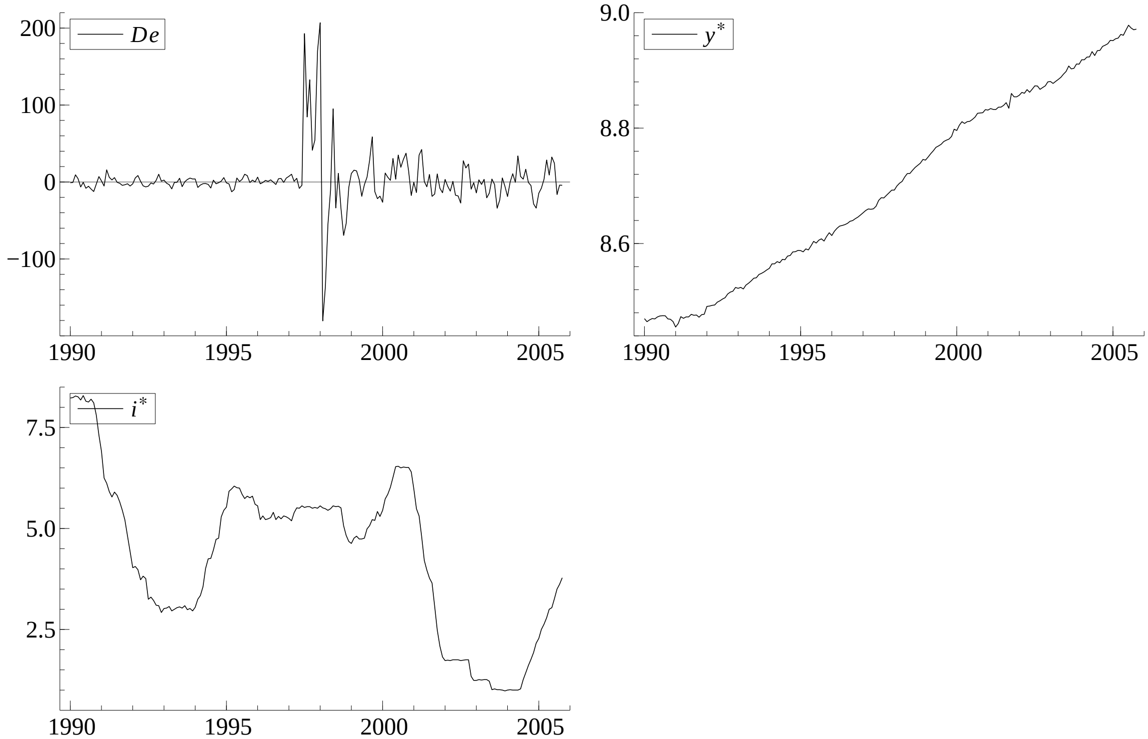

FIGURE 2: VARIABLES IN MODEL

as proxies for the world output y_t^* and the world interest rate i_t^*. Although trading with the U.S. accounts for less than 20% of Thailand's total trade, we argue that the U.S. variables are better indicators than other alternatives such as trading partners' variables for three reasons. First, many Thailand's trading partners has their output markets tied to the U.S. economy.[5] Their domestic production indicators may not truly represent their final demand of Thai goods (such as Singapore or Hong Kong). Second, it is known that financial integration in the East and Southeast Asia is lagged behind intraregional trade (Kato, 2005). Thailand's financial market is not influenced directly by the neighbors' markets but mainly by the U.S. market. Lee and Wu (2004) show that Asian interest rates, including Thailand's call money rate, converge to the U.S. rate rather than Japan's rate. Lastly, as an experiment, we used the trade-weighted pro-

[5]The top-10 trading partners of Thailand over our entire sample include Japan, U.S., Singapore, China, Malaysia, Taiwan, Hong Kong, Germany, South Korea and the United Kingdom.

duction index as a proxy of the world output in our system. We could not find the cointegrating relationship between the MPI and the trade-weighted production index as suggested by the IS relation.

In the end, our system has a total of 7 variables, 5 Thai variables and 2 U.S. variables. Our vector of variables X_t is defined as $X_t = [y_t, i_t, \pi_t, s_t, De_t, y_t^*, i_t^*]$. Unit root test on each of these variables can be found in Appendix A. The results show that only De_t in the pre-crisis periods is stationary. The evidence on π_t in the post-crisis period is mixed. All other variables are proved to be nonstationary.

4 Econometric Methodology

Although macroeconomic variables are usually found to be integrated and can wander without bound, Granger (1981) suggests that integrated variables can be "cointegrated". There may exist stationary combinations between nonstationary variables. The variables are not free to move further away from the stationary relationship. In this sense, the cointegrated relationships are often regarded as the long-run relationships between the variables. In this section, we briefly explain the Johansen's reduced rank estimation technique and the bootstrap simulation used in this paper. All estimations and simulations are performed in CATS in RATS and OxMetrics4.

We start with a brief explanation of partial system and Johansen's reduced rank technique. Consider a vector error correction model (VECM)

$$\Delta X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \Theta D_t + \epsilon_t, \tag{5}$$

where X_t is a p-dimensional vector of I(1) processes, D_t is a vector of deterministic terms, ϵ_t is i.i.d. $N_p(0, \Omega)$, and X_{-k+1}, . . ., X_0 are fixed initial values. We adopt the bottom-up strategy to choose the lag length k. That is, we begin with $k = 2$ and increase it until residuals ϵ_t

follow the Gaussian process. If there exist $p \times r$ matrices α and β such that $\Pi = \alpha\beta'$ and $0 < r < p$, then X_t are cointegrated. There are r long-run relationships between variables, which are given by $\beta' X_t$. Estimation is performed using Johansen (1988, 1991)'s reduced rank method. Restriction on Θ can be made to allow the time trend to enter the cointegrating vectors. The number of cointegrating relationship (rank of Π) is based on the trace test as suggested by Johansen (1988) and Johansen and Juselius (1990).

The full system described above involves a large number of parameters to be estimated. This usually leads to a low power of tests. Greenslade, Hall and Henry (2002) propose that by imposing weak exogeneity restrictions from the beginning we can greatly increase the power. To show the concept of weak exogeneity, we decompose X_t into a p_1-dimensional vector Y_t and a p_2-dimensional vector Z_t such that $X_t' = \left(Y_t', \; Z_t'\right)$. The vector Z_t is weakly exogenous with respect to β if (i) β is a function of parameters in ΔY_t equations alone and (ii) parameters in ΔY_t and ΔZ_t equations do not have any joint restrictions (Engle, Hendry and Richard, 1983). In practice, a necessary and sufficient condition for weak exogeneity for a variable is simply $\alpha = 0$ in the equation for that variable. Once we find which variables are weakly exogenous, the model can be reduced to a partial system (or a conditional model)

$$\Delta Y_t = \alpha\beta X_{t-1} + \omega\Delta Z_t + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1}\Delta X_{t-k+1} + \Theta D_t + \epsilon_t \tag{6}$$

with no loss of information about the cointegrating vector β. The advantage of (6) over (**??**) is clearly the lower number of parameters. This especially benefits us because our sample size is small. Asymptotic critical values of the rank test for partial systems can be found in Harbo, Johansen, Nielsen and Rahbek (1998).

A problem with the reduced rank technique is that the estimated α and β are not uniquely identified. For any $r \times r$ non-singular matrix Q, we can define matrices $\alpha^* = \alpha Q$ and $\beta^{*\prime} = Q^{-1}\beta'$ such that they yield the same matrix Π (i.e. $\Pi = \alpha\beta' = \alpha^*\beta^{*\prime}$). Results from the rank tests only tell us about the space spanned by α and β but not their economic meanings.

To identify β, we apply many H-form restrictions. We rewrite β into coefficients that need to be estimated as

$$\beta = H\varphi = (H_1\varphi_1, H_2\varphi_2, \ldots, H_r\varphi_r) \tag{7}$$

where H_i is a $p \times s_i$ matrix of restrictions, and φ_i is $s_i \times 1$ matrix of free parameters.[6] The validity of the restrictions can be tested using the likelihood ratio (LR) test, which is asymptotically distributed as χ^2 with $\sum_{i=1}^{r}(p_1 - r - s_i + 1)$ degrees of freedom. Once the cointegrating vectors are identified, we need to identify α, ω and Γ_i in (6). Economic theories usually do not give us enough information about the short-run adjustments. Therefore, we rely mainly on a statistical approach. Following the general-to-specific approach, we drop insignificant variables from the model. Our aim is to reduce the system to its most parsimonious specification. While searching for restrictions on α, ω and Γ, we take the value of β as fixed. This practice is allowed because of the super-consistency property of the estimated β.

One major drawback of the traditional cointegration analysis is that distribution of test statistics in finite-samples often depends on nuisance parameters and is difficult to derive. It becomes a tradition to compare the test statistics with asymptotic critical values instead. The asymptotic results, however, are proved to be poor approximations of finite-sample results. Li and Maddala (1997), for example, show by a Monte Carlo simulation that rank tests based on asymptotic critical values suffer serious size distortions in small samples. Two solutions have been proposed—adjusting the test statistics (e.g. Bartlett-correction trace statistics by Johansen, 2000) and estimating finite-sample critical values (e.g. bootstrap p-value by Li and Maddala, 1997). We find that the bootstrap approach has a slight advantage because it can be applied to all of our tests, not just the rank tests. Thus, we conduct bootstrap simulation for small sample inference in this paper.

We use the overlapping block bootstrap simulation with 10,000 replications for each

[6] The detail on how to estimate the model with these restriction is gruesome, and we leave the readers to consult, for example, Johansen (1995, chap.7).

test. The simulation is carried out by first estimating all parameters with no restriction. Bootstrap errors are then drawn *with replacement* from the estimated residuals based on the moving-block bootstrap approach (Künsch, 1989). That is, instead of drawing one observation at the time as in the original bootstrap approach, we resample a block of observations with fixed length of $T^{\frac{1}{3}}$. The blocks are allowed to be overlapping.[7] Then, based on initial estimators and bootstrap errors, we construct bootstrap samples under the null hypothesis. Bootstrap statistic is calculated using the bootstrap sample. The procedure is repeated 10,000 times to obtain 10,000 bootstrap statistics whose distribution gives us the finite-sample distribution of the test statistic under the null hypothesis. Once the simulation is completed, the bootstrap p-value can be easily calculated from the fraction of bootstrap samples in which the bootstrap statistics is larger than the estimated statistic.

5 Empirical Results

5.1 Model Selection and Rank Tests

Following Greenslade et al. (2002)'s suggestion, we start the analysis with the weak exogeneity test. Table 2 shows the test results. The null of weak exogeneity can be rejected for y_t, i_t and π_t during both pre-crisis and post-crisis periods. On the other hand, y_t^* and i_t^* are weakly exogenous in both subsamples. The U.S. variables are important in determining the equilibrium in Thailand, but the Thai equilibrium does not affect these variables. This supports our claim that the *small* open-economy framework should be used. We also find that s_t and De_t are weakly exogenous in the pre-crisis period but are determined within the model in the post-crisis period. The result is consistent with the exchange rate regimes adopted at different time. Before the crisis, the exchange rate was pegged at 25 baht per dollar with a very small

[7] Lahiri (1999) shows that the moving-block bootstrap is preferred over other block bootstrap methods. It gives lower variance of bootstrap variance estimators than non-overlapping block or random-length block methods. MacKinnon (2006) provides a recent literature review on the bootstrap method.

bandwidth allowed. Since the exchange rate does not move away from the pegged level, it is not affected by economic factors. Under the managed floating system adopted after the crisis, the exchange rate is allowed to move according to fundamentals. Thus, other economic factors can influence the level of exchange rate as well as s_t and De_t. Given the weak exogeneity results, we defined the vector of endogenous variables and the vector of exogenous variables as $Y_t = [y_t, i_t, \pi_t]$ and $Z_t = [s_t, De_t, y_t^*, i_t^*]$ for the pre-crisis model and $Y_t = [y_t, i_t, \pi_t, s_t, De_t]$ and $Z_t = [y_t^*, i_t^*]$ for the post-crisis model.

TABLE 2: TESTS FOR WEAK EXOGENEITY

Variable	LR Statistics	Asymptotic p-values	Bootstrap 5% c.v.	Bootstrap p-values
Pre-Crisis Model				
y_t	54.752	0.000	30.620	0.000
i_t	37.061	0.000	29.514	0.007
π_t	45.556	0.000	31.256	0.001
s_t	16.980	0.018	28.432	0.459
De_t	26.799	0.000	29.044	0.082
y_t^*	23.512	0.001	27.678	0.123
i_t^*	26.291	0.000	28.397	0.077
Post-Crisis Model				
y_t	44.210	0.000	31.292	0.003
i_t	45.019	0.000	36.333	0.010
π_t	46.459	0.000	29.027	0.000
s_t	38.441	0.000	31.409	0.009
De_t	35.636	0.000	32.549	0.026
y_t^*	22.438	0.002	32.368	0.308
i_t^*	28.907	0.000	30.102	0.065

Notes: The null hypothesis is that the variable is weakly exogenous. The likelihood (LR) statistics is asymptotically distributed as $\chi^2(7)$ with the asymptotic critical value of 14.067. The pre-crisis model is a UVAR(2) with constant and time trend; while the post-crisis model is a UVAR(3) with a constant, a time trend, three permanent dummies for Oct 2000, Jan 2002, May 2005 and a transitory dummy for Sept 2001. Bootstrap results are based on overlapping block bootstrap simulations with 10,000 replications.

The next task is to choose the lag lengths for our UVAR models. We apply the bottom-up strategy with the minimum lag length set at 2. Results from various univariate and multivariate misspecification tests are presented in Table 3. Descriptions of these tests are given in

Appendix B. In the pre-crisis period, the model with minimum lag length turns out to be sufficient to produce Gaussian residuals. The unrestricted model passes all misspecification tests. The time trend is not significant and is dropped from the pre-crisis model.[8] For the post-crisis model, we select a UVAR model with 3 lags, a time trend, 2 permanent dummies for October 2000 ($dum0010$) and January 2002 ($dum0201$), and a transitory dummy for September 2001 ($dum0110T$).[9] The three dummies are aimed to capture the effect of the establishment of Thai Asset Management Company in dealing with the non-performing loans, the introduction of Euro and the 9-11 incidence, respectively. This post-crisis model passes all misspecification tests as well.

TABLE 3: MODEL MISSPECIFICATION TESTS

Equation	Normality	AR(1)	AR(6)	Hetero	ARCH(6)
Pre-Crisis Model, 2 lags					
y_t	1.743	0.779	0.523	0.804	2.032*
i_t	2.429	1.780	0.924	0.860	0.642
π_t	0.019	0.047	0.617	0.830	0.452
Vector	3.638	0.635	1.163	0.597	NA
Post-Crisis Model, 3 lags					
y_t	4.096	0.013	1.908	0.042	1.440
i_t	0.452	0.033	0.618	0.041	0.788
π_t	4.351	3.058*	1.375	0.044	0.458
s_t	2.046	1.224	0.372	0.032	0.588
De_t	1.179	0.614	0.544	0.044	0.650
Vector	9.125	1.166	0.996	728.200	NA

Notes: * = significant at 10%, ** = 5%, and *** = 1%. Column Normality gives the test stistics from Doornik and Hansen (1994)'s normality test. Column AR(1) and AR(6) give the test statistics from the autocorrelation test at 1 and 6 lags. Column Hetero gives the test statistics on the White (1980)'s heterskedasticity test using square. Column ARCH(6) represent the test statistics from the Engle (1982)'s ARCH test at 6 lags. The NA means not enough observation to conduct the test.

[8] As an experiment, we proceed with trend in the pre-crisis model. We find the rank to be 3. The exclusion test then suggests that the trend can be excluded from the cointegrating vector anyway. The bootstrap p-values of the test is 0.422. Since it is uncommon that the trend enters the model but not in the cointegrating vector, this experiment confirms that the trend should be excluded from the pre-crisis model.

[9] Permanent dummy takes a value of 1 in a specified period, and 0 otherwise. Transitory one takes a value of 1 in a specified period, –1 in the period immediately followed, and 0 otherwise.

Table 4 reports the results from rank tests. Based on the asymptotic critical values from Harbo et al. (1998), we would have chosen rank equal to 3 for the pre-crisis model and 5 for the post-crisis model.[10] However, as shown by Li and Maddala (1997), the rank tests based on the asymptotic critical values can suffer serious size distortions in small samples. The results from our bootstrap p-values suggest instead that the rank should be 3 in both periods. The result presented here and elsewhere in this paper underline the importance of small-sample inferences based on bootstrap simulations in the cointegration analysis. The rest of the paper focuses solely on the finite-sample inference based the bootstrap simulation, while the asymptotic results are provided in tables for comparison but are not discussed.

TABLE 4: TESTS FOR COINTEGRATION RANK

Null Hypothesis	LR Statistics	Asymptotic 5% c.v.	Bootstrap 5% c.v.	Bootstrap p-values
Pre-Crisis Model				
$r = 0$	141.580	69.7	83.605	0.000
$r \leq 1$	66.458	44.5	50.676	0.002
$r \leq 2$	25.857	22.9	23.804	0.028
Post-Crisis Model				
$r = 0$	241.890	108.0	146.740	0.000
$r \leq 1$	157.370	80.9	107.010	0.000
$r \leq 2$	85.915	56.3	75.490	0.008
$r \leq 3$	41.479	35.5	45.021	0.103
$r \leq 4$	19.682	17.9	22.967	0.124

Notes: Asymptotic critical values are from Harbo et al. (1998). Bootstrap results are based on overlapping block bootstrap simulation with 10,000 replications.

5.2 Single Restrictions on Cointegrated Relations

Having determined the number of long-run relationships between variables, we next try to identify what theoretical relations these cointegrating vectors represent. To help us better

[10] In a full system, the number of rank equal the number of endogenous variables usually indicates that all variables are stationary. The same conclusion does not hold with a partial system where at least one weakly exogenous variable is nonstationary. That is because the maximum number of columns of β (the number of endogenous variables) is always less than the number of rows (the number of endogenous and exogenous variables). The matrix β can never have full rank.

understand the cointegrating relationships, we first impose exclusion restrictions on the cointegrating vectors β. We ask if any variable can be excluded from all cointegrating vectors (but *not* from the system). The test is carried out by imposing zeros on a row of β associated with the interested variable. The results presented in Table 5 suggest that De_t can be excluded from the pre-crisis model, and y_t^* may be excluded from the post-crisis model. The evidence on y_t^* however is not clear since the p value is barely larger than 5%. We do not impose these results on β immediately, but we will use this information to help us identify the cointegrating relations in the later stage.

TABLE 5: TEST OF VARIABLE EXCLUSION AND STATIONARITY

	Pre-crisis			**Post-Crisis**		
Variable	LR Statistics	Asymptotic p-values	Bootstrap p-values	LR Statistics	Asymptotic p-values	Bootstrap p-values
y_t	25.981	0.000	0.001	14.196	0.048	0.000
i_t	24.290	0.000	0.008	14.664	0.041	0.000
π_t	33.183	0.000	0.000	14.585	0.042	0.041
s_t	12.191	0.058	0.000	14.245	0.047	0.049
De_t	11.059	0.087	0.177	15.602	0.029	0.000
y_t^*	15.263	0.018	0.000	13.359	0.064	0.053
i_t^*	17.938	0.006	0.000	13.680	0.057	0.002
t				13.511	0.061	0.003

Notes: The null hypothesis is that the variable can be excluded from the cointegration vectors. Asymptotic critical values follows $\chi^2(6)$ for the pre-crisis model and $\chi^2(7)$ for the post-crisis model. Bootstrap results are based on overlapping block bootstrap simulations with 10,000 replications.

Next, we impose the theoretical restrictions as discussed in Section 2 on the cointegrating vectors. It is informative to start with one restriction at the time before we combine all restrictions. This gives us some ideas about which restriction can be a good candidate for the system restrictions. The results of single restrictions are shown in Table 6 for the pre-crisis model and Table 7 for the post-crisis model.

The IS relation as in (1) holds during the pre-crisis period, but it is no longer stable after the crisis. The bootstrap p-value is 0.036 in the post-crisis model, and y_t^* enters the relation with a wrong sign. One explanation is that the U.S. real consumption is not a good proxy for the

TABLE 6: SINGLE STATIONARITY RESTRICTIONS, PRE-CRISIS MODEL

Variable	IS	SUIP	WUIP	i-i^*	Fisher	PPP
y_t	1.00					
i_t	0.01	1.00	1.00	1.00	1.00	
π_t					-1.00	
s_t	0.31					
De_t		-1.00	-0.03			1.00
y_t^*	-2.54					
i_t^*		-1.00	-1.62	-1.00		
LR Statistics	0.08	17.77	3.71	8.19	10.90	24.71
Asym p-value	0.77		0.16	0.08	0.03	0.00
Boot p-value	0.80	0.03	0.28	0.21	0.09	0.02

Notes: Bootstrap results are based on overlapping block bootstrap simulations with 10,000 replications.

world demand of Thai goods after the crisis. Since the 1997 financial crisis, Thailand's trade with the U.S. has been declining from 18.7% of Thailand's total trade in 1998 to 11.2% in 2005. The U.S. economy also went into a recession in 2001, while Thailand was still recovering from the crisis. As also suggested by the exclusion test in Table 2, we drop y_t^* from the IS relation and rely on the time trend to represent the demand for Thai goods during this period. We argue that this is possible because Thailand's trade with China has been consistently increasing. Trading with China accounts for 3.7% of total trade in 1998, but the number increased to almost 9% in 2005. China's economic growth is very impressive itself averaging 9% a year over this period. After dropping y_t^* from the IS relation, the new restriction (IS$'$) has the bootstrap p-value of 0.052 and cannot be rejected. We conclude that the equilibrium in goods market is in fact stable in both periods, but the external factor driving Thailand's economy has changed. The U.S. is loosing its importance on Thai economy, while intraregional trade, especially trade with China, becomes increasingly important.

The fact that the time trend is excluded from the pre-crisis model limits the number of relationships, which can be tested. Any theoretical relation that contains the time trend can not be assessed in the pre-crisis period. This includes the Phillips curve and the Taylor

TABLE 7: SINGLE STATIONARITY RESTRICTIONS, POST-CRISIS MODEL

Variable	IS	IS′	SUIP	WUIP	Fisher	Taylor	OPC	CPC	PPP
y_t	1.00	1.00				22.40	-21.46	-4.79	
i_t	0.01	0.02	1.00	1.00	1.00	1.00			
π_t					-1.00	-4.56	1.00	1.00	
s_t	0.43	0.55							1.00
De_t			-1.00	-0.54			-3.19		
y_t^*	1.61						-15.62		
i_t^*			-1.00	-0.53					
t	-0.01	-0.01				-0.21	0.41	0.04	
LR Statistics	7.68	9.68	9.56	4.92	27.69	23.73	0.93	24.10	30.24
Asym p-value	0.01	0.00	0.09	0.18	0.00	0.00	0.33	0.00	0.00
Boot p-value	0.04	0.05	0.32	0.42	0.01	0.00	0.47	0.00	0.01

Notes: see Table 6 for descriptions.

rule. Both of these relations can be tested in the post-crisis sample since the time trend is included in the model. In fact, the open-economy Phillips curve (OPC) can represent one of the cointegrating relations in the post-crisis period. The bootstrap p-value is 0.469, and we cannot reject the OPC restriction. The importance of open-economy model for Thailand is again underlined here as the close-economy version of Phillips curve (CPC) clearly fails. We can reject the CPC restriction even at 1% significance level. The Taylor rule, on the other hand, is rejected in the post-crisis sample. The bootstrap p-value is less than 1%, and the output gap enters the rule with wrong sign. This result is quite surprising since Thailand has adopted the inflation targeting since May 2000. One possible explanation is that Thailand has followed two monetary regimes, the money supply targeting and the inflation targeting regimes, during our post-crisis sample. Yet, we hesitate to reduce our sample to the inflation targeting regime mainly because of the sample size. This leaves room for future research on this topic.

The strong form of uncovered interest parity (SUIP) can be rejected only in the pre-crisis model but not in the post-crisis model. The bootstrap p values are 0.026 and 0.318 respectively. Although he weak form of UIP (WUIP) holds in both periods, the coefficients of De_t and i_t^* are strikingly different. They are respectively –0.03 and –1.62 in the pre-crisis model and –0.54 and –0.53 in the post-crisis model. Many evidences seem to point to De_t for the failure of pre-crisis SUIP. First, the exclusion test from Table 5 indicates that De_t can be excluded from the cointegrating vectors in the pre-crisis sample. Second, the coefficient of De_t in the pre-crisis WUIP is very close to zero. Third, as an experiment, we proceed with WUIP, but we still cannot identify the 3 cointegrating vectors based on theory discussed in Section 2. The LR statistics in such experiment rejects the combined restrictions of IS, WUIP and Fisher relation, even though each of them is not rejected as shown in Table 6. We proceed by dropping De_t from the SUIP. The new relation is simply the comovement of domestic and foreign interest rates. An increase in foreign rate is associated with an increase in the domestic rate by the same amount. The level of the two rates does not need to be the same because factors such as country premium can cause one rate to be higher than the other. Table 6 shows that we cannot reject the comovement of interest rates during the pre-crisis period ($i_t - i_t^*$ restriction). The bootstrap p-value of the restriction is 0.21. This result is not surprising because the pegged exchange rate regime requires the comovement of interest rates. The BOT must follow any action to alleviate pressure on the baht, including adjusting the domestic rate in line with the foreign rate.[11]

Other relations that we investigate are the Fisher hypothesis and the purchasing power parity (PPP). The Fisher hypothesis holds before the crisis; that is, the real interest rate is mean reverting. In contrast, we find that the real interest rate is not stationary after the crisis. The bootstrap p-values are 0.09 and 0.01 in pre-crisis and post-crisis samples, respectively. The stability of PPP relation is tested by focusing on the stationarity of s_t. The results suggest that the PPP does not hold in both periods. The p-values are 0.02 in the pre-crisis sample and 0.01 in the post-crisis sample.

[11] It is possible that the comovement of interest rates is a result of UIP with zero expected change of exchange rate. This also explains why the SUIP restriction fails during the pre-crisis period. The current rate of depreciation (De_t) is a bad proxy for the zero expected depreciation. The zero expected change is not unrealistic given the fact that the BOT had successfully defended the value of Thai baht for a long time. More information on the exchange rate expectations is needed to decide whether the UIP holds with zero expected change of exchange rate or the UIP simply does not hold.

5.3 Final Models

In this section, we combine all possible restrictions to identify all cointegrating vectors. We describe the results for pre-crisis and post-crisis model separately.

5.3.1 The Pre-Crisis Model

For the pre-crisis model, we identify three cointegrating vectors as the IS relation, the comovement of interest rates and the Fisher relation. The restricted cointegrating relations are shown in (8) to (10) with standard deviations of the estimates in parentheses. The LR statistics for all restrictions is 17.088 with the bootstrap p-value of 0.160. All coefficients are statistically significant. Identifications of the short-run structure are given in equation (11) to (13).

$$y_t = -\underset{(0.001)}{0.0254}\ i_t + \underset{(0.128)}{1.8073}\ y_t^* - \underset{(0.175)}{1.4481}\ s_t + \varepsilon_{1t} \tag{8}$$

$$i_t = i_t^* + \varepsilon_{2t} \tag{9}$$

$$\pi_t = i_t + \varepsilon_{3t} \tag{10}$$

According to the IS relation, an increase of overnight rate by 100 basis points reduces the output by 2.5%. Elasticity of Thai output with respect to the U.S. real consumption is estimated at 1.807, while the elasticity with respect to the real exchange rate is –1.448. The estimated αs reveal a rapid restoration of equilibrium in the output market. If the output is 1% above its equilibrium level, the equilibrium relation will drive down the output in the following month by 0.919%. The interest rate also correct to the disequilibrium in output market. Excess output by 1% is associated with 18.6 basis-point increase in interest rate, which in turn reduces the output as suggested by the IS relation. Inflation does not adjust to the real disequilibrium in the long-run. The inflation also does not change with the output. This result confirms the Phillips relation does not exist before the crisis. There is no link between real shocks and movement of prices.

$$
\begin{aligned}
\widehat{\Delta y_t} \;=\;& -\underset{(0.0996)}{0.9190}\ \varepsilon_{1,t-1} + \underset{(0.0025)}{0.0226}\ \varepsilon_{2,t-1} - \underset{(0.0005)}{0.0014}\ \varepsilon_{3,t-1} - \underset{(0.0010)}{0.0021}\ \Delta i_{t-1} \\
& + \underset{(0.0004)}{0.0010}\ \Delta\pi_{t-1} + \underset{(0.5118)}{1.5027}\ \Delta y_t^* - \underset{(0.0113)}{0.0169}\ \Delta i_t^* \\
& + \underset{(0.0113)}{0.0246}\ \Delta i_{t-1}^* - \underset{(0.6058)}{5.5890} \qquad (11)
\end{aligned}
$$

$$
\begin{aligned}
\widehat{\Delta i_t} \;=\;& \underset{(10.04)}{18.6302}\ \varepsilon_{1,t-1} - \underset{(0.2657)}{0.9667}\ \varepsilon_{2,t-1} - \underset{(0.0463)}{0.0632}\ \varepsilon_{3,t-1} - \underset{(51.90)}{75.5776}\ \Delta s_{t-1} \\
& - \underset{(0.0409)}{0.0949}\ \Delta De_t - \underset{(56.44)}{162.037}\ \Delta y_{t-1}^* + \underset{(1.249)}{3.3202}\ \Delta i_t^* \\
& - \underset{(1.245)}{1.8801}\ \Delta i_{t-1}^* + \underset{(61.02)}{115.529} \qquad (12)
\end{aligned}
$$

$$
\begin{aligned}
\widehat{\Delta \pi_t} \;=\;& -\underset{(0.2520)}{1.0182}\ \varepsilon_{2,t-1} - \underset{(0.1321)}{1.0437}\ \varepsilon_{3,t-1} + \underset{(0.0835)}{0.2406}\ \Delta\pi_{t-1} - \underset{(112.9)}{772.218}\ \Delta s_t \\
& + \underset{(132.6)}{373.985}\ \Delta s_{t-1} + \underset{(0.1187)}{0.3169}\ \Delta De_t + \underset{(124.2)}{219.269}\ \Delta y_{t-1}^* - \underset{(0.9830)}{1.4238} \qquad (13)
\end{aligned}
$$

The second cointegrating relationship is identified as the comovement of interest rates. This relation is required by the pegged exchange rate system used at the time. The domestic interest must follow the foreign interest in order to reduce the depreciation/appreciation pressure on Thai baht. A change in the fed funds rate is translated to an equal change in Thailand's overnight rate. The estimated loading parameter in the interest equation is equal to –0.952, which emphasizes the importance of interest rate equalization even more. If the domestic interest rate is 100 basis points above the equilibrium level, the equilibrium relationship will force it to decline by 96.7 basis points in the next period.

The last stable relationship in the pre-crisis model is the Fisher relation. The stability of real interest rate is possible because the inflation adjusts to the change in nominal interest rate as shown in (10). Even if the inflation does not adjust to the change in interest rate in a period, it will completely correct the disequilibrium of real interest rate in the following month. In fact, the inflation usually overshoot the equilibrium level as the loading parameter is –1.04, slightly more than unity. The output also has an expected decline to any real interest rate above its equilibrium level.

5.3.2 The Post-Crisis Model

The three cointegrating relations during the post-crisis period are identified as the IS relation, the weak form of uncovered interest parity and the open-economy Phillips curve. The identified relations are presented in (14) to (16) along with the standard deviations. The likelihood ratio for all three restrictions is 26.265, and the bootstrap p-value is 0.056. Results of dynamic restrictions are given in (17) to (21).

$$y_t = -\underset{(0.004)}{0.0199}\ i_t - \underset{(0.081)}{0.6043}\ s_t + \underset{(0.000)}{0.0075}\ t + \varepsilon_{1t} \tag{14}$$

$$i_t = \underset{(0.047)}{0.5074}\ (i_t^* + De_t) + \varepsilon_{2t} \tag{15}$$

$$\pi_t = \underset{(11.061)}{55.144}\ (y_t - 0.0076t) + \underset{(49.089)}{375.314}\ (y_t^* - 0.0025t) + \varepsilon_{3t} \tag{16}$$

As tested earlier, y_t^* does not enter the IS relation. The U.S. economy becomes less important to Thailand's economy. The response of output to interest rate is about the same in the post-crisis period as in the pre-crisis period. A 100-basis-point increase in i_t now reduces the output by 1.99%, compared to 2.5% in the pre-crisis sample. On the contrary, the output is less elastic with respect to the real exchange rate. The elasticity of y_t with respect to s_t drops from –1.45 to –0.60. Adjustments toward equilibrium becomes slower in the post-crisis period. Any 1-percentage excess output will be corrected only by 0.29% in the following period.

$$
\begin{aligned}
\widehat{\Delta y_t} \;=\;& -\underset{(0.0706)}{0.2946}\;\varepsilon_{1,t-1} + \underset{(0.0007)}{0.0026}\;\varepsilon_{3,t-1} - \underset{(0.0866)}{0.4486}\;\Delta y_{t-1} - \underset{(0.0800)}{0.2574}\;\Delta y_{t-2} \\
& + \underset{(0.0051)}{0.0084}\;\Delta i_{t-2} - \underset{(0.0006)}{0.0017}\;\Delta \pi_{t-1} + \underset{(0.1310)}{0.3242}\;\Delta s_{t-2} \\
& + \underset{(0.0114)}{0.0367}\;\Delta i^*_{t-1} + \underset{(2.388)}{11.315} \qquad (17) \\
\widehat{\Delta i_t} \;=\;& -\underset{(0.0027)}{0.0040}\;\varepsilon_{2,t-1} - \underset{(0.0106)}{0.0380}\;\varepsilon_{3,t-1} + \underset{(0.0672)}{0.2360}\;\Delta i_{t-2} + \underset{(0.0103)}{0.0465}\;\Delta \pi_{t-1} \\
& + \underset{(0.0087)}{0.0337}\;\Delta \pi_{t-2} - \underset{(6.6520)}{14.363}\;\Delta y^*_{t-1} + \underset{(0.1781)}{0.6094}\;\Delta i^* - \underset{(0.1859)}{0.6917}\;\Delta i^*_{t-2} \\
& - \underset{(37.06)}{134.04} - \underset{(0.2872)}{0.4947}\;dum0201 \qquad (18) \\
\widehat{\Delta \pi_t} \;=\;& -\underset{(0.1074)}{0.2437}\;\varepsilon_{2,t-1} - \underset{(0.0761)}{0.4426}\;\varepsilon_{3,t-1} - \underset{(0.5840)}{2.0870}\;\Delta i_{t-2} - \underset{(0.0694)}{0.3063}\;\Delta \pi_{t-1} \\
& - \underset{(68.01)}{154.92}\;\Delta s_{t-1} + \underset{(0.0129)}{0.0342}\;\Delta De_{t-1} + \underset{(66.35)}{145.48}\;\Delta y^*_t - \underset{(57.84)}{107.41}\;\Delta y^*_{t-1} - \underset{(268.7)}{1561.8} \\
& + \underset{(2.291)}{15.576}\;dum0201 + \underset{(1.933)}{8.4047}\;dum0109T - \underset{(2.391)}{13.439}\;dum0010 \qquad (19) \\
\widehat{\Delta s_t} \;=\;& -\underset{(0.0443)}{0.2376}\;\varepsilon_{1,t-1} - \underset{(0.0002)}{0.0015}\;\varepsilon_{2,t-1} - \underset{(8.41e-005)}{0.0005}\;\varepsilon_{3,t-1} + \underset{(0.0477)}{0.1402}\;\Delta y_{t-1} \\
& + \underset{(0.0030)}{0.0090}\;\Delta i_{t-2} + \underset{(0.0003)}{0.0004}\;\Delta \pi_{t-2} - \underset{(0.1246)}{0.5453}\;\Delta s_{t-1} - \underset{(0.3224)}{1.1157}\;\Delta y^*_{t-1} \\
& - \underset{(0.3215)}{0.5865}\;\Delta y^*_{t-2} + \underset{(0.0139)}{0.0383}\;dum0010 \qquad (20) \\
\widehat{\Delta De_t} \;=\;& -\underset{(54.33)}{286.18}\;\varepsilon_{1,t-1} - \underset{(0.1919)}{1.1053}\;\varepsilon_{3,t-1} + \underset{(58.60)}{158.41}\;\Delta y_{t-1} + \underset{(3.756)}{9.1779}\;\Delta i_{t-2} \\
& + \underset{(0.1527)}{0.3268}\;\Delta \pi_{t-1} + \underset{(0.4107)}{0.6958}\;\Delta \pi_{t-2} - \underset{(99.80)}{730.35}\;\Delta s_{t-1} + \underset{(102.0)}{154.65}\;\Delta y^*_t \\
& - \underset{(401.8)}{1544.8}\;\Delta y^*_{t-1} - \underset{(396.7)}{881.44}\;\Delta y^*_{t-2} - \underset{(2.405)}{5.8921}\;\Delta i^*_{t-2} - \underset{(573.3)}{1978.4} \\
& + \underset{(17.13)}{45.991}\;dum0010 - \underset{(4.397)}{17.274}\;dum0201 \qquad (21)
\end{aligned}
$$

Because the exchange rate is no longer pegged, the expected depreciation of Thai baht becomes an important determinant in the relationship between the domestic and foreign interest rates after the crisis. As shown in Table 7, the weak form of UIP is stable, and the coefficients on De_t and i_t^* are very close in magnitude. Thus, as shown in (15) we restrict the two variables to enter the relation with equal coefficients. The result shows that an increase of i_t^* by 100 basis points or a depreciation of Thai baht by 1% would increase the domestic interest rate by 51 basis points. Both domestic interest rate and exchange rate adjust to bring back the equilibrium. However, the adjustment toward equilibrium tend to be slow. If i_t is 1% above the equilibrium level, the equilibrium relation will force it to decline only by 0.004% in the next period.

The inflation after the crisis responds to the domestic and foreign excess output as indicated by the open-economy Phillips relation. This result is consistent with Bhanthumnavin (2002) who also finds that the Phillips relation exists only after the crisis. In this relation, we restrict the potential growth to be the mean of output growth during this sample, which is 0.76% a month for Thai output and 0.25% a month for U.S. real consumption.[12] Our estimated Phillips curve shows that the domestic output gap of 1% pushes the inflation up by 55 basis points. A foreign output gap of 1 % increases π_t by 375 basis points. Although the response to foreign output gap may seem to be too strong, it should be noted that the average size of foreign output gap is much smaller than the domestic output gap. For 95% of the time, the growth of y_t^* fluctuates between -0.12% and 0.68% per month, while the growth of y_t varies between -2.24% and 3.76%. Adjustment of inflation is fairly slow after the crisis. Before the crisis, the inflation adjusts fully correct in the following month to the disequilibrium in real interest rate. However, after the crisis, the inflation reduces only by 0.24 basis point if the inflation is 1 basis point above the equilibrium suggested by the Phillips curve.

[12] In an unrestricted version, we make no restriction on the coefficient of time trend and find the estimated coefficient to be similar to our restriction.

6 Concluding Remarks

This paper identifies the equilibrium relationships between Thai macroeconomic variables using the system-based cointegration method. We focus specifically on the changes of these relationships between pre-crisis and post-crisis periods. In contrast to the existing literature, we explicitly employed the small open-economy theoretical framework, which is more appropriate to the Thai economy than close-economy models. We also follow the modeling strategy proposed by Greenslade et al. (2002) by imposing weak exogeneity from the outset to increase the power of the tests. Further, it is widely known that asymptotic tests in the cointegration analysis usually face severe size distortions. To alleviate the distortions and to improve the power, we perform bootstrap simulations to estimate the finite-sample critical values in most tests.

Our results from the weak exogeneity tests confirm that the small open-economy theory should be used to identify the equilibrium relations in Thai economy. Foreign output and interest rate are proved to be weakly exogenous. They enter the cointegrating vectors but are not affected by disequilibrium in Thai economy. The exchange rate is found to be weakly exogenous before the crisis because of the pegged exchange rate system adopted at the time. It becomes determined within the model once the BOT switched to the managed float system.

According to the rank tests with bootstrap p-values, we find 3 cointegrating vectors during both pre-crisis and post-crisis periods. We prove that they represent different theoretical relations during the two periods. The three vectors are identified as as the IS relation, the comovement of interest rates and the Fisher relation before the crisis. In the post-crisis period, one of the equilibrium relation remains to be the equilibrium in the goods market. The other two vectors are identified instead as the weak form of uncovered interest parity and the Phillips curve. The domestic and foreign interest rates no longer move in one-to-one relation after the crisis. Also, the real interest rate does not show the mean reverting pattern as in the pre-crisis period. Based on this evidence, we conclude that the relationships between macroeconomic variables in Thailand have changed after the 1997 financial crisis.

Our conclusion has an important implication on future research related to Thai economy. It is not appropriate to combine data before and after the crisis since the underlying economic structure has changed after the financial crisis. Instead, future studies should try to estimate models using the pre-crisis and post-crisis data separately. We note that the literature on estimating models with structural change has grown rapidly in recent years (see for example Perron, 1989). These methods however assume that changes occur only in deterministic components. In the case of Thailand, not only do we observe a shift in level, but we also witness the change in relationships among macroeconomic variables. The interest rate for example move one-to-one with the fed funds rate before the crisis but less than one-to-one after the crisis. Our model is also small and easy to manage. This can be a supplement to the large-scale macro model currently used by the BOT.

Appendix

A Unit Root Tests

The concept of cointegration becomes trivial in a system where all variables are stationary. Our task in this section is to apply unit root tests on each variable. We focus on the M_{α}^{GLS} and M_{t}^{GLS} tests proposed by Ng and Perron (2001) and the *DF-GLS* test proposed by Elliott et al. (1996). The lag length k for each unit root test is chosen based on the modified Akaike Information Criterion (MAIC) also proposed by Ng and Perron (2001).[13] We set the maximum lag length at 13 in all cases. The tests are conducted for the pre-crisis and post-crisis samples separately. We do not run the tests using the full sample since the results can be misleading. Perron (1989) and Rappoport and Reichlin (1989) argue that the standard unit root tests may falsely "accept" the nonstationary hypothesis when the true data generating process contains a structural break. We note the development in the unit root tests that allows for a structural break in deterministic terms (see for example Zivot and Andrews, 1992; Perron and Rodriguez, 2003). However, Figure 1 and 2 suggests that most variables have different variances, not only different levels or slopes, after the crisis. Since the entire DGPs seem to change, it is more appropriate to conduct the unit root tests on each period separately.

Results of the unit root tests on the variables used in this paper are shown in Table 8. In sum, most variables are nonstationary. The exceptions are Δe_t during the pre-crisis period and the π_t during the post-crisis period. The evidence is clear that Δe_t is stationary in the pre-crisis period as all tests reject the null at 1% significance level. However, it is less clear in the case of π_t in the post-crisis period. Only the M_{α}^{GLS} and M_{t}^{GLS} tests reject the nonstationarity of π_t in the post-crisis period, while the *DF-GLS* test can not reject the null even at 10% level.

We present the evidence in this section that most of the variables are nonstationary.

[13] They point out that the conventional criteria such as AIC or BIC tend to choose the lag length too low because those criteria assign penalties to overfitting the model but not underfitting. The MAIC considers different penalties for integrated data.

TABLE 8: UNIT ROOT TESTS, PRE-CRISIS AND POST-CRISIS

	Pre-Crisis			Post-Crisis		
Variable	M_{α}^{GLS}	M_{t}^{GLS}	*DF-GLS*	M_{α}^{GLS}	M_{t}^{GLS}	*DF-GLS*
y_t	-3.86	-1.13	-1.54	-7.97	-1.98	-1.59
i_t	-4.65	-1.48	-1.84*	0.01	0.01	-0.30
π_t	0.21	0.33	-0.11	-36.20***	-4.25***	-1.03
s_t	0.70	0.64	0.60	-4.68	-1.46	-0.76
Δe_t	-142.20***	-8.43***	-2.72***	-2.37	-1.04	-0.61
y_t^*	-2.10	-0.93	-0.97	-3.38	-1.26	-1.02
i_t^*	-2.90	-1.17	-0.64	-3.12	-1.25	-1.25

Notes: * = significant at 10%, ** = 5%, and *** = 1%. A constant and a t are included when testing y and y^*; otherwise, only a constant is included. The pre-crisis covers the period from Feb. 1990 to Jun. 1997, while the post-crisis covers the period from Jul. 1998 to Oct. 2006.

Even if some variables are stationary, the vector process $X_t = [y_t, i_t, \pi_t, s_t, \Delta e_t, y_t^*, i_t^*]'$ is still nonstationary. Further, since there are more than two nonstationary variables in each period, it is possible to have at least one cointegrating vector. This provides a justification of cointegration analysis.

B Diagnostic Tests

In this appendix, we describe the tests used to examine the properties of residuals from UVAR model. The tests include normality test, Lagrange multiplier (LM) test for autocorrelation at 1 and 6 lags, heteroskedasticity test using squares, and autoregressive conditional heteroskedasticity (ARCH) test. Let $u = (u_1, \ldots, u_T)$ be the residuals from (6).

B.1 Normality Test

The normality test used in this paper is proposed by Doornik and Hansen (1994). The test statistic is

$$e_2 = z_{12} + z_{22}$$

where z_{12} and z_{22} denote the transformed skewness and kurtosis of u_t. The null hypothesis is that $\widehat{u}$ is normally distributed. The test statistic is approximately distributed as $\chi^2(2)$.

Let $(x_1, \ldots, x_n)$ be a sample of n independent observations on a 1-dimensional random variable. Sample skewness (b_1) and kurtosis (b_2) are defined by:

$$\bar{x} = \frac{1}{n}\sum_{i=1}^{n} x_i, m_i = \frac{1}{n}\sum_{i=1}^{n}(x_i - \bar{x})^i, \sqrt{b}_1 = \frac{m_3}{m_2^{3/1}}, \text{ and } b_2 = \frac{m_4}{m_2^2}.$$

The transformation of b_1 into z_{12} is

$$\begin{aligned}
\beta &= \frac{3(n^2+27n-70)(n+1)(n+3)}{(n-2)(n+5)(n+7)(n+9)}, \\
\omega^2 &= -1 + 2(\beta-1)^{\frac{1}{2}}, \\
\delta &= \frac{1}{\{\log(\sqrt{\omega^2})\}^{\frac{1}{1}}}, \\
y &= \sqrt{b_1}\left\{\frac{\omega^2-1}{2}\frac{(n+1)(n+3)}{6(n-2)}\right\}^{\frac{1}{2}} \\
z_1 &= \delta \log\left\{y + (y^2+1)^{\frac{1}{2}}\right\}.
\end{aligned}$$

The kurtosis b_2 is transformed from a gamma distribution to χ^2, which is then translated into standard normal z_2:

$$\begin{aligned}
\delta &= (n-3)(n+1)(n^2+15n-4), \\
a &= \frac{(n-2)(n+5)(n+7)(n^2+27n-70)}{6\delta}, \\
c &= \frac{(n-7)(n+5)(n+7)(n^2+2n-5)}{6\delta}, \\
k &= \frac{(n+5)(n+7)(n^3+37n^2+11n-313)}{12\delta}, \\
\alpha &= a + b_1 c, \\
\chi &= (b_2 - 1 - b_1)2k, \\
z_2 &= \left\{\left(\frac{\chi}{2\alpha}\right)^{\frac{1}{3}} - 1 + \frac{1}{9\alpha}\right\}(9\alpha)^{\frac{1}{2}}.
\end{aligned}$$

B.2 LM Test for Autocorrelation

The LM tests for autocorrelation can be done by estimating

$$u_t = \sum_{i=1}^{k} \varrho_i u_{t-i} + \varepsilon_t$$

where we choose k = 1 and 6 in this study. The null hypothesis is that there is no autocorrelation ($H_0 : \varrho_i = 0$ for all $1 = 1, \ldots, k$). The LM test is TR^2 where R^2 is coefficient of determination from the regression above. The test is distributed as $\chi^2(k)$.

B.3 Heteroskedasticity Test Using Squares

This test is based on White (1980). It involves the following regression:

$$u_t^2 = c_0 + \sum_{j=1}^{p} \kappa_{1j} x_{jt} + \sum_{j=1}^{p} \kappa_{2j} x_{jt}^2$$

where x_{jt} are the elements of X_t in (**??**) and (6). The null hypothesis is that the variance of u_t does not depend on the regressors and their squares ($H_0 : \kappa_{ij} = 0$ for all $i = 1, 2$ and $j = 1, \ldots, p$). The usual F-test can be used in this case.

B.4 Autoregressive Conditional Heteroskedasticity Test

The ARCH (AutoRegressive Conditional Heteroscedasticity) test is proposed by Engle (1982). The tests considers the variance of the current error term to be a function of the variances of the previous time period's error terms as

$$u_t^2 = c_0 + \sum_{i=1}^{k} \gamma_i u_{t-i}^2.$$

The null hypothesis is that the variance should be constant ($H_0 : \gamma_i = 0$ for all $i = 1, \ldots, k$). We set k equal to 6 in this study. Again, the F-test can be used to test the hypothesis.

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