

Factors Accounting for the Thai Intermediation Premium and Conditional Heterocedasticity

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Abstract

Asymmetries in the Thai lending-deposit rate spread (intermediation premium) are documented. Empirical results revealed that the intermediation premium adjusts to the threshold faster when the deposit rates increase relative to the lending rates rather than when the deposit rates move in the opposite direction. Additionally, empirical findings indicate that Thai commercial banks exhibit predatory rate setting behaviour. These findings also show bidirectional Granger causality between the Thai lending rate and the deposit rate, indicating that the lending rate and the deposit rate affect each other's movements. These results suggest that monetary authority can use its countercyclical monetary policy instruments to achieve macroeconomic objectives in the short run. However, the estimation results of the GARCH (2,2)-in-Mean model indicate that more frequent intervention by small policy measures is necessary to minimize the conditional variance of the intermediation premium in order to minimize the magnitude of the cycle of the lending rate.

Keywords: Asymmetry; lending rate; deposit rate; intermediation premium; monetary policy.

JEL classification codes: C22; E44; G21.

1. Introduction

Financial intermediation drives profitability or financial distress; and, is a critical facilitator of investment and economic growth (Schumpeter 1912; Patrick 1966; McKinnon 1973). Commercial banks play a crucial role in determining the spread between the lending rate charged to borrowers and accounting for the cost of funds. In addition to creating interest income to financial intermediaries, this spread affects the economy's savings and investment level, the effectiveness of a central bank's monetary policies, as well as economic development and social progress. Some of the spread is risk related to the instrument; that is, the intermediation premium over and above the "accounting cost of funds" level. This "risk" portion provides useful insights into banking behaviour. Accordingly, this paper explores the behaviour of Thai banks in particular—with an emphasis on the factors that affect the spread between Thai lending rates and deposit rates (or henceforth the "intermediation premium"), and the dynamic relationships amongst these factors.

Additionally, the Thai banking sector has become more and more internationalized just as the international economic landscape, over the last three decades or so, has been dotted with international political and social turmoil. These developments exacerbate the variance of the intermediation premium and cause the variance to be different from some sub-

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periods to others over the sample period. Therefore, another important question is whether the fluctuations in the variance of the intermediation premium from one month affect the premia and the variances in subsequent months. Countercyclical monetary policy makers rely heavily upon this very important information and thus to the question: should these policy makers less frequently and by large policy measures intervene to bring the economy to its long-term trend; or, should they intervene more frequently and by small policy measures? As we show in this study, these two alternative policy provisions result in a dissimilar variance of the intermediation premium. An increase in variance is indicative of increased risk associated with the debt instrument, which, in turn, exacerbates the lending rate in the banking sector. This investigation specifies and estimates a simple GARCH (s, r)-in-Mean (GARCH-M) model to discern this possibility.

In theory, banks operating in a free market economy should be expected to consider all sources of risk in determining and setting the spread that separates the rate paid to lenders from the rate charged to borrowers. If banks set an intermediation premium either too high or too low, market forces would nominally force an adjustment back to some equilibrium spread. Three main hypotheses explain this rate-setting behaviour: the bank concentration hypothesis, the consumer characteristic hypothesis, and the consumer reaction hypothesis. The bank concentration hypothesis theorizes that oligopolistic banks are slow to raise deposit rates, but quickly raise lending rates when market forces allow it. Conversely, banks in declining markets quickly adjust downward the rates paid to depositors and only slowly reduce the rates charged to borrowers (Neumark and Sharpe, 1992, Hannan and Berger, 1991). The consumer characteristic hypothesis posits that banks can adjust rates to widen the spread and increase their profitability to the extent that consumers are unsophisticated and/or are saddled with higher costs of searching and switching (Calem and Mester, 1995, Hutchison, 1995, Rosen, 2002).

The consumer reaction hypothesis proposes that asymmetric adjustments in lending rates may actually benefit consumers, because the presence of asymmetric information can foster an adverse selection problem in lending markets such that higher interest rates will tend to attract riskier borrowers (Stiglitz and Weiss, 1981). Therefore, even if the market rates rise, banks would be reluctant to raise lending rates, because the expected cost to the banks of not raising the lending rates (when their marginal cost of funds increases) is offset by the risk reduction benefits of not encouraging the higher-risk borrowers.

As is discussed in the next section, in the last three decades which were dotted with considerable international, political, and social turmoil, the Thai banking sector has gone through a drastic reform process. Interest rates were allowed to adjust to the rate of credit expansion and were very much affected by international rates as a result of the Thai open economy. Domestic prices also were largely determined by world price movements as a result of the country's open economy and minimal domestic price controls. Accountants across the world stress the need to for understanding economic modeling as well as the link between fair value accounting (on securities) to better inform policy decisions.

Therefore, it is now of special interest to assess the Thai commercial banks' rate setting behavior after almost two decades of reforms of the banking sector and understand the contributing factors, including accounting's effect on economic modeling. It is also of interest to compare the Thai banks' rate setting behavior to those of their counterparts in advanced market economies to gauge the effectiveness of the aforementioned liberalization and reforms. To this end, this paper explores whether asymmetries exist in the Thai lending-deposit rate spread and, if such asymmetries are present, how lending and deposit rates respond to these asymmetries. Furthermore, this paper explores whether responses to such asymmetries are independent or are dynamically interrelated. Also, this analysis seeks to determine whether the Thai lending institutions exhibit competitive or predatory pricing

behaviours, and to what extent. Finally, this study investigates whether the variance of the intermediation premium from one month affects the variances and premia in subsequent months.

The remainder of this study is organized as follows: section 2 summarizes the literature on asymmetric rate adjustments by international lending institutions and Thai banking sector; section 3 describes the data and the descriptive statistics used in the analysis; section 4 describes the methodology used in the investigation; section 5 reports the empirical results; and section 6 provides a summary of the study's takeaways and offers concluding remarks and policy implications.

2. Asymmetric Rate Adjustments and the Thai Banking Sector

2.1 Asymmetric Rate Adjustments – Literature Summary

The rationale for theoretically hypothesizing asymmetric responses to the national countercyclical monetary policy can be attributed to the documented asymmetric rate-setting behavior of commercial banks in the context of rates of return on financial market instruments. Dueker (2000) and Tkacz (2001) have reported asymmetries in the U.S. prime lending rate in the past. Thompson (2006) found asymmetries in the U.S. prime lending-deposit rate spread. Sarno and Thornton (2003) found asymmetries in U.S. Treasury securities in their studies. Frost and Bowden (1999) and Scholnick (1999) reported asymmetries in mortgage rates in New Zealand and Canada. Hofmann and Mizen (2004) indicated asymmetric behavior of retail rates in the United Kingdom. Hannan and Berger (1991), and Neumark and Sharpe (1992) examined various deposit rates for the same behavior. Several studies have found asymmetric cointegration between bank lending and deposit rates. For instance, Nguyen et al. (2008) documented similar asymmetries in Mexican lending and deposit rates. Nguyen and Islam (2010) reported asymmetries in the Thai bank lending and deposit rates. Nguyen and Henney (2013) found asymmetries in the US housing mortgage market. Chang and Su (2010) reported nonlinear cointegration between lending and deposit rates in ten Eastern European countries. Also, Haug and Basher (2011) found nonlinear cointegration in the purchasing power parity relationships for Canada, Japan, Switzerland, the U.K., Belgium, France, Germany, Italy and the Netherlands.

2.2 The Thai Banking Sector

Thailand: A Country Study (1987) articulated that monetary policy was traditionally passive. Control over the rate of credit extension was the primary means for supporting growth, maintaining price stability, and monitoring the balance of payments. Interest rates were allowed to adjust to the rate of credit expansion and were very much affected by international rates as a result of the Thai open economy. Domestic prices were also largely determined by world price movements as a result of the country's open economy and minimal domestic price controls.

As reported in the July 2010 Asia Focus, regulators in Thailand began to consider the need for systemic reform after recognizing that weaknesses inherent in the financial system had intensified the 1997-98 Asian Financial Crisis, spawning a banking crisis in addition to the currency crisis. In 2004 Thailand began to implement the Financial Sector Master Plan (FSMP), a very important long-term reform program aimed at creating a more efficient, transparent, and internationally competitive financial sector that can serve a larger proportion of the Thai population. With the first phase of the FSMP completed in 2009, Thailand announced that a second phase of financial sector reforms would begin in 2010.

Bank of Thailand, the central bank of Thailand, (2015) reported that for the past 10 years, the banking sector has remained the main channel of financial intermediation, especially for the private sector, by competing for most deposits and loan business.

Specially, specialized financial institutions (SFIs) have played a greater role in financial intermediation in Thailand. Within the Thai banking sector, SFIs have gained more importance in credit extension, initially as a main part of the government's countercyclical policies during the global financial crisis (GFC), and to support the government's spending and stimulus measures in the subsequent periods.

At the end of 2013, total credits extended by all deposit-taking institutions (excluding the central bank) accounted for 47% of total sources of financing. In terms of participants, the four largest domestic commercial banks have maintained their dominance as financial intermediaries for households and firms, and thus have maintained their strong influence on the pass-through of monetary policy to the economy. Unlike the 1997 Asian economic crisis, the GFC did not directly bring significant changes to the Thai financial landscape due to the country's sound economic fundamentals and banking system. The largest four domestic commercial banks, comprising Bangkok Bank, Kasikorn Bank, Siam Commercial Bank and Krungthai Bank, made up 61% of commercial banks' total assets at the end of 2013.

Also, over the past decade, the Thai banking sector has maintained its predominance in intermediating financial resources to the real economy. Government policies have played a part in influencing the degree of competition within the sector. In addition, changes in economic and financial conditions at home and abroad have altered banks' balance sheets and business operations.

Foreign banks are defined as commercial banks with foreign control of 50% or more of total ownership. They are categorized as either branches or subsidiaries. Foreign banks remain of limited importance, accounting for just 14% of the banking sector's total assets. In contrast to SFIs and domestic banks, foreign banks participate in the major areas of Thai bank business, retail lending and deposits, to only to a small extent. With only a small number of branches to facilitate retail lending, foreign banks instead tend to focus on areas where they can bring more expertise and comparative advantage to bear, such as wholesale funding, investment banking and trading services, as well as credit extension to affiliated foreign firms. Despite the second phase of the Financial Sector Development Plan (2010–14), which aimed to enhance the level of competition within the banking sector partly via promoting greater foreign participation, competition from foreign banks remains small-scale, with only a marginal increase in the number of foreign bank branches.

As to the current state of the Thai banking sector, IMF Country Report no. 15/144 (2015) argued that while the current monetary stance is accommodative, there is scope for further easing if the recovery is weaker than anticipated. The Bank of Thailand cut the policy rate by 25 basis points in November 2013 and again in March 2014 to support the economy during a period of heightened political unrest, and has maintained the rate at 2 percent ever since. The Thai baht depreciated more than 8% during 2015 (CIA World Factbook March 14, 2016.)

3. The Data and Descriptive Statistics

This research uses the data series of monthly lending rates and the deposit rates from Thai commercial banks from 1985:01 to 2015:12, as reported by the International Monetary Fund, which constitutes the post-1997 Asian financial crisis. Consequently, the results describe how Thai commercial banks behaved after the Asian financial crisis that affected financial markets internationally. The monthly Thai lending rates, deposit rates, and the resulting rate spread or intermediation premium, are denoted by LR_t , DR_t , and IP_t , respectively. Figure 1 describes the movements of LR_t , DR_t , and IP_t over the sample period.

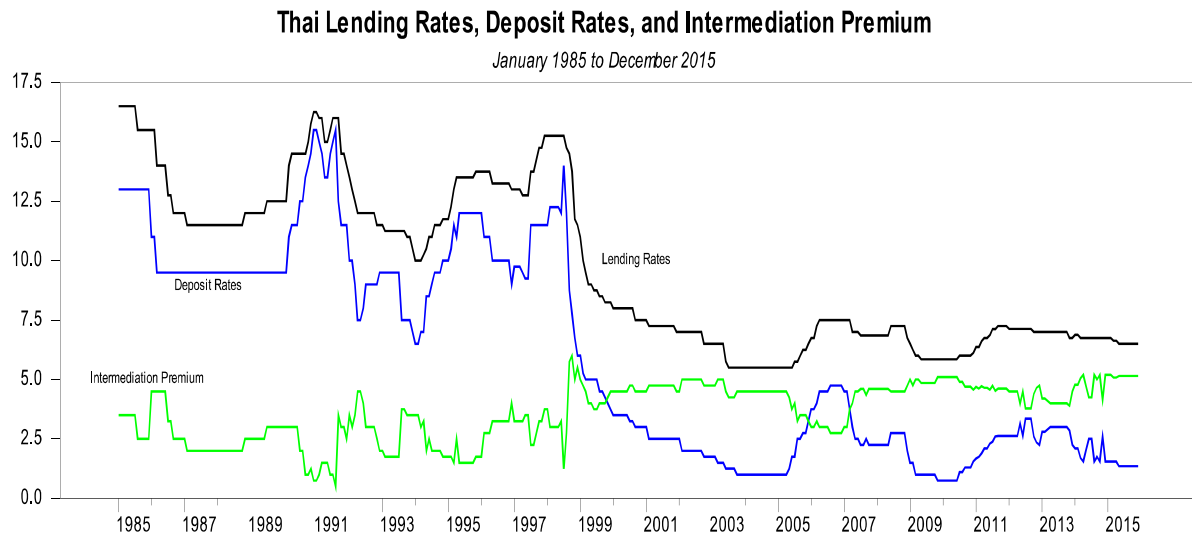


Figure 1: Thai Lending Rate, Deposit Rate, and Intermediation Premium since 1985

The mean lending rate during this period is 9.66 percent and ranges from 5.50 to 16.50, with a standard deviation of 3.40. The mean deposit rate over the same period is 6.01 percent and ranges from 7.50 to 15.50, with a standard deviation of 4.32. The mean intermediation premium during this period is 3.65 percent, and ranges from 0.50 to 6.0, with a standard deviation of 1.19. Their correlation is 98.08 percent. Figure 1 suggests the Thai lending-deposit rate spread experiences a structural change over the sample period. That structural change occurred in late 1998 and early 1999 coinciding with the Global Financial Crisis (GFC).

4. Methodological Issues and Analytical Framework

4.1 Structural Break

This study specifies and estimates Perron's (1997) endogenous unit root test function with the intercept, slope, and the trend dummy to test the hypothesis that the Thai lending-deposit rate spread has a unit root.

$$IP_t = \mu + \theta DU + \alpha t + \gamma DT + \delta D(T_b) + \beta IP_{t-1} + \sum_{i=1}^k \psi_i \Delta IP_{t-i} + v_t \quad (1)$$

where $DU = 1(t > T_b)$ is a post-break constant dummy variable; t is a time trend; $DT = 1(t > T_b)$ is a post-break slope dummy variable; $D(T_b) = 1(t = T_b + 1)$ is the break dummy variable; and ε_t are white-noise error terms. The null hypothesis of a unit root is stated as $\beta = 1$. The break date, T_b , is selected based on the minimum t -statistic for testing $\beta = 1$ (Perron, 1997).

4.2 Nonlinear Cointegration

Breitung (2001) articulated that there is often a nonlinear relationship between economic and financial time series, implying that LR_t and DR_t may be nonlinearly cointegrated. To discern this possibility, this investigation utilizes Breitung's nonparametric procedure to test for their nonlinear cointegration.

Breitung's nonparametric testing procedure consists of the cointegration test, known as the rank test for cointegration, and the nonlinearity test, referred to as the score statistic for a rank test of neglected nonlinear cointegration. To calculate the rank test for cointegration, this study first defines a ranked series as $R_T(LR_t)$ [of LR_t among $LR_1, \dots, LR_T$] and $R_T(DR_t)$. Breitung's two-sided rank test statistic, testing for the null hypothesis of no nonlinear cointegration, denoted by Ξ_T^* , is then calculated as follows:

$$\Xi_T^* = T^{-3} \sum_{i=1}^T (r_i^R)^2 / (\sigma_{\Delta r}^2) \quad (2)$$

where T is the sample size, r_i^R is the least squares residual from a regression of $R_T(LR_t)$ on $R_T(DR_t)$. According to Haug and Basher (2011), $\sigma_{\Delta r}^2$ is the variance of Δr_i^R , which is included to adjust for the potential correlation between the two time series LR_t and DR_t . The critical values for this rank test are found in Table 1 of Breitung (2001).

Given a positive result of the rank test, the first step in calculating Breitung's score statistic for a rank test of neglected nonlinear cointegration (testing for the null hypothesis of nonlinearity) is to regress the Thai lending rate, LR_t , on a constant, the deposit rate, DR_t , the ranked series of the deposit rate, $R_T(DR_t)$, and the disturbance ζ_t .

$$LP_t = \delta_0 + \delta_1 DR_t + \delta_2 R_t^*(DR_t) + \zeta_t \quad (3)$$

where $\delta_0 + \delta_1 DR_t$ is the linear part.

Under the null hypothesis, $R_t^*(DR_t) = 0$, implies that LR_t and DR_t are linearly cointegrated. Under the alternate hypothesis, $R_t^*(DR_t) \neq 0$, implies that LR_t and DR_t are nonlinearly cointegrated. The score test statistic is given by $T.R^2$, where R^2 is the coefficient of determination of the least squares regression of ζ_t , under the null hypothesis, on a constant, the ranked series of the deposit, $R_T(DR_t)$, and a disturbance term. T is the sample size. As explained by Breitung (2001), under the null hypothesis of linear cointegration, the score statistic for a rank test of neglected nonlinear cointegration is asymptotically Chi-Square distributed with one degree of freedom.

4.3 Threshold Autoregressive (TAR) Model

If the results of Breitung's nonparametric tests are positive, this study follows Thompson (2006) to regress the intermediation premium, IP_t , on a constant, a linear trend and an intercept dummy (with values of zero prior to the structural break date and values of one on the structural break date and thereafter) to formally examine the Thai LR_t , DR_t , and IP_t . The estimation results are reported in the appendix I. The saved residuals from the above estimated model, denoted by $\hat{\varepsilon}_t$, are then used to estimate the following TAR model:

$$\Delta \hat{\varepsilon}_t = I_t \rho_1 \hat{\varepsilon}_{t-1} + (1 - I_t) \rho_2 \hat{\varepsilon}_{t-1} + \sum_{i=1}^p \alpha_i \Delta \hat{\varepsilon}_{t-p} + \hat{u}_t \quad (4)$$

where $\hat{u}_t \sim i.i.d.(0, \sigma^2)$, and the lagged values of $\Delta \hat{\varepsilon}_t$ are meant to yield uncorrelated residuals. As defined by Enders and Granger (1998), the Heaviside indicator function for the TAR specification is given as:

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{t-1} \geq \tau \\ 0 & \text{if } \hat{\varepsilon}_{t-1} < \tau \end{cases} \quad (5)$$

The threshold value, τ , is endogenously determined using Chan's (1993) procedure, which obtains τ by minimizing the sum of squared residuals after sorting the estimated residuals in ascending order, and eliminating the largest and smallest 15 percent of values. The elimination of the largest and the smallest values assures that the $\hat{\varepsilon}_t$ series crosses through the threshold in the sample period.

The threshold autoregressive (TAR) model allows the degree of autoregressive decay to depend on the state of the intermediation premium, i.e. the "deepness" of cycles. The estimated TAR model reveals whether the intermediation premium reverts back to the long-run position faster when the premium is above or below the threshold. Therefore, the TAR model indicates whether troughs or peaks persist more when countercyclical monetary policy actions or economic shocks push the premium out of its long-run equilibrium path. The null hypothesis (that the intermediation premium contains a unit root) is expressed as $\rho_1 = \rho_2 = 0$, while the hypothesis that the premium is stationary with symmetric adjustments is expressed as $\rho_1 = \rho_2$.

4.4 The Asymmetric Error-Correction Model

If the results of the above asymmetric co-integration tests are positive, a Threshold Autoregressive Vector Error-Correction (TAR-VEC) model is specified and estimated to continue an investigation into any asymmetric short-run dynamic behaviours that occur between lending rates and deposit rates. Results of this model can be used to study the Granger causality between lending rates and deposit rates. The Granger causality will help to evaluate empirically (through statistics) how the Thai lending rates and deposit rates respond to the widening and narrowing of the intermediation premium caused by external economic shocks or countercyclical policy measures. Again, conventional error-correction models do not suffice for this purpose, because they do not allow the asymmetric adjustments toward the long-run equilibrium that the TAR-VEC model does.

$$\Delta LR_t = \alpha_0 + \rho_1 I_t \hat{\varepsilon}_{t-1} + \rho_2 (1 - I_t) \hat{\varepsilon}_{t-1} + A_{11}(L) \Delta LR_{t-i} + A_{12}(L) \Delta DR_{t-i} + u_{1t} \quad (6)$$

$$\Delta DR_t = \tilde{\alpha}_0 + \tilde{\rho}_1 I_t \hat{\varepsilon}_{t-1} + \tilde{\rho}_2 (1 - I_t) \hat{\varepsilon}_{t-1} + A_{21}(L) \Delta LR_{t-i} + A_{22}(L) \Delta DR_{t-i} + u_{2t} \quad (7)$$

where $u_{1,2t} \sim i.i.d.(0, \sigma^2)$ and the Heaviside indicator function is set in accordance with (5). This assumes that the Thai lending rates may respond differently depending on whether the intermediation premium is widening or narrowing as a result of expansionary, contractionary monetary policy or external shocks.

4.5 GARCH (s, r)-M Model

As aforementioned, this investigation specifies and estimates the following GARCH (s, r)-in-Mean (GARCH-M) model to discern the important question that is whether the fluctuations in the intermediation premium and hence its variance from one month affect the

premiums and variances in future months. This information is very important for countercyclical monetary policy makers with regard to whether they should intervene to bring the economy to its long-term trend less frequently and by large policy measures or more frequently and by small policy measures as these two alternative policy actions result in different variances of the intermediation premium. It is of some interest to note that GARCH-M models have been very popular and effective for modeling the volatility dynamics in many asset markets.

$$IP_t = c + \lambda \ln(\omega_t^2) + \varepsilon_t \quad (8)$$

$$\omega_t^2 = \alpha + \sum_{l=1}^r \beta_l \varepsilon_{t-l}^2 + \sum_{m=1}^s \eta_m \omega_{t-m}^2 \quad (9)$$

where IP_t is the intermediation premium, $\ln$ is the natural logarithm and ω_t^2 is its variance at time t ; ε_t is a disturbance; c is a constant; λ , α , β_l , and η_m are the parameters to be estimated of the model. The retentions of these estimated coefficients are determined by the calculated z-statistics at the 5 percent level of significance. The r and s indices are the highest subscripts l and m of retained β_l and η_m .

5. Empirical Results

5.1 Structural Break

Exhibit 1 summarizes the results of Perron's endogenous unit root tests.

Exhibit 1: Perron's Endogenous Unit Root Test, Thai Data, and 1985:01 - 2015:11

$$IP_t = 0.46016 + 0.43030 DU + 0.000042t - 0.000048 DT - 2.25837 D(T_b) + 0.80766 IP_{t-1} + v_t$$

(4.37406*) (2.93573*) (0.60332) (-0.57834) (-5.69733*)

(25.61212*)

Number of augmented lags: $k = 4$ Break Date: June 1998 $t(\alpha = 1) = -6.09957 *$

Notes: Critical values for t-statistics in parentheses. Critical values based on $n = 100$ sample for the break date (Perron, 1997). “*” indicates significance at 1 percent level.

The post-break intercept dummy variable, DU , is positive and significant at the 1 percent level. The post-break slope dummy variable, DT , is negative and statistically insignificant, while the break dummy, $D(T_b)$ is negative, and is significant at any conventional level. The time trend, t , is negative and is insignificant. These results suggest a stationary trend process, with a break date of June 1998, for the Thai intermediation premium. Moreover, strength of the test statistic, $t(\alpha = 1) = -6.09957$, confirms the structural break. This break date suggests a possible connection to the 1997 Asian financial crisis.

5.2 Results of Breitung's Nonparametric Tests

Breitung's nonparametric rank tests calculates to be 0.000000, a result that rejects the null hypothesis of no nonlinear cointegration, while the score test calculates to be 375.8201, which rejects the null hypothesis of nonlinear cointegration. These results show that, at all conventional levels of significance, Thai lending rates and deposit rates are linearly cointegrated.

5.3 Results of the Cointegration Test with Asymmetric Adjustment

Also, analysing the overall estimation results of the TAR model (summarized in Exhibit 2) indicates that the estimation results are without serial correlation and have good predicting power, as shown by the Ljung-Box statistics and the overall F -statistics, respectively. The model confirms that the Thai lending-deposit rate spread is stationary, as statistic $\Phi_{\mu} = 15.4017$ indicates that the null hypothesis of no cointegration, $\rho_1 = \rho_2 = 0$, should be rejected at the 1 percent significant level.

Exhibit 2: Unit Root and Tests of Asymmetry, Thai Data, 1985:01- 2015:11

ρ_1	ρ_2	τ	$H_0 : \rho_1 = \rho_2 = 0$	$H_0 : \rho_1 = \rho_2$	aic	sic
-0.2106*	-0.4235*	-0.60648	$\Phi_{\mu} = 15.4017^*$	$F = 3.6564^{**}$	-1.66344	-
$Q_{LB(12)} =$		$\ln L = -212.2703$		$F_{(4,365)} = 11.851$	D.W. = 2.0306	

Notes: The null hypothesis of a unit root, $H_0 : \rho_1 = \rho_2 = 0$, uses the critical values from Enders and Siklos (2001). "*" and "**" indicate 1 percent and 10 percent levels of significance, respectively. The null hypothesis of symmetry, $H_0 : \rho_1 = \rho_2$, uses the standard F distribution. τ is the threshold value determined via the Chan (1993) method. $Q_{LB(12)}$ denotes the Ljung-Box Q-statistic with twelve lags.

The results also show that both ρ_1 and ρ_2 are statistically significant at the 1 percent level. In fact, the estimation results reveal that the intermediation premium tends to decay at the rate of $|\rho_1| = 0.2106$ for $\hat{\varepsilon}_{t-1}$ above the threshold, $\tau = -0.60648$, and at the rate of $|\rho_2| = 0.4235$ for $\hat{\varepsilon}_{t-1}$ below the threshold. On the strength of the partial $F = 3.6564$, the null hypothesis of symmetry, $\rho_1 = \rho_2$, should be rejected at the 5 percent significance level, indicating statistically asymmetric adjustments around the threshold value of the Thai intermediation premium.

Specifically, the adjustment of the intermediation premium toward the long-run equilibrium tends to persist more when the premium is widening than when it is shrinking, given the finding of $|\rho_2| > |\rho_1|$. This suggests that Thai commercial banks react differently to rising deposit rates than they do to declining deposit rates. These findings may also show that these institutions react differently to expansionary monetary policy than to contractionary monetary policy, since the deposit rate is itself affected by countercyclical monetary policies. The empirical results further indicate the predatory pricing behaviour of the Thai lending institutions. These results also parallel those reported in advanced and emerging economies. Furthermore, these empirical findings support the aforementioned consumer characteristic and consumer reaction hypotheses.

5.4 Results of the Asymmetric Error-Correction Model

The estimation results of the TAR-VEC model, specified by Equations (5), (6), and (7), using the Thai lending rates and the deposit rates are summarized in Exhibit 3. Therein, $A_{ij}(L)$ is the first-order polynomials in the lag operator L . F_{ij} is the calculated F -statistic (with the p-value in brackets), which tests the null hypothesis that all coefficients of A_{ij} are equal to zero. $Q_{(12)}$ is the Ljung-Box statistic (with its significance in brackets), which tests whether the first twelve of the residual autocorrelations are both equal to zero. $\ln L$ is the log likelihood.

The empirical results suggest that the estimated equations (6) and (7) are without serial correlation and have good predicting power, as shown by the Ljung-Box statistics and the overall *F-statistic*, respectively. The estimation results of equation (6) of the TAR-VEC model indicate that both ρ_1 and ρ_2 are insignificant at conventional levels. This finding shows that the Thai lending rates do not respond to the intermediation premium both when it widens and when it narrows in the long run, and suggests that Thai lending institutions do not respond to both expansionary and contractionary monetary policy in the long run. Regarding the long-term adjustment of the deposit rates, the estimation results of Equation (7) show that both $\tilde{\rho}_1$ and $\tilde{\rho}_2$ are significant at the 1 percent level, suggesting that Thai deposit rates are affected by lending rates in the long run.

Exhibit 3: Thai Lending and Deposit Rates, Monthly Data, 1985:01- 2015:11

$\Delta LR_t = 0.0059 + 0.0257 I_t \hat{\varepsilon}_{t-1} + 0.0487 (1 - I_t) \hat{\varepsilon}_{t-1} + A_{11}(L) \Delta LR_{t-1} + A_{12}(L) \Delta DR_{t-1} + u_{1t}$		
(0.3610)	(0.7209)	(1.8733) $F_{11} = 5.720[0.0000]$
$F_{12} = 15.4442[0.0000]$		
$Q_{(12)} = 4.9580[0.9594] \quad \ln L = 47.2998 \quad F_{(15,333)}\text{-statistic} = 12.3307^*$		
$\Delta DR_t = 0.0110 + 0.1815 I_t \hat{\varepsilon}_{t-1} + 0.2134 (1 - I_t) \hat{\varepsilon}_{t-1} + A_{21}(L) \Delta LR_{t-1} + A_{22}(L) \Delta DR_{t-1} + u_{2t}$		
(2.8060*)	(2.6490*)	(4.2569*) $F_{21} = 9.1551[0.0000]$
$F_{22} = 7.1865[0.0000]$		
$Q_{(12)} = 8.066 [0.7799] \quad \ln L = -193.7717 \quad F_{(11,336)}\text{-statistic} = 8.8656^*$		

Note: * indicates 1 percent level of significance.

In addition to estimating the long-run equilibrium relationship and asymmetric adjustment, the estimated TAR-VEC model also allows for determinations of the Granger causality between the Thai lending rates and the deposit rates. Equation (6) reveals in the partial *F-statistics* that the lending rate responds to its own lagged changes, as well as the lagged changes in the deposit rate. Similarly, the estimation results show that the deposit rate responds to both its own lagged changes and lagged changes of the lending rates. These findings suggest a bidirectional Granger-causality between the Thai lending rate and the deposit rate in the short run, and reveal that the Thai lending rate and the deposit rate affect the movements of each other's rates in the short run,

5.5 GARCH (s, r)-M Model

As aforementioned, the retentions of the estimated coefficients of equations (7) and (8) are determined by the calculated z-statistics at the 5 percent level of significance. The *r* and *s* indices are the highest subscripts *l* and *m* of retained β_l and η_m which are *l*=2 and *m*=2, respectively. The values of *l* and *m*, in turn, suggest GARCH (2, 2) be the best model for this investigation. The estimation results of the GARCH (2, 2)-M model are reported in Exhibit 5

An analysis of the estimation results of the GARCH (r, s)-M model suggests the presence of GARCH (2, 2) effect on the Thai intermediation premium and its variance. Financially, the empirical results indicate that the fluctuations in the Thai premium and hence its variance from the one month affect the premia and the variances in the subsequent months.

Exhibit 5- GARCH (2, 2)-M Model Results, Thai Data, 1985:01 - 2015:11

$$IP_t = 3.5756 - 0.5134 \ln(\omega_t^2) + \varepsilon_t$$

$$(49.3331^*) \quad (-14.2066^*)$$

$$\omega_t^2 = 0.0505 + 0.6022 \varepsilon_{t-1}^2 + 0.4172 \varepsilon_{t-2}^2 - 0.32473 \omega_{t-1}^2 + 0.2853 \omega_{t-2}^2$$

$$(4.8873^*) \quad (10.0181^*) \quad (3.7631^*) \quad (-2.2472^{**}) \quad (6.4098^*)$$

Notes: Akaike info criterion=1.9699; Schwarz criterion =2.0436; Hannan-Quinn Criterion=1.9991; Log likelihood = -359.3916; Durbin-Watson Statistic = 0.2493.

“*” and “**” indicate the 1 percent and 5 percent levels of significance, respectively.

6. Concluding Remark and Policy Implications

This study investigates the behaviour of Thai lending rates, deposit rates and the intermediation premium in the post-1997 Asian financial crisis era, by estimating the threshold autoregressive (TAR) model developed by Enders and Siklos (2001).

First, the study tested the hypothesis that the Thai intermediation premium has a unit root by specifying and estimating Perron's (1997) endogenous unit root test function with the intercept, slope, and trend. This test suggested that the premium followed a stationary trend process with a structural break in July 2008, which may be attributable to the influence of the U.S. subprime financial crisis.

Second, the study tested whether Thai lending rates and deposit rates are linearly and/or nonlinearly cointegrated. Breitung's nonparametric rank tests reveal nonlinear cointegration at all conventional levels of significance.

Third, the estimation results of the TAR model reveal that Thai commercial banks react differently to rising versus declining deposit rates. These findings suggest that these institutions react differently to expansionary rather than contractionary monetary policy. Furthermore, these results on asymmetric responses reveal the predatory pricing behaviour of the Thai institutions.

Finally, the study tested for Granger causality between the lending rate and the deposit rate in the short run by the empirical estimation of the TAR-VEC model. This revealed bidirectional Granger causality, and indicates that the lending rate and the deposit rate affect each other's movement in the short run. This finding also reveals asymmetric responses of financial markets to contractionary and expansionary monetary policy actions, and confirms the ability of the Thai monetary authority to use its countercyclical monetary policy instruments to achieve its short-run macroeconomic objectives. However, the GARCH (2,2)-M model results suggest that monetary policymakers should intervene to bring the economy to its long-term trend more frequently and by small policy measures to minimize the conditional variance of the intermediation premium to minimize the magnitude of the cycle of the lending rate.

The modeling utilized in this study is likely one of the most impactful methodologies for use by Thai policymakers in the monitoring overall debt funding costs in the Thai economy. Utilization of these models for analysis or reported data helps capture all factors when accounting for costs associated lending and deposit-taking activities. Future research by accountants can benefit from understanding economic modeling as well as the link between fair value accounting (on securities) to better inform policy decisions.

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Appendix I

Exhibit 4: Estimation Results, Thai Monthly Data, 1985:01 - 2015:11

$$IP_t = 5.9443 - 0.0058Trend_t - 2.0561Dummy_t + \varepsilon_t$$

(42.4756*) (-4.4247*) (-9.6356*)

$$\ln L = -301.7137 \quad R^2 = 0.6001 \quad DW \text{ statistic}^{(a)} = 0.3382 \quad F_{(2,223)} = 167.3195^*$$

Notes: “*” indicates significance at 1 percent level.

(a) As articulated by Enders and Siklos (2001, p. 166), in this type of model specification, ε_t may be contemporaneously correlated.