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Behavior of lending and deposit rates in the economy where conventional and Islamic finance systems coexist: the case of Malaysia

Abstract

Nonparametric test statistics document the nonlinear cointegration between the Malaysian lending and deposit rates. Empirical investigation further reveals that their spread asymmetrically adjusted to the threshold faster when the deposit rate fell relative to the lending rate than when the deposit rate moved in the opposite direction. These findings may be attributable to robust Islamic and conventional financial systems that operate parallel to each other. These empirical findings also suggest that Malaysian lending institutions in the conventional segment of the economy do not exhibit predatory pricing behavior. Furthermore, empirical results indicated that the Malaysian lending rate and deposit rate adjustments affected each other's movements. As to the long-run behavior of lending and deposit rates, the empirical findings indicated that while the Malaysian lending rate responded to both, the deposit rate responded to neither the expansionary monetary policy, which widens the lending-deposit rate spread; nor the contractionary policy, which narrows the spread in the short run.

Keywords: Islamic finance, asymmetry, lending rate, deposit rate, lending-deposit rate spread, nonlinear cointegration, Malaysia.

JEL Classification: C22, E44, G21.

Introduction

Since independence in 1957, the most important socioeconomic issue in Malaysia has been the improvement of the economic standing of ethnic Malays and other indigenous people within the country, collectively known as "bumiputras". In an effort to eradicate poverty and end the identification of economic function with ethnicity, the Malaysian government established the national economic policy known as the New Economic Policy (NEP) in 1971. The NEP of 1971 and the concerted effort of the government to develop its Islamic financial industry and institutions defined the characteristics and the complexity of the Malaysian economy and have differentiated it from those of developed and emerging countries. Malaysia became one of the largest – if not the largest – international Islamic financial centers where robust Islamic and conventional financial systems operate parallel to each other.

The central feature of the Islamic financial operation is that Shariah, the Islamic law, prohibits the payment or acceptance of a fixed or predetermined rate of interest, known as *riba*. This central characteristic of Islamic finance alters the "normal" relationship between lending rates and deposit rates. These distinctive characteristics of the Malaysian financial system create enormous intellectual curiosity and lead to questions as to how the lending institutions set their lending and deposit rates. To formally investigate this matter, Breitung's nonparametric testing procedure is used to test for the possible nonlinear relationship between the Malaysian lending and deposit rates. Given the positive results of the cointegration tests, this study utilizes Enders and Siklos (2001) procedure to test for asymmetric cointegrat-

ing relationships and the Granger causality between the Malaysian lending and deposit rates.

The remainder of this paper is organized as follows. Section 1 briefly reviews the literature. Section 2 characterizes the Malaysian financial sector and its operational environment. Section 3 summarizes data for this study. Section 4 briefly describes the methodology that will be used in the investigation. Section 5 reports Breitung's rank test statistics and empirical test results for cointegration allowing for asymmetric adjustment to a threshold. It presents the results of the cointegration and asymmetry tests. Section 6 examines the results of the asymmetric error-correction model to determine the Granger causality between the lending rates, and the deposit rates, and the Final section provides some concluding remarks.

1. A brief literature review

As articulated by Thompson (2006), commercial banks may set their lending rates as some markup or premium over their deposit rates. If the premium or the intermediation margin is perceived to be too high or too low, the market forces will discipline banks to adjust back to some equilibrium spread. The evidence of asymmetric rate-setting behavior in the banking industry supports the literature hypothesizing the asymmetric effects of monetary policy on output. There are three main theoretical explanations for commercial bank interest rate asymmetries: bank concentration hypothesis, consumer characteristic hypothesis, and consumer reaction hypothesis¹.

The bank concentration hypothesis posits that banks in more concentrated markets are slower to adjust deposit rates upward and faster to adjust them downward while exhibiting the opposite behavior

¹ Scholnick (1999) provides the survey on these three types of explanations for commercial banks' interest rate asymmetries in the literature.

regarding lending rates (Neumark and Sharpe, 1992; Hannan and Berger, 1991). The consumer characteristic hypothesis asserts that the greater the proportion of unsophisticated consumers relative to sophisticated consumers in the market is, together with the potential search and switching costs, the greater is the banks' ability to adjust interest rates to their advantage (Hutchison, 1995; Rosen, 2002).

Interestingly, the asymmetric adjustment in lending rates may be influenced by a further asymmetry. As mentioned by Stiglitz and Weiss (1981), the presence of asymmetric information may create an adverse selection problem in lending markets such that higher interest rates will tend to attract riskier borrowers. Therefore, banks would be reluctant to raise lending rates, even if market rates rise. The expected cost to the banks of not raising the lending rates when their marginal cost of fund increases, will be offset by the benefits from not encouraging the higher-risk consumers to borrow.

Additionally, in traditional market economies, the asymmetric rate-setting in the context of rates of return on financial market instruments in developed economies has been empirically documented in financial economics literature. 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 the 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 Philippine bank lending and deposit rates. Nguyen et al. (2010) found asymmetries in the Bangladeshi lending deposit rates. Chang and Su (2010) reported nonlinear cointegration between the lending and the deposit rate in ten Eastern European countries. Lately, Haug and Basher (2011) found nonlinear cointegration in the purchasing power parity relationships for Canada, Japan, Switzerland, the UK, Belgium, France, Germany, Italy, and the Netherlands.

2. The Malaysian financial sector and operating environment

As aforementioned, the Malaysian government established the national economic policy known as the New Economic Policy (NEP) in 1971 in an effort to

eradicate poverty and end the identification of economic function with ethnicity, the government provided funds to the bumiputras population to purchase foreign-owned shareholdings. Bank Bumiputras was also created to promote the economic development of the indigenous population. The National Development Policy, that followed the NEP, also contained many of the NEP's goals, although without specific equity targets and timetables (Malaysia Country Study Guide, 2007, International Business Publications, USA).

Since the early 1980s, Malaysian policymakers have supported the development of an Islamic financial industry with strong Islamic institutions. The government established regulatory and tax systems to support the market-driven environment in which Islamic finance coexists with conventional finance. As pointed out in the recent IMF Country Report on Malaysia (IMF Country Report No. 09/253, 2009, p. 14), Malaysia has become one of the largest – if not the largest – international Islamic financial centers and now has robust Islamic and conventional financial systems that operate parallel to each other.

The assets of the Islamic banks have doubled since 2000, and accounted for about 17 percent of total banking sector assets as of May 2009. Shariah-compliant stocks account for about 88 percent of stocks listed and 64 percent of total market capitalization of the Malaysian stock market. *Takaful* (Islamic insurance) operators have a 7 percent share of total insurance and *takaful* assets, and about 13 percent of funds managed by unit trust management companies are Islamic. Malaysia hosts the world's largest *sukuk* (Islamic bonds) market (estimated at RM 155 billion, or 59 percent of total outstanding bonds in Malaysia). As of the end of 2008, Malaysian *sukuk* accounted for about 61 percent of the total global *sukuk* outstanding (both domestic and international issues). In 2008, Malaysia also led in terms of global *sukuk* issuance with a share of 53 percent, followed by the United Arab Emirates, Saudi Arabia, and Bahrain. In other sectors of Islamic finance, international Islamic stock indices have been developed and new licenses issued to foreign Islamic banks, Islamic fund management companies, and *takafuls* (IMF Country Report No. 09/253, 2009, p. 14). The central feature of the Islamic financial operation is that Shariah, the Islamic law, prohibits the payment or acceptance of a fixed or predetermined rate of interest that is known as *riba*. But in spite of this, the Malaysian Islamic financial sector has grown into one of the robust financial markets in the world.

3. The data

To empirically investigate the Malaysian lending and deposit rate behavior, this analysis utilizes monthly

data from International Financial Statistics, published by the IMF, over the period from January 1991 to October 2010. The monthly lending rates and the deposit rates are denoted by LR_t and DR_t , respectively.

Figure 1 displays the behavior of the respective lending and deposit rates over the sample period. As Figure 1 suggests, the Malaysian lending and deposit rates oscillated around an upward trend from the beginning of the sample period to early 1992. They then fluctuated around fairly steep

downward trends through the middle of 1994. These rates again oscillated around their very steep upward trends until the beginning of the Asian financial crisis in the middle of 1997. Over the duration of the crisis, these rates took huge jumps to their highest peaks ever, dropped sharply to their pre-crisis levels in early 1999, and then oscillated around their moderately downward trends until the end of 2008. Both rates took a sharp drop and then stabilized at their low levels before increasing significantly at the beginning of 2010.

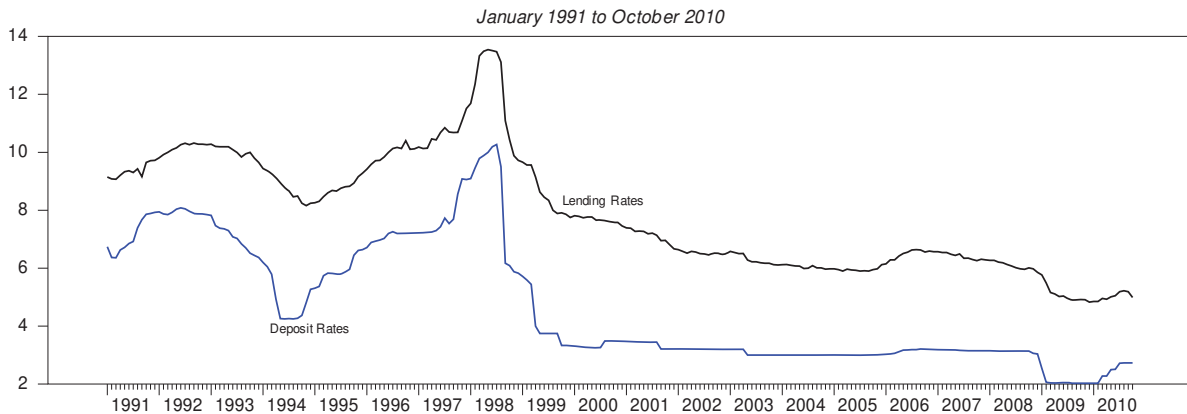


Fig. 1. Malaysian lending rates and deposit rates

The mean lending rate during the sample period was 7.87 percent, and ranged from 4.83 percent to 13.54 percent with the standard error being 2.03 percent. The mean deposit rate over the same period was 4.69 percent, and ranged from 2.03 percent to 10.27 percent with the standard error being 2.15 percent. Their correlation was 95.56 percent.

4. Methodology

As mentioned, the Malaysian economy and its financial sector are distinctly different from both developed and emerging economies and have gone through many changes and experienced many economic shocks; therefore, it is possible that the lending-deposit rate relationship might experience structural breaks over the sample period. To discern this possibility, this study followed Perron's (1997) procedure to specify and estimate the following endogenous unit root test function with the intercept, slope, and the trend dummies to test the hypothesis that the Malaysian lending rate has experienced structural shifts over the sample period. The equation's specification and the estimation results are reported in Appendix A.

The empirical results of these tests suggest that the Malaysian lending-deposit rate relationship followed a stationary trend process with a break date of September 1997, corresponding to the worldwide financial chaos created by the Asian financial crisis. However, the test statistics failed to support this suggested structural shift at conventional levels of significance. To reevaluate the strong suggestion of

the data in Figure 1, the Chow test was performed and the result statistically confirmed the structural shift at all conventional significant levels. Furthermore, as posited by Breitung (2001, p. 331), economic theory suggests in many cases a nonlinear relationship between economic and financial time series. This implies that we have to test for the nonlinear cointegration. The analysis of the behaviors of and the cointegration between the Malaysian lending and deposit rates as well as their spread should be done next if the results of these tests are statistically significant. In this study, we use Breitung's nonparametric procedure to test for 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. Following Breitung (2001), this study defines a ranked series as $R_T(LR_t)$ [of LR_t among LR_1, \dots, LR_T] and $R_T(DR_t)$ accordingly. Breitung's two-sided rank test statistic, testing for cointegration, denoted by Ξ_T^* , is calculated as follows:

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

where T is the sample size, r_i^R is the least squares residual from a regression of $R_T(LR_t)$ on $R_T(DR_t)$. As pointed out by Haug and Basher (2011, p. 187),

$\sigma_{\Delta r}^2$ is the variance of Δr_t^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 given in Table 1 in Breitung (2001, p. 334).

Given the 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 nonlinearity) is to regress the Malaysian 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 .

$$LR_t = \delta_0 + \delta_1 DR_t + R_t^*(DR_t) + \zeta_t, \quad (2)$$

where $\delta_0 + \delta_1 DR_t$ is the linear part. Under the null hypothesis, $R_t^*(DR_t) = 0$ implying that LR_t and DR_t are linearly cointegrated. Under the alternate hypothesis, $R_t^*(DR_t) \neq 0$ implying that LR_t and DR_t are nonlinearly cointegrated. The score test statistic is given by $T \cdot R^2$ from the least squares regression of ζ_t on a constant, the deposit rate, DR_t , the ranked series of the deposit rate, $R_t(DR_t)$, and a disturbance term. T is again the sample size and R^2 is the coefficient of determination of the regression. As articulated by Breitung (2001, p. 337), 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.

If the results of Breitung's nonparametric tests are positive, the threshold autoregressive (TAR) model

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

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

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

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 15 percent of the largest and the small

developed by Enders and Siklos (2001), specified by the following equations (3), (4) and (5), is estimated to formally examine the behavior of the Malaysian lending rates, deposit rates, and their spread. The threshold autoregressive model allows the degree of autoregressive decay to depend on the state of the Malaysian lending-deposit rate spread in the previous period (i.e., "deepness" of cycles). Empirical result reveals if the spread tends to revert to the long-run position faster when it is above or below the threshold. Therefore, the TAR model indicates whether troughs or peaks persist more when shocks push the Malaysian lending-deposit rate spread out of its long-run path.

Enders and Siklos (2001) extended the popular two-step symmetric Engle-Granger's (1987) methodology to test for long-run relationships between two time series allowing for asymmetry. As demonstrated by Enders-Siklos (2001), the first step in the procedure is to estimate the following long-run relationship between the Malaysian lending rate and deposit rate using ordinary least squares.

$$LR_t = \beta_0 + \beta_1 DR_t + \beta_2 Dummy_t + \beta_3 Trend_t + \epsilon_t, \quad (3)$$

where LR_t and DR_t are denoted as the lending rate and the deposit rate, respectively. $Dummy_t$ is a dummy (with values of zero prior to September 1997 and values of one for September 1997 and thereafter). $Trend_t$ is a linear trend. The estimation results are reported in Appendix B. The saved residuals, ϵ_t from the estimation of equation (3), are defined as the Malaysian lending-deposit rate spread, denoted by $\hat{\epsilon}_t$, and are then used to estimate the following TAR model:

lest values. The elimination of the largest and the smallest values is to assure that the $\hat{\epsilon}_t$ series crosses through the threshold in the sample period. As pointed out by Thompson (2006, p. 326), if the TAR model rejects the null hypothesis of symmetric adjustment, the sample mean is a biased estimate of the threshold. Chan's method yields the consistent estimate of the threshold. In this model's specification, the null hypothesis that the Malaysian lending-deposit rate spread contains a unit root can be expressed as $\rho_1 = \rho_2 = 0$, while the hypothesis that the spread is stationary with symmetric adjustments can be stated as $\rho_1 = \rho_2$.

5. Empirical results

Empirical calculations indicate that Breitung’s non-parametric rank tests and score test are 0.00045 and 72.921, respectively. These statistics reveal that the Malaysian lending and deposit rates are nonlinearly cointegrated at all conventional levels of significance. Additionally, the estimation results of the TAR model are summarized in Table 1. An analysis of the overall empirical results indicates that the estimation results are devoid of serial correlation and have good predicting power, as evidenced by the Ljung-Box statistics and the overall F -statistics, respectively. With the calculated statistic $\Phi_{\mu} = 13.7792$, the null hypothesis of a unit root ($\rho_1 = \rho_2 = 0$) is rejected at the 1 percent significance level (i.e., the spread is stationary). As to the

speed of adjustment, based on the partial test statistic $F = 19.1461$, the null hypothesis of symmetry, $\rho_1 = \rho_2$, is rejected at 1 percent level. Thus, the empirical results indicate that adjustments around the long-run threshold value of the Malaysian lending-deposit rate spread are asymmetric. In fact, the point estimates suggest that the spread tends to decay at the rate of $|\rho_1| = 0.5499$ for $\hat{\epsilon}_{t-1}$ above the threshold, $\tau = 0.3789$, and at the rate of $|\rho_2| = 0.0525$ for $\hat{\epsilon}_{t-1}$ below the threshold. Additionally, while ρ_1 is statistically significant at 1 percent level, ρ_2 is insignificant at any conventional level. Furthermore, the estimates of ρ_1 and ρ_2 satisfy the stationary (convergence) conditions¹.

Table 1. Unit root and tests of asymmetry, Malaysian monthly data January 1991 to October 2010

ρ_1	ρ_2	τ	$H_0 : \rho_1 = \rho_2 = 0$	$H_0 : \rho_1 = \rho_2$	Aic
-0.5499*	-0.0525	0.3789	$\Phi_{\mu} = 13.7792$ *	$F = 19.1461$ *	-3.7859
$Q_{LB(12)}=16.1510[0.3162]$			$\ln L = 70.8177$	$F_{(4,231)}=9.7339$ *	

Notes: The null hypothesis of a unit root, $H_0 : \rho_1 = \rho_2 = 0$, uses the critical values from (2001, p. 170, Table 1, for four lagged changes and $n = 100$). * indicates 1 percent level of significance. 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 12 lags.

With regard to the adjustment process, given $|\rho_1| > |\rho_2|$, the Malaysian lending-deposit rate spread adjusts to the threshold value faster when monetary policy action or economic shock causes the deposit rates to fall relative to the lending rates, widening the spread, than when the deposit rates move in the opposite direction, narrowing the spread. These findings contradict those reported by Thompson (2006) in the U.S. with respect to the prime rate and the secondary market one-month CD rate. Interestingly, these empirical results seem to support the consumer-reaction hypothesis articulated by Stiglitz and Weiss (1981) which may be attributable to the influences of the co-existence of the Is-

lamic financial industry and the traditional financial institutions. These empirical findings also suggest that Malaysian lending institutions in the conventional segment of the economy do not exhibit predatory pricing behavior.

6. Results of the asymmetric error-correction model

The presence of asymmetric adjustments in the Malaysian lending-deposit rate spread, as indicated by the above estimation results, allows the use of a TAR VEC model to further investigate the short-run and long-run dynamics with respect to the lending rate (LR_t) and the deposit rate (DR_t).

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

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

where $u_{i,t} \sim i.i.d. (0, \sigma^2)$, $i = 1, 2$ and I_t is set in accordance with equation (5).

As pointed out by Thompson (2006, pp. 327-328), the above specified TAR VEC model differs from the convention error-correction models by allowing asymmetric adjustments toward the long-run equilibrium. Also, the asymmetric error correctional model replaces the single symmetric error correction term with two error correction terms. Thus, in addition to estimating the long-run equilibrium relationship and asymmetric adjustment, the model

also allows for tests of the short-run dynamic between changes in lending rates and deposit rates. This in turn reveals the nature of their Granger causality.

The estimation results of the asymmetric error correction model are reported in Table 2. In the summary of the estimation results, $A_{ij}(L)$ represents the first-order

¹ As shown by Petrucelli and Woolford (1984), the necessary and sufficient condition for the basis to be stationary is: $\rho_1 < 0$, $\rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2) < 1$.

polynomials in the lag operator L . The F_{ij} represents the calculated partial F -statistics with the p-value in square brackets testing the null hypothesis that all coefficients of A_{ij} are equal to zero. The t -statistics are reported with “***” and “**” respectively indicating the 5 and 10 percent significant

levels. $Q_{LB(12)}$ is the Ljung-Box statistics and its significance is in square brackets, testing for the first twelve of the residual autocorrelations to be jointly equal to zero. $\ln L$ is the log likelihood. The overall F -statistic with “*” indicates the significance level of 1 percent.

Table 2. Asymmetric error correction model, Malaysian monthly data, January 1991 to October 2010

$\Delta LR_t = -0.0075 - 0.0670I_t\hat{\varepsilon}_{t-1} - 0.0754(1 - I_t)\hat{\varepsilon}_{t-1} + A_{21}(L)\Delta LR_{t-i} + A_{22}(L)\Delta DR_{t-i} + u_{1t}$ -0.6073 -1.7710*** -1.9720** $F_{11}=14.9453[0.0000]$ $F_{21}=17.9627[0.0000]$		
$Q_{LB(12)} = 7.3970[0.8303]$	$\ln L = 109.3821$	$F_{(10,215)} = 17.6275^*$
$\Delta DR_t = -0.0303 + 0.0498I_t\hat{\varepsilon}_{t-1} - 0.0541(1 - I_t)\hat{\varepsilon}_{t-1} + A_{21}(L)\Delta LR_{t-i} + A_{22}(L)\Delta DR_{t-i} + u_{2t}$ -1.5087 0.8369 -0.8209 $F_{21}= 12.8952[0.000]$ $F_{22}=21.7026[0.0000]$		
$Q_{LB(12)} = 8.2660[0.7640]$	$\ln L = 0.8376$	$F_{(9,213)} = 10.7374^*$

An analysis of the overall empirical results indicates that the estimated equations (6) and (7) are devoid of serial correlation and have good predicting power as evidenced by the Ljung-Box statistics and the overall F -statistics, respectively. As to the short-run dynamic adjustment, the calculated partial F -statistics in equations (6) and (7) indicate bidirectional Granger-causality between Malaysian lending and deposit rates. These results imply that the Malaysian lending rate and deposit rate adjustments affected each other’s movements, which parallel those reported by Thompson with respect to the prime lending rate and the one-month CD rate in the U.S. banking industry.

In addition to revealing the short-run dynamic Granger-causality, the asymmetric error correction model also allows the investigation of the long-run adjustments of the lending rate. The empirical results of the TAR model revealed $|\rho_1| > |\rho_2|$ in equation (6), which indicated that the lending rate adjusted to the long-run equilibrium slower when the shock widened than when it narrowed the lending-deposit rate spread. Economically, this result seems to suggest that the Malaysian lending rates respond to both contractionary and expansionary countercyclical monetary policy in the long run. With regard to the long-run adjustment of the deposit rate, the estimation results for equation (7) show that $|\tilde{\rho}_2| > |\tilde{\rho}_1|$; however, both $\tilde{\rho}_1$ and $\tilde{\rho}_2$ are statistically insignificant at any conventional significant level. These findings seem to indicate that the deposit rate responds to neither the expansionary monetary policy, which widens the lending-deposit rate spread, nor the contractionary monetary policy, which narrows the spread in the long run.

Conclusion

This study used Breitung’s nonparametric testing procedure to document the nonlinear cointegration between the Malaysian lending and deposit rates. A threshold autoregressive (TAR) model (Enders and Siklos, 2001) was used to investigate the behavior

of the Malaysian lending and deposit rates and their spread. First, following Perron’s (1997) procedure, an endogenous unit root test function with the intercept, slope, and trend were specified and estimated to test the hypothesis that the Malaysian lending rate has a unit root. The results of this test suggest that the lending rate may have experienced a possible structural break in September 1997. The Chow test result confirmed this structural shift.

The TAR model’s estimation results indicate that the adjustments of the Malaysian lending-deposit rate spread toward the long-run equilibrium are asymmetric and tend to adjust faster when the deposit rate is declining than when it is rising. The empirical findings may be attributable to the fact that Malaysia has become one of the largest – if not the largest – international Islamic financial centers. As a result, Malaysia now has robust Islamic and conventional financial systems that operate parallel to each other. These empirical findings also suggest that Malaysian lending institutions in the conventional segment of the economy do not exhibit predatory pricing behavior.

Furthermore, the estimation results of the Asymmetric Error-Correction Model suggested a bidirectional Granger-causality between Malaysian lending and deposit rates. These results imply that both the Malaysian lending and deposit rates adjustments affected each other’s movements, which parallel those reported by Thompson with respect to the prime lending rate and the one-month CD rate in the U.S. banking industry.

As to the long-run behavior of lending and deposit rates, the empirical results indicated that while the Malaysian lending rate responded to both, the deposit rate responded to neither the expansionary monetary policy, widening the lending-deposit rate spread, nor the contractionary monetary policy, narrowing the spread in the long run.

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Appendix A

To endogenously search for the structural break possibility in the time series data LR_t , Perron (1997) procedure with the intercept, slope, and trend dummy is specified as:

$$LR_t = \mu + \theta DU + \alpha t + \gamma DT + \delta D(T_b) + \beta LR_{t-1} + \sum_{i=1}^k \psi_i \Delta LR_{t-1} + v_t, \quad (6)$$

where $DU = 1(t > T_b)$ is a post-break constant dummy variable; t is a linear 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 v_t is white-noise error term. The break date, T_b , is selected based on the minimum t-statistic for testing $\beta = 1$ (see Perron, 1997, pp. 358-359). The estimation results using the Malaysian lending rate LR_t are summarized in Table 3.

Table 3. Perron's endogenous unit root test, Malaysian data, January 1991 to October 2010

$LR_t = 0.3936 + 0.2701DU - 0.0001t - 0.0014DT - 1.0798D(T_b) + 0.8734LR_{t-1} + v_t$ (4.5214*) (2.8577*) (-0.1034) (-1.1183) (-6.0270*) (30.7879*)		
No. of augmented lags: k = 12	Possible Break Date: September 1997	$t(\alpha = 1) = -4.4619$

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

Appendix B

Table 4. Estimation results for equation (3), Malaysian data, January 1991 to October 2010

$LR_t = 4.6145 + 0.8024DR_t + 1.1061Dummy_t - 0.0103Trend_t + \varepsilon_t$ (22.913*) (32.2882*) (9.4521*) (-10.4254*)			
$\ln L = -164.5434$	$R^2 = 0.2792$	DW statistic ^(a) = 0.1605	$F_{3,234} = 1291.5747^*$

Notes: * indicates significance at the 1 percent level. As articulated by Enders and Siklos (2001, p. 166), in this type of model specification, ε_t may be contemporaneously correlated.