Currency Regime and Monetary Independence in the Post-Crisis East Asia: An Application of Regime-Switching Model with Endogenous Explanatory Variables

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Abstract

This paper investigates whether the choice of exchange rate regimes influences the sensitivity of domestic interest rates to U.S interest rates in the East Asian countries. We employ a regime-switching model that allows for the possibility of a structural break in the extent of monetary autonomy at an unknown time and the endogeneity of U.S. interest rate shocks. We find that the sensitivity of local to U.S interest rates has declined in Korea and Thailand after they adopted floating exchange rate regimes. We also find that Japan with a floating exchange rate regime has greater independence in monetary policy than a pegged economy like Hong Kong throughout the period since 1987. These empirical findings suggest that exchange rate flexibility provides a larger extent of monetary independence.

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I. Introduction

The choice of an appropriate exchange rate regime is one of the most fundamental policy issues in open economies. Each regime has advantages and disadvantages. Fixing the exchange rate helps to reduce transaction costs and exchange rate risks. It can also work as a credible nominal anchor for monetary policy. On the other hand, a floating exchange rate regime allows the domestic monetary authority to pursue an independent monetary policy.

The issue of monetary independence under floating exchange regimes, however, has been controversial. In theory, under a credible pegged exchange rate regime with perfect capital mobility, domestic interest rates cannot be set independently. The 'impossible trinity' says that countries can pursue two of three options—fixed exchange rates, domestic monetary autonomy, and capital mobility. Thus, without restrictions on capital flows, the monetary authority that wishes to retain domestic interest rate as a policy instrument must adopt a flexible exchange rate arrangement. But, even if formally floating, countries may not allow exchange rate to fluctuate beyond the magnitude that they can tolerate. This 'fear of floating' view argues that the floating countries actually behave much like they would in a pegged system by keeping domestic interest rates close to the base rate (Calvo and Reinhart, 2002).

The purpose of this paper is to examine empirically the independence of monetary policy in the context of East Asian economies. In the wake of the financial crisis, most of crisis-affected East Asian economies including Indonesia, Korea, the Philippines and Thailand shifted their exchange rate regimes from *de facto* US dollar pegs to floating ones. However, critics say that East Asian central banks have intervened heavily in foreign exchange rate markets in order to prevent the appreciation of their currencies for the pursuit of an ‘export-led growth strategy’ (Dooley, Folkerts-Landau, and Garber, 2003). If so, the *de jure* free-floating system may not have enabled East Asian countries to retain their monetary autonomy.

Recent empirical studies—including Frankel (1999), Hausman, Panizza, and Stein (1999), Borensztein, Zettelmeyer, and Philippon (2001), Frankel, Schmukler and Serven (2002), and Shambaugh (2004)—formally investigate whether the choice of the currency regime affects monetary policy independence in practice. These studies estimate the sensitivity of local interest rates to changes in international interest rates, examining whether rates are less sensitive to changes under floating exchange rate regimes than under pegged regimes. In principle, under floating regimes, changes in the exchange rates would absorb the effects of international interest rate shocks, and thereby provide "insulation" for domestic interest rates. These existing studies, however, do not focus on East Asia. Frankel et al. (2002) find that in the long run local interest rates are adjusted fully to international interest rates regardless of the exchange rate regime for Hong Kong, Singapore, Thailand, and the Philippines, despite the differences in their exchange regimes. But, the research does not consider the change of regimes after the crisis explicitly.

Our methodology for empirical investigation follows closely those in the previous literature, but improve them in several ways. First, our empirical techniques allow for the possibility that there is a structural break in the sensitivity of local interest rates to changes in international interest rates over the sample period, and furthermore that the timing of actual structural break in the degree of monetary independence and the timing of *de jure* exchange regime change may not coincide. We actually estimate the timing of the potential structural break in monetary autonomy. Second, our model takes account of the endogeneity of international interest rate shocks.

Our empirical results show that in some East Asian countries such as Korea and Thailand, the sensitivity of local to U.S. interest rates declined after they adopted floating exchange rate regimes. We also find that a floater such as Japan has had more monetary autonomy than a pegged economy such as Hong Kong since 1987 up to present. These empirical findings suggest that the choice of exchange rate regime is an important factor for the independence of monetary policy. Floating regimes appear to provide some degree of monetary independence after the East Asian crisis.

The next section discusses the exchange rate regimes in East Asia. Section III explains the empirical framework and methodology. Section IV discusses the empirical results, and Section V discusses the implications of the results and concludes.

II. Exchange Rate Regimes in East Asia

Since the financial crisis in 1997, exchange rate regimes have changed significantly in many East Asian economies. Appendix Table A reports the official IMF classifications of currency regimes for East Asian countries. The IMF classifications rely exclusively on each government's own declaration of its exchange rate regime. But, there are often contradictions between the exchange rate regime that prevails *de jure* and the way exchange rate policy is conducted *de facto*. For example, a regime that is classified as floating (independently or managed) might be in effect a peg with which the country defends its exchange rate within a narrow margin around a fixed rate. Recent studies such as Levy-Yeyati and Sturzenegger (2002) and Reinhart and Rogoff

(2004) highlight the contrast between countries' official declarations concerning their exchange rate regimes and the way they actually conduct exchange rate policy. The *de facto* classifications of exchange rate regimes for East Asian countries by Reinhart and Rogoff (2004) are reported in Appendix Table B.

Before the 1997 currency crisis, most East Asian economies opted to adopt *de facto* U.S. dollar peg systems for their exchange rate arrangements. The Hong Kong dollar was fixed to the dollar. The Chinese yuan was also been pegged to the U.S. dollar, but with occasional adjustments. The Thai baht and the Malaysian ringgit were similarly stable against the dollar, although these monetary authorities officially adopted a multiple currency basket system. Singapore, Korea, and the Philippines also targeted their currencies to the dollar rather loosely by combining discretion and market pressure with varying weights. Indonesia was on a *de facto* crawling peg to the U.S. dollar by sliding the rupiah by several percent per year to offset the inflation gap between home and abroad.

Most of the crisis-affected Asian economies have shifted their exchange rate regimes from *de facto* US dollar pegs to floating ones.[1] The dollar peg system they adopted before the crisis was thought to have contributed to the loss of confidence in their currencies in 1997. On July 2, 1997, Thailand adopted a managed floating exchange rate regime in which the value of the baht was determined by market forces and the Bank of Thailand would intervene in the market, only when it was necessary to avoid excessive volatilities. In July 1998, Thailand moved to substantially greater exchange rate flexibility. Indonesia, Korea, and the Philippines have also moved to floating exchange rate systems since the crisis. In contrast, Malaysia started pegging to the U.S. dollar in September 1998. Hong Kong and China have also kept their currencies pegged to the U.S. dollar.

Both the *de jure* and *de facto* definitions seem to have reflected these exchange regime changes in the post-crisis Asian countries. In fact, for the whole period of the 1990s, they render quite similar classifications of the exchange rate regime for each of the East Asian economies, at least in terms of three broad categories—fixed, intermediate and floating regimes. One exception is the Philippines. The IMF-based classification defines the country's exchange rate regime as floating since October 1984, while the *de facto* one categorizes it as intermediate before August 1995, fixed between September 1995 and June 1997, and intermediate from December 1997 until the present.

III. Empirical Test of Monetary Independence in East Asia

3.1. Empirical Specification of the Model

The basic estimation model can be specified as follows:[2]

$$\Delta r_t^{lc} = \alpha + \beta \Delta r_t^* + \rho \Delta r_{t-1}^{lc} + e_t, \qquad e_t \sim \ (0, \sigma_t^2)$$

Here r_t^{lc} represents the domestic nominal interest rate in the local currency of each country at time t, α is a constant term, r_t^* is the international interest rate, and β is the sensitivity of the local interest rate to foreign rates. This specification takes into account the dynamic adjustment of local interest rates to international interest rate

[1] See Baig (2001) for a detailed description of exchange rate behavior in East Asia after the crisis.
[2] The specification is derived easily from uncovered interest parity condition with open capital markets. See Shambaugh (2004) for the discussions of how omitted variables such as expected changes in the

shocks. The speed of adjustment towards the long-run equilibrium is defined by 1-ρ, and the long-run adjustment of local interest rates to international interest rate is measured by β/(1-ρ). Due to a failure to reject the null of a unit root for each of the interest rates under investigation we specify the model in the first-differenced interest rates. We test if the local and U.S. interest rates are cointegrated and cannot reject the null hypothesis of no cointegration.

We estimate the specification for each of the East Asian economies. By comparing the values of β across economies, we can assess whether the choice of the currency regime influences monetary policy independence across East Asian economies. According to the conventional view, more flexible exchange rate regimes allow countries additional room to pursue their independent monetary policies. Therefore, the sensitivity of local interest rates to international base rates should increase with the rigidity of the exchange rate regime. In other words, for a given degree of capital mobility, real integration and other factors, we would expect $\beta_{fixed} > \beta_{\text{int}\,ermediate} > \beta_{floating}$. In fact, in a fixed exchange rate regime with full capital mobility, β_{fixed} equals 1. In countries with similar exchange rate regimes, the sensitivity of the local interest rate to the foreign rate increases with the degree of capital account liberalization.

In the model, the parameter ρ captures the country heterogeneity in adjustment speed. The adjustment of domestic interest rate to the movement of the foreign base rate may not occur immediately (within a month). Differences in capital market openness and financial market development can cause heterogeneity in the adjustment dynamics among countries, even under the same exchange rate regimes. We expect that, all other

exchange rate and risk premium affect the estimate of β coefficient.

factors being equal, the adjustment speed would be higher (i.e., lower value of ρ) in fixed exchange rate regimes.

There occur some important issues in the estimation of the basic specification. Previous studies such as Frankel et al. (2002) assume that the changes in the U.S. interest rates are exogenous, and contemporaneously uncorrelated with the error term. This assumption can be justified in the sense that the financial markets of East Asian countries are relatively smaller than the U.S. financial market.

However, as Borensztein et al. (2002) point out, this assumption is unlikely to be true. Common shocks that affect both U.S. and domestic interest rates cause a potential endogeneity problem. Since East Asian economies are highly linked to the U.S. in trade, their business cycles tend to become synchronized. Hence, shocks to U.S. activities are likely to affect the outputs of Asian economies, leading to co-movements of U.S. and domestic interest rates. In general, any variables omitted from the specification, which are correlated with the U.S. interest rates, can cause biased estimates of the sensitivity parameter, β.

Another issue is that the sensitivity of local to international interest rates can also change over time within an economy where the exchange rate regime has changed during the sample period. Furthermore, the timing of actual structural breaks in the degree of monetary independence and the timing of exchange regime change may not coincide.

This paper adopts an estimation technique that accounts for the endogeneity problem. This technique also allows for the possibility of a structural break at an unknown time, and attempts to actually estimate the timing of the potential structural break in monetary autonomy.

3. 3 The Regime-Switching Model with Endogenous Explanatory Variables

For our empirical investigation of the issue, we employ a version of the regime-switching model with endogenous explanatory variables as introduced by Kim (1993). For simplicity of exposition, we focus on the estimation of the following benchmark model of ours in which the parameters undergo a structural break at an unknown break point and one of the regressors (Δr_t^*) is correlated with the disturbance term (v_t):

$$\Delta r_t = \alpha_{D_t} + \beta_{D_t}\Delta r_t^* + \rho_{D_t}\Delta r_{t-1} + v_t \tag{1}$$

$$\Delta r_t^* = z'_t d + \varepsilon_t \,, \tag{2}$$

$$\begin{bmatrix} \varepsilon_t \\ v_t \end{bmatrix} \sim i,i,d.N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon v,D_t} \\ \sigma_{\varepsilon v,D_t} & \sigma^2{}_{v,D_t} \end{pmatrix}\right), \tag{3}$$

where r_t is domestic interest rate and r_t^* is foreign interest rate. Equation (3) suggests that Δr_t^* and v_t are correlated. And z_t is a vector of instrumental variables that are correlated with Δr_t* but not with $e_{t.}$

The subscript D_t indicates that the parameters of interest undergo a structural break at an unknown break point τ, in the following way:

$$\theta_{D_t} = \theta_0(1-D_t) + \theta_1 D_t \qquad \theta = \alpha, \beta, \rho, \sigma_v^2 \text{, or } \sigma_{\varepsilon,v} \tag{4}$$

$$D_t = \begin{cases} 1, & if\ t \le \tau \\ 0, & otherwise \end{cases}$$

To complete the model, we specify the dynamics of the latent variable D_t as follows:

$$\Pr[D_t = 1 | D_{t-1} = 0] = 1 - q \quad , \quad \Pr[D_t = 0 | D_{t-1} = 1] = 0 \,, \tag{5}$$

which is designed to capture one-time permanent structural break in the sample at an unknown point in time. The estimate of the transition probability q provides an estimate of the break date, as the expected duration of a regime before a structural break is given by $\frac{1}{1-q}$.

In the absence of endogeneity in the explanatory variable, the above regime-switching model can be consistently estimated using the approach proposed by Hamilton (1989). In the above specification, however, Hamilton's approach is invalid as one of the regressors in (1) is correlated with the disturbance term. As Kim (2003) suggests, one way to solve the problem is to transform the model in (1) so that all the explanatory variables and disturbances are uncorrelated. To this end, we note that we can rewrite ε_t and v_t as functions of two independent standard normal random variables:

$$\varepsilon_t = \sigma_{\varepsilon,D_t} w_{1t}, \quad w_{1t} \sim i.i.d.N(0,1), \tag{6}$$

$$v_t = b_{21,D_t} w_{1t} + b_{22,D_t} w_{2t} = \frac{b_{21,D_t}}{\sigma_{\varepsilon,D_t}} \varepsilon_t + b_{22,D_t} w_{2t}, \quad w_{2t} \sim i.i.d.N(0,1), \tag{7}$$

where w_{1t} and w_{2t} are uncorrelated. From (2), (6) and (7), we rewrite the disturbance term in (1) in the following way:

$$v_t = \frac{b_{21,D_t}}{\sigma_{\varepsilon,D_t}} (\Delta r_t^* - z_t\,' d) + b_{22,D_t} w_{2t} \tag{8}$$

Then, by substituting (8) into equation (1), we get:

$$\Delta r_t = \alpha_{D_t} + \beta_{D_t} \Delta r_t^* + \rho_{D_t} \Delta r_{t-1} + \gamma_{D_t} (\Delta r_t^* - z_t\,' d) + w_t, \quad w_t \sim N(0, b_{22,D_t}^2), \tag{1'}$$

where $\gamma_D = \frac{b_{21,D_t}}{\sigma_{\varepsilon,D_t}}$.

Equation (1') is our transformed model in which all the explanatory variables are uncorrelated with the new disturbance term. As the transformed model does not suffer from endogeneity problem, we now can consistently estimate the parameters of our interest by employing the maximum likelihood estimation based on the Hamilton filter.

In addition, as the disturbance terms in (2) and (1') are uncorrelated, we can employ a simple two-step procedure described below:

<u>1 st Step</u> : [Estimation of the equation (2) using OLS]

Estimate equation (2): $\Delta r_t^* = z_t'\hat{d} + \hat{\varepsilon}_t$ and get the residuals $\hat{\varepsilon}_t$.

<u>2 nd Step</u> : [MLE founded on the Hamilton filter]

After replacing the $(\Delta r_t^* - z_t'\hat{d})$ with $\hat{\varepsilon}_t$, estimate the following equation using the MLE Proposed by Hamilton (1989):

$$\Delta r_t = \alpha_{D_t} + \beta_{D_t}\Delta r_t^* + \rho_{D_t}\Delta r_{t-1} + \gamma_{D_t}\hat{\varepsilon}_t +, w_t, \quad w_t \sim N(0, b_{22,D_t}^2), \qquad (1'')$$

Note that the above two-step procedure is further justified by the following decomposition of the log likelihood function:

$$\ln L(\alpha_0, \alpha_1, \beta_0, \beta_1, \rho_0, \rho_1, \gamma_0, \gamma_1, b_{22,0}, b_{22,1}, q)$$

$$= \sum_{t=1}^{T} \ln[\sum_{d_t=0}^{1} f(\Delta r_t \mid \Delta r_t^*, I_{t-1}, D_t = d_t]\Pr[D_t = d_t \mid I_{t-1}) + \sum_{t=1}^{T} \ln[f(\Delta r_t^* \mid I_{t-1})] \qquad (9)$$

3.4 Data

Our basic source of interest rate data is the International Monetary Fund (IMF), *International Financial Statistics*. We work with monthly data on 90-day local money market rates. As the international interest rate, we use the 90-day US T-bill rate. We focus on the sample period of January 1987 to April 2002. The vector of instrumental variables includes the four lags of domestic and U.S. interest rates (Δr_{t-1}^{lc}, Δr_{t-2}^{lc}, Δr_{t-3}^{lc}, Δr_{t-4}^{lc}, Δr_{t-1}^{*}, Δr_{t-2}^{*}, Δr_{t-3}^{*}, Δr_{t-4}^{*}).

4. Empirical Results

4.1. Testing for Structural Breaks

Before we apply the model introduced in Section 3, we employ a test of a structural break for each country. We then apply the regime-switching model to the sample of the countries for which the null hypothesis of no structural break is rejected. For the other countries for which no structural break is detected, we estimate the model without imposing any structural break during the sample period.

When the break date is unknown, the break date is a nuisance parameter that exists only under the alternative hypothesis but not under the null. The appearance of such nuisance parameter implies that the Wald tests of equality of the coefficients do not have the standard asymptotic properties. We employ the test developed by Andrews (1993) in order to deal with the problem of the nuisance parameter. The test is based on the following equation:

$$\Delta r_t = \alpha_{D_t} + \beta_{D_t} \Delta r_t^* + \rho_{D_t} \Delta r_{t-1} + \gamma_{D_t} \hat{\varepsilon}_t + w_t, \quad w_t \sim N(0, \sigma_t^2), \tag{10}$$

$$\hat{\varepsilon}_t = \Delta r_t^* - z_t'\hat{d} \tag{11}$$

$$D_t = \begin{cases} 1, & if\ t \le \tau \\ 0, & otherwise \end{cases} , \tag{12}$$

where τ is the break that needs to be estimated under the alternative hypothesis. Define the function $W_T(\tau)$ as the Wald statistic for the null hypothesis of no structural break in the coefficients, for each possible value of the break date τ. With T being the sample size, we assume that τ lies in a range $\tau_1 < \tau < \tau_2$, where $\tau_j = c_j T, \quad 0 < c_1 < c_2 < 1$. Andrews (1993) shows the asymptotic properties of the statistic $\sup_{\tau_1<\tau<\tau_2} W_T = \sup W_T(\tau)$ and reports the asymptotic critical values.

For each possible break date τ, we estimate equation (10) using the OLS methodology. We then calculate the Wald statistic for test of no structural break for given break date, $W_T(\tau)$, using the heteroscedasticity-consistent variance covariance matrix for the estimated coefficients. Estimated break date is the value of τ that maximizes $W_T(\tau)$, and the value of the maximized $W_T(\tau)$ is the $\sup W_T(\tau)$ test statistic.

The results of the Wald test are reported in Table 1. Among the eight East Asian economies under investigation, structural breaks were significant at a 5% significance level for the three crisis-hit countries: Korea, Malaysia, and Thailand. The test results show that structural breaks were not significant in the other five economies. They include Hong Kong, Japan, and Singapore that did not experience a currency crisis in addition to Indonesia and Philippines that experienced a currency crisis.

4.2. Estimation of the Model for Countries with and without Structural Break

Once structural break have been identified for Korea, Thailand, and Malaysia, we investigate the nature of structural breaks in the interest rate equations for these countries. In order to account for high uncertainty during the period of currency crisis, we apply a modified version of the model such as:

$$\Delta r_t = \alpha_{D_t} + \beta_{D_t}\Delta r_t^* + \rho_{D_t}\Delta r_{t-1} + \gamma_{D_t}\hat{\varepsilon}_t +, w_t, \quad w_t \sim N(0,\sigma_t^2), \qquad (9')$$

$$\sigma_t^2 = \begin{cases} \sigma_1^2, & if\ t < 1997:05 \\ \sigma_2^2, & if\ 1997:05 \le t \le 1998:10 \\ \sigma_3^2, & if\ t \ge 1998:10 \end{cases}$$

where the variance of the disturbance term is assumed to have different values for three sub-sample periods: the pre-crisis period from January 1987 to April 1997, the crisis period from May 1997 to September 1998, and the post-crisis period from October 1998 to April 2002.

Estimation results are reported in Table 2. For the three countries under investigation, structural break occurred during the crisis period: August 1997 for Malaysia, December 1997 for Korea, and August 1998 for Thailand.. For Korea and Thailand, there is a discernable difference in the coefficients describing the response of domestic interest rate to foreign interest rate before and after the structural breaks. For the period before the structural break, the estimated β coefficients, 2.95 (standard error=1.01) and 3.51 (s.e.=1.63) for Korea and Thailand respectively, are statistically significant. The estimates are large in magnitude but not statistically different from 1.

This confirms the theory: in the pre-crisis period when the countries adopted intermediate or *de facto* pegged regimes, domestic interest rates respond closely to foreign rates. By contrast, the estimates of β became smaller, close to zero, in magnitude and statistically insignificant in the post-crisis period when the economies chose to float. We also find that the speed of adjustment (1-ρ) became smaller with floating exchange rates after the structural break.

For Malaysia, the result shows that the estimated β coefficients indicating the response of domestic interest rate to foreign rate are statistically insignificant in both periods. That is, in contrast to the conventional theory, local interest rates did not become more sensitive to foreign rates in Malaysia after it changed its exchange rate regime from an intermediate regime in the pre-crisis period to a fixed one in the post-crisis period. This must be related to the tight capital controls adopted by the Malaysian government after the crisis. Even in pegged exchange rate regimes the monetary policy can be used independently by imposing capital control measures. Since capital controls permit domestic rates to react less to foreign interest rate shock, the estimates of β become statistically insignificant in the post-crisis period. Shambaugh (2004) show that in a large sample of countries, the response of local rates to foreign interest rates is smaller in pegs with capital controls than in pegs with open capital markets. Thus, the nature of structural break for Malaysia is somewhat different from that for Korea or Thailand. For Malaysia, structural break mainly occurred in the intercept term of the interest rate equation.

Note that most East Asian economies except Indonesia and Malaysia have accelerated the process of capital account liberalization over the 1990s (Kaminsky and Schmukler, 2001). An increase in the degree of capital account liberalization would have reduced monetary independence, if the same exchange rate arrangements had been maintained before and after the structural break. Thus, the significant increase in monetary independence in Korea and Thailand after the structural break must have come from the adoption of more floating regimes, rather than the change in capital account liberalization.

Table 3 presents the estimation results for the five countries—Hong Kong, Indonesia, Japan, Singapore, and the Philippines—which did not have any structural break during the sample period. For this group, our main interest is in examining whether the β parameters have different values across the economies that adopted different exchange regimes.

The results are broadly consistent with conventional predictions: interest rates in Hong Kong, which has maintained a credible fixed exchange regime, were very responsive to U.S. interest rates. The β estimate is positive and statistically significant (1.64, s.e.= 0.64). On the contrary, for Japan in which the floating regime has been adopted, the domestic interest rates did not respond to U.S. rates. The β estimate is 0.60 (s.e.= 2.88).

The estimates of the β coefficients turned out to vary a lot in intermediate regimes. In the Philippines, the local interest rates reacted closely to U.S. rates. The β estimate is statistically significant (1.73, s.e. =0.80). The Philippines has maintained *de factor* intermediate exchange rate regimes over the period except the period from September 1995 to June 1997 in which the country kept de factor fixed regime. For Singapore, the estimate of the response of local interest rates to U.S. interest rates was small in magnitude and not statistically significantly different from zero (0.12, s.e.= 0.40). For

Indonesia, which had adopted *de factor* intermediate exchange rate regimes before the crisis and then floated, the estimate of β was statistically insignificant too (0.84, s.e.=2.88).

IV. Conclusion

This paper examines whether the choice of an exchange rate arrangement has an important impact on monetary independence in East Asian countries, focusing on the possible presence of structural breaks before and after the 1997 Asian crisis. For our empirical investigation of the issue, we employ a regime-switching model with endogenous explanatory variables that allows for the endogeneity of foreign interest rate shocks and the possibility of a structural break in monetary autonomy at an unknown time.

We find that the sensitivity of local to U.S interest rates has declined in Korea and Thailand after they adopted floating exchange rate regimes. We also find that countries under more floating exchange rate regimes like Japan had greater independence in monetary policy than pegged economies such as Hong Kong. These empirical findings suggest that greater exchange rate flexibility offers a larger extent of monetary independence for East Asian economies.

Our methodology assesses monetary autonomy based on the observed degree of co-movements of local and US interest rates. Thus, we do not specify the actual operation of monetary policy. Monetary policy would respond to local conditions such as output gap and inflation too. In addition, the actual degree of monetary independence may differ from what is observed from the interest rates co-movements if the monetary authority’s exercise of its monetary autonomy is influenced by other factors such as business cycle synchronization, exposure to international capital markets, and political and institutional circumstances. In subsequent research, we plan to include additional variables such as domestic output gap, future expected inflation, and measures of linkages with the base country in the domestic interest rate equation. The application of our estimation technique to the extended model would be straightforward.

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Table 1. Tests of Structural Break

$$\Delta i_t = \alpha_{D_t} + \beta_{D_t} \Delta i_t^* + \gamma_{D_t} \widehat{w_t} + \rho_{D_t} \Delta i_{t-1} + e_t, \quad e_t \sim N(0, \sigma_{et}^2)$$
$$\Delta i_t^* = z_t ' \delta + w_t, \quad w_t \sim N(0, \sigma_{wt}^2)$$
$$D_t = 0 \quad for \quad t = 1, 2, \ldots, \tau; \quad D_t = 1 \quad for \quad t = \tau + 1, \ldots, T$$
$$\tau : break\ point$$

where the $\widehat{w_t}$ term is a bias correction term and z_t is a vector of instrumental variables.

	Sup-W statistic	Break date
Japan	7.87	--
Singapore	6.32	--
Korea	26.10*	1997:12
Thailand	25.51*	1998:08
Philippines	4.85	--
Malaysia	26.36*	1997:08
Indonesia	10.04	--
Hong Kong	9.91	--

Notes:
i) Test of structural break: Andrews' (1993) test with an unknown breakpoint.
ii) Sample period: 1987:01 - 2002:04 (Monthly data)
iii) Break dates were searched over the period 1993:01 – 1999:12.
iv) * significant at the 5% significance level

Table 2. Estimation of the Model for Countries with Significant Structural Break

$$\Delta i_t = \alpha_{D_t} + \beta_{D_t} \Delta i_t^* + \gamma_{D_t} \widehat{w_t} + \rho_{D_t} \Delta i_{t-1} + e_t, \quad e_t \sim N(0, \sigma_{et}^2)$$
$$\Delta i_t^* = z_t ' \delta + w_t, \quad w_t \sim N(0, \sigma_{wt}^2),$$
$$\Pr[D_t = 0 \mid D_{t-1} = 0] = q, \quad \Pr[D_t = 1 \mid D_{t-1} = 1] = 1$$
$$\frac{1}{1-q} : \textit{estimated break date}$$

			α	β	γ	ρ
	Before	Break	0.080	2.946*	-0.518	-0.062
			(0.109)	(1.018)	(1.248)	(0.052)
Korea	After	Break	-0.038	-0.290	0.568	0.752*
			(0.066)	(0.426)	(0.482)	(0.073)
	Break Date: 1997:08					
	Before	Break	-0.067	3.514*	-4.884*	-0.282*
			(0.173)	(1.633)	(1.954)	(0.090)
Thai	After	Break	-0.060	0.392	-0.521	0.379*
			(0.170)	(1.073)	(1.265)	(0.127)
	Break Date: 1998:08					
	Before	Break	0.059	0.152	-0.245	0.174
			(0.036)	(0.322)	(0.394)	(0.138)
Malaysia	After	Break	-0.171*	-0.427	0.879	-0.074
			(0.074)	(0.501)	(0.590)	(0.155)
	Break Date: 1997:10					

Notes: (i) Standard errors are in the parentheses.
(ii) * Significant at a 5% significance level.

Table 3. Estimation of the Model for Countries without Significant Structural Break

$$\Delta i_t = \alpha + \beta \Delta i_t^* + \gamma \widehat{w_t} + \rho \Delta i_{t-1} + e_t, \quad e_t \sim N(0, \sigma_{et}^2)$$
$$\Delta i_t^* = z_t{}'\delta + w_t, \quad w_t \sim N(0, \sigma_{wt}^2)$$

where the $\widehat{w_t}$ term is a bias correction term and z_t is a vector of instrumental variables.

	α	β	γ	ρ
Japan	0.017	0.595	0.839	0.177*
	(0.448)	(2.880)	(3.415)	(0.085)
Singapore	-0.018	0.118	0.545	0.102
	(0.047)	(0.395)	(0.529)	(0.062)
Philippines	0.067	1.731*	-1.777*	0.215*
	(0.095)	(0.795)	(0.855)	(0.100)
Indonesia	0.017	0.595	0.839	-0.177*
	(0.448)	(2.880)	(3.415)	(0.085)
Hong Kong	0.018	1.641*	-1.057	-0.326*
	(0.070)	(0.639)	(0.784)	(0.091)

Notes: (i) Standard errors are in the parentheses.
(ii) * Significant at a 5% significance level.

Appendix Table: Exchange Rate Regimes in East Asia

***A. De Jure* Regime (Official Classification by the IMF)**

Country	Period		Exchange rate regime classification	
	From	**To**	**Narrow**	**Broad**
Philippines	Oct81	Jun82	Limited flexibility wrt U.S. dollar	Intermediate
	Jul82	Sep84	Managed floating	
	Oct84	Jan02	Independently floating	Floating
Singapore	Jun73	Jun87	Limited flexibility wrt. a basket	Intermediate
	Jul87	Jan02	Managed floating	
Thailand	Jan77	Feb78	Peg to U.S. dollar	Fixed
	Mar78	Jun81	Limited flexibility wrt a basket	Intermediate
	Jul81	Mar82	Managed floating	
	Apr82	Oct84	Limited flexibility wrt U.S. dollar	
	Nov84	Jun97	Limited flexibility wrt a basket	
	Jul97	Jun98	Managed floating	
	Jul98	Jan02	Independently floating	Floating
Hong Kong	Jul72	Oct74	Peg to U.S. dollar	Fixed
	Nov74	Oct83	Independently floating	Floating
	Oct83	Jan02	Peg to U.S. dollar	Fixed
Indonesia	Nov78	Jul97	Managed floating	Intermediate
	Aug97	Jan02	Free floating	Floating
Japan	Dec71	Jan73	Peg to U.S. dollar	Fixed
	Feb73	Jan02	Independently Floating	Floating
Korea	Aug76	Jan80	Peg to U.S. dollar	Fixed
	Feb80	Nov97	Managed floating	Intermediate
	Dec97	Jan02	Independent floating	Floating
Malaysia	Sep75	Mar93	Limited flexibility wrt U.S. dollar	Intermediate
	Apr93	Aug98	Managed floating	
	Sep98	Jan02	Pegged to U.S dollar	Fixed
China	Mar81	Jul87	Pegged to a basket	Intermediate
	Aug87	Aug98	Managed floating	
	Sep98	Jan02	Limited flexibility wrt U.S. dollar	Fixed

Sources: Frankel, et al. (2002) and IMF *Annual Report on Exchange Arrangements and Exchange Restriction.*

Note: This classification of exchange rate regimes is based on a quarterly database from the IMF which encompasses a total of ten regime categories, based on officially reported exchange arrangements.

B. *De Facto* Regime (Reinhart and Rogoff, 2004)

Country	Period		Exchange rate regime classification	
	From	To	Narrow	Broad
Philippines	Dec72	Sep83	De facto crawling band around US	Intermediate
	Oct83	Feb85	Managed floating	
	Mar85	Apr92	De facto crawling peg to US dollar	
	May92	Aug95	De facto band around US dollar	
	Sep95	Jun97	De facto peg to US dollar	Fixed
	Jul97	Dec97	Freely floating/Free falling*	Floating
	Dec97	Dec01	Managed floating	Intermediate
Singapore	Jun72	Jun73	Peg to US dollar	Fixed
	Jun73	Nov98	De facto moving band around US$	Intermediate
	Dec98	Dec01	Managed Floating	
Thailand	Oct63	Mar78	Peg to US dollar	Fixed
	Mar78	Jul97	De facto peg to US dollar	
	Jul97	Jan98	Freely floating/Free falling*	Floating
	Jan98	Dec01	Manage floating	Intermediate
Hong Kong	Aug62	Jul72	Peg to pound sterling	Fixed
	Jul72	Oct83	De facto moving band around US $	Intermediate
	Oct83	Dec01	Currency board system/Peg to US$	Fixed
Indonesia	Dec70	Aug71	Peg to US dollar	Fixed
	Aug71	Oct78	De facto crawling band to US dollar	Intermediate
	Nov78	Jul97	De facto crawling peg to US dollar	
	Aug97	Jan02	Freely floating/Free falling*	Floating
	Apr99	Dec01	Freely floating	
Japan	Aug71	Dec71	Managed floating	Intermediate
	Dec71	Jan73	Bretton Woods Basket Peg	Fixed
	Feb73	Nov77	De facto moving band around US	Intermediate
	Dec77	Dec01	Independently Floating	Floating
Korea	May74	Feb80	Peg to US dollar	Fixed
	Feb80	Nov94	Pre announced crawling band	Intermediate
	Nov94	Nov97	De facto crawling peg to US dollar	
	Dec97	Jun98	Freely falling*	Floating
	Jul98	Dec01	Freely floating	
Malaysia	Jun67	Aug75	Peg to pound sterling	Fixed
	Sep75	Jul97	Limited flexibility wrt US dollar	Intermediate
	Aug97	Sep98	Freely floating/Free falling*	Floating
	Sep98	Dec01	Pegged arrangement	Fixed
China	Mar81	Jul92	Managed floating	Intermediate
	Aug92	Jan94	De facto crawling band around US$	
	Jan94	Dec01	De facto peg to US dollar	Fixed

Source: Reinhart and Rogoff (2004).
Note: Free falling is a new separate category for countries whose twelve – month rate of inflation is above 40%.