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FOREWORD

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ABSTRACT

The exchange rate peg to the United States dollar is widely believed to have been a major cause of the Asian financial crisis of 1997–1998. Rigid exchange rates may have invited massive capital inflows into East Asia by creating a false sense of security among investors. A substantial empirical literature examines the actual behavior of pre-crisis exchange rates in the region. This paper seeks to contribute to this literature by using daily data compared to other studies that tend to use monthly data and other lower-frequency data. The paper applies the GARCH-M model to investigate the statistical properties of East Asia's bilateral exchange rates vis-à-vis the United States dollar. Systematic relationships among the dollar exchange rates of the regional currencies are sought. The results confirm the rigidity of East Asia's pre-crisis bilateral dollar exchanges and support the existence of systematic relationships consistent with financial contagion. The findings lend some support to the region's post-crisis moves toward more flexible exchange rates.

I. INTRODUCTION

According to conventional wisdom, the rigidity of pre-1997 pegged exchange rate regimes in East Asia was one of the main causes of the currency crisis.[1] The basis for assuming that East Asia's pre-crisis exchange rates were rigid is the declaration of central banks about exchange rate policies. In practice, however, the declared exchange rate regime itself may not indicate rigidities—a tightly controlled managed float regime may be more rigid than a pegged regime with a wide band, *ceteris paribus*. Ghosh et al. (1995) and Reinhart (2000) point out that declared exchange rate regimes may differ from the actual characteristics of exchange rates. Therefore, it is worthwhile to pursue a statistical examination of East Asia's pre-crisis exchange rates in order to identify their distinctive characteristics, rather than take central bank declarations at full face value.

This paper investigates the issue of whether systematic relationships existed among East Asia's bilateral exchange rates vis-à-vis the United States (US) dollar in the pre-crisis period. For example, was there a systematic relationship between the Malaysian ringgit-US dollar exchange rates and the Thai baht-US dollar exchange rate? The existence of such systematic relationships in the pre-crisis period would lend further support to the contagion nature of the Asian currency crisis. At the same time, the pre-crisis existence of relationships raises the interesting question of how the peg regimes were able to last so long in light of their vulnerability to external shocks and contagion effects.

This paper investigates the behavior of the pre-crisis daily bilateral US dollar exchange rates of the four newly industrialized economies (NIEs) of Hong Kong, China; Republic of Korea; Singapore; and Taipei,China; and four members of the Association of Southeast Asian Nations (ASEAN-4), Indonesia, Malaysia, Philippines, and Thailand. Preliminary analyses and robust misspecification tests are performed to explore the distinctive characteristics of East Asian exchange rates. Empirical evidence of systematic relationships among the region's pre-crisis bilateral exchange rates vis-à-vis the US dollar are then established. Although the paper is essentially a study of exchange rate behavior rather than exchange rate policy or regime, it attempts to draw some policy implications from the results. However, the primary focus of the analysis remains exchange rate behavior rather than policy or regime.

II. A BRIEF OVERVIEW OF EAST ASIAN EXCHANGE RATE POLICIES

Conventional wisdom suggests that overvaluation of the East Asian currencies played a major role in precipitating the Asian currency crisis of 1997–1998 through its effect on the current account.[2] The twin roots of the overvaluation of those currencies were (i) those currencies' being closely pegged to the US dollar, and (ii) the sharp appreciation of the US dollar against the Japanese yen since

[1] See for example, International Monetary Fund (1998).

[2] See, for example, Grilli (2002), Glick and Rose (1999), and International Monetary Fund (1998).

April 1995. Since Japan was a major export market for the NIEs and the ASEAN-4, the consequent loss in the competitiveness of exports to Japan contributed significantly to the deterioration of their current account balances, which, in turn, contributed significantly to loss of confidence and outflows of capital from the region. The peg to the US dollar served the export-oriented East Asian economies well when the US dollar remained relatively weak against the other major currencies prior to its pre-crisis appreciation. However, the sudden strengthening of the US dollar caught those economies off guard, and turned an advantage into a disadvantage. The peg to the US dollar not only contributed to the crisis through current account channels but also through capital account channels. As Frankel (2003) points out, the peg and the consequent illusion of security from depreciation invited excessive capital inflows and moral hazard problems.[3] The abrupt and massive reversal of those excessive capital inflows was the immediate cause of the Asian crisis.

A well-known common structural characteristic of East Asian currencies in the pre-crisis period was their high degree of rigidity vis-à-vis the US dollar. The fact that the NIEs and the ASEAN-4 were small open economies highly dependent on exports for growth, combined with the well-known benefits of stable exchange rates for international trade, led those economies to more or less fix the US dollar price of their currencies. An additional perceived benefit of a fixed exchange rate system for a region that relied heavily on foreign capital was that it fostered a greater sense of confidence among foreign investors. In a seminal study, Frankel and Wei (1994) developed and popularized a method for uncovering the implicit weights assigned to major international currencies constituting a currency basket. Each weight picks up not only the direct impact of the major currency on an East Asian currency but also the indirect impact via the regional currencies. Applying this method to nine East Asian countries for the period from 1972 to mid-1992, Frankel and Wei found that the implicit weight of the US dollar in the currency basket ranged from 0.9 to 1. The only regional exception to the overwhelming dominance of the US dollar in the currency basket was the Singapore dollar, which was significantly influenced by the Japanese yen in addition to the US dollar.

Kwan (1995) updates the Frankel and Wei study to 1991–1995, and reconfirms the position of the US dollar as the dominant anchor for East Asian currencies, although he found that the implicit weight of the Japanese yen increased for the Korean won, Thai baht, Singapore dollar, and Malaysian ringgit. In an extension of the Frankel and Wei study, Chow et al. (2007) reconfirm the dominant position of the US dollar in East Asian countries' currency baskets in the pre-crisis period. The study explicitly includes regional competitors' currencies in the model in light of the export-oriented nature of East Asian economies and the real risk of competitive devaluation within the region. Furthermore, in order to overcome the simultaneity bias, the study replaces the regression model by a vector autoregressive model that allows for endogenous interactions among the exchange rate variables. An interesting additional finding for the pre-crisis period is that despite the peg, country-specific shocks also have a significant effect on East Asian currency fluctuations.

In light of the conventional wisdom suggesting that the de facto pegging of the regional currencies to the US dollar was a significant cause of the Asian crisis, many economists, including Fischer (2001) and Mishkin (1999), have called for more flexible exchange rates in the post-crisis period. The balance of evidence suggests that by and large East Asian exchange rates have indeed become more flexible in the post-crisis period.[4] More generally, whereas East Asian policy regimes were relatively homogeneous in the pre-crisis period—i.e., virtually all regional countries pegged

[3] Khan (2004) and Rajan, Siregar and Sugema (2003), among others, also express the same views.
[4] See, for example, Baig (2001).

their currencies to the US dollar—they have become far more diverse in the post-crisis period. The People's Republic of China and Malaysia adopted a fixed dollar exchange rate until July 2005 and Hong Kong, China continues to do so. On the other hand, Indonesia; Korea; Philippines; Thailand; and Taipei,China have chosen greater flexibility in their exchange rate management. Among the countries that have opted for greater flexibility, the implicit weight of the US dollar in the currency basket tends to be lower than in the pre-crisis period and the implicit weight of the Japanese yen tends to be higher.5 According to Kawai (2007), Korea; Thailand; and Taipei,China appear to have shifted to a managed float in which the weight of the US dollar and the Japanese yen is around 60–70% and 20–30%, respectively. The Philippine peso had shown greater volatility in the post-crisis period although much less so than the Indonesian rupiah, which has become more or less free floating.

The basic reason for this paper's focus on bilateral exchange rates vis-à-vis the US dollar and on the pre-crisis period is that, as noted, during this time the exchange rate regimes of the region were relatively homogeneous and based on a hard peg to the US dollar. Furthermore, the dollar peg is widely seen as having contributed to the crisis, which makes the examination of pre-crisis bilateral dollar exchange rates especially relevant for policymakers. This paper contributes to the existing empirical literature on pre-crisis exchange rates in East Asia by using daily exchange rate data, whereas other studies use lower-frequency data such as monthly data. Baillie and Bollerslev (1990) recommend the use of high-frequency data to investigate possible relationships among exchange rates. While fully realizing that policymakers' horizons extend well beyond one day and therefore caution must be exercised in interpreting the results, it is nevertheless true that daily data often contain rich and complex information that lower-frequency data fail to capture. For example, daily exchange rate volatility may be significantly higher than monthly exchange rate volatility and be subject to different dynamics. The choice of the GARCH-M model in this paper provides a neat and consistent way of modeling pre-crisis daily exchange rate behavior in the eight East Asian economies.

III. DATA CHARACTERISTICS

Evidence shows that intraday, daily, or weekly exchange rates are approximated by martingales, and the returns exhibit conditional volatility, fat-tail distribution, skewness, and excess kurtosis. Autoregressive conditional heteroscedasticity (ARCH) models have been used successfully to describe foreign exchange rate data. More recently, Engle and Gau (1997) examine the conditional volatility of the exchange rates under the target zone regime of the European Monetary System (EMS). They find that exchange rates within the EMS show the same characteristics as free-floating exchange rates except for strong negative autocorrelations.

Conditional volatility in exchange rates has been attributed to the irregular arrival of new information in the market. Engle, Ito, and Lin (1990) interpret volatility persistence in foreign exchange markets to either the time it takes market traders to act on new information or the autocorrelation of news across countries arising from potential policy coordination. They employ Ito and Roley's (1987) decomposition of the intraday yen–US dollar exchange rates based on the operating times of different markets to show volatility spillover across markets. On the other hand, Baillie and Bollerslev (1990) examine systematic relationships in the conditional returns or variances of the hourly bilateral exchange rates of Germany, Japan, Switzerland, and United Kingdom vis-à-vis

[5] See, for example, Kawai (2007) and MAS (2000).

the US dollar. They use the robust inference procedures of Wooldridge (1990) and Bollerslev and Wooldridge (1992) in order to handle the skewness and excessive kurtosis in the data. The results fail to show systematic relationships among the four major exchange rates.

The ASEAN-4 and NIE daily exchange rate data, quoted in US dollars, are from Datastream. The sample period is between 2 January 1991 and 1 July 1997, the day before the adoption of the de facto floating exchange rates by Thailand and the start of the Asian crisis.[6] Only bilateral exchange rates vis-à-vis the US dollar are considered, since the majority of East Asia's economic transactions are in US dollars. In addition, as noted earlier, according to Frankel and Wei (1994), the implicit weights of the US dollar in East Asian currencies range from 0.9 to 1.

Before the Asian currency crisis, the International Monetary Fund (IMF) classified East Asian exchange rate regimes based on the countries' description of their exchange rate policies. The regimes vary from fixed exchange rate regimes at one end, to the managed but more "free-floating" regimes at the other end. Hong Kong, China's currency was pegged to the US dollar at 7.8 to 1. Thailand's currency had an undisclosed band based on a basket of its main trading partners' currencies, while Indonesia and Korea had preannounced bands based on the previous day's rates. On the other hand, Malaysia; Philippines; Singapore; and Taipei,China had either managed or independently floating regimes. However, according to the IMF, all East Asian central banks intervened in the foreign exchange markets to limit short-term fluctuation in their exchange rates, using the US dollar as the main intervention currency (International Monetary Fund, various years).

Table 1 shows the results of the unit root tests and the summary statistics of exchange rate returns. The unit root hypothesis is tested using the Phillips-Perron (1988) test, which accounts for serial correlation and is robust to heteroscedasticity in the error terms. The Phillips-Perron (1988) test rejects the null of unit roots at the 5% and higher significance level in the bilateral exchange rates versus the US dollar of Indonesia and Hong Kong, China. In addition, the Phillips-Perron test also rejects the unit roots hypothesis for Thailand.[7] On the other hand, the Phillips-Perron (1988) test cannot reject unit roots for the other six East Asian currencies. Thus, in contrast to previous findings on high-frequency exchange rate data, the results give some support for trend stationarity in bilateral exchange rates, especially allowing for the possibility of heteroscedascity in the data.

[6] Before 1991, the exchange rates in many East Asian countries were based on government announced "official" rates rather than "freely-traded" exchange rates.

[7] The paper considers a 5% or higher level of significance.

TABLE 1
UNIT ROOT TESTS FOR BILATERAL RATES VERSUS THE US DOLLAR AND SUMMARY STATISTICS OF EXCHANGE RATE RETURNS ($\log(S_{i,t}/S_{i,t-1})$)

ECONOMY	UNIT ROOT TESTS[a] (PHILLIPS-PERRON Z_t–STATISTIC[b])	SUMMARY STATISTICS OF EXCHANGE RATE RETURNS					
		MEAN[c] (x10²)	STANDARD ERROR[c]	SKEWNESS[d]	KURTOSIS[d]	$Q_{i,5}$[d]	$Q^2_{i,5}$[d]
ASEAN-4							
Indonesia	−4.02**	0.459	0.183	2.775	42.702	34.440**	43.163**
Malaysia	−1.89	0.055	0.024	0.693	17.789	8.002	234.472**
Philippines	−2.35	0.193	0.080	0.602	61.155	3.165	19.601**
Thailand	−4.21**	0.189	0.078	0.708	146.857	609.089**	935.750**
NIEs							
Hong Kong, China	−4.24**	0.121	0.050	0.691	22.822	2.747	145.313**
Korea	−1.31	0.400	0.160	1.247	75.565	3.065	6.133
Singapore	−1.17	0.021	0.014	-0.204	5.764	23.992**	54.982**
Taipei,China	−3.21	0.196	0.080	-0.038	5.245	11.945**	322.386**

Note: RATS 4.3 is used for the summary statistics and the various statistical tests.
[a] The null is unit root for the Phillips-Perron (1988) tests and the alternative hypothesis is trend stationarity. The Phillips-Perron test is also robust to possible heteroscedasticity and serial correlation.
[b] The critical values are −3.66 and −3.96 for Phillips-Perron Z_t-statistics at the 5%(*) and 1% (**) significance levels. See Hamilton (1994).
[c] F-tests for variances of the bilateral exchange rates reject the null that the variances are equal for all the economies except Philippines; Taipei,China; and Thailand,at the 5% significance level.
[d] As in the Baillie and Bollerslev (1990), the standardized residuals ($\bar{\varepsilon}_{i,t} = (y_{i,t} - \mu_{i,t}) / \sqrt{h_{i,t}}$) in the preliminary analysis are calculated from an MA(1) model. The skewness, kurtosis, and Ljung Box Q-statistics are based on the standardized residuals. $Q_{i,5}$ and $Q^2_{i,5}$ are the Ljung-Box (1978) Q-statistic for five serial correlation of the standardized residuals and squared residuals respectively.

Table 1 also shows the summary statistics of exchange rate returns. Despite the finding of trend stationarity in some of the series, the standard practice of examining the relationships in the returns is followed, rather than log levels in all exchange rate data. Unit root tests have low power when the alternative is close to but still larger than 1. Furthermore, the theoretical justification for the relationship between exchange rates in returns and log levels is inadequate.[8]

In addition, Table 1 presents the estimates of the skewness, kurtosis, and Ljung-Box (1978) Q and Q-squared statistics for the standardized residuals.[9] The sample estimates have significantly nonzero skewness and excess kurtosis that indicate non-normal distribution. Furthermore, the Ljung-Box Q and Q-squared statistics are large for all the series. As Baillie and Bollerslev (1990) point out, despite the presence of heteroscedasticity and excess kurtosis, large Q and Q-squared estimates may be an indication of serial correlation and conditional heteroscedasticity, respectively. The data characteristics of eight East Asian exchange rates are not consistent with the previous findings for major currencies in three of the eight cases. Hong Kong, China; Indonesia; and Thailand exchange rates show evidence of trend stationarity rather than unit roots, although preliminary analysis of characteristics of exchange rate returns also show non-normality.

[8] Papell (1992) and then Alba (1997) make the same argument, but for first differencing all of the series, despite the finding of a second root in some of the data.
[9] The standardized residuals ($\bar{\varepsilon}_{i,t} = (y_{i,t} - \mu_{i,t}) / \sqrt{h_{i,t}}$) in the preliminary analysis are calculated from an MA(1) model (Baillie and Bollerslev 1990).

IV. MODEL SPECIFICATION

This paper adopts as its general model Engle and Gau's (1997) "official band" model, which is an extension of the GARCH-M model. The model specification includes the impact of volatility on returns, and a location parameter that ensures consistency of non-Gaussian Quasi-Maximum-Likelihood Estimators (QMLE):

$$y_{i,t} = \alpha_{i,0} + \alpha_{i,1}h_{i,t} + \varepsilon_{i,t} + \alpha_{i,2}\varepsilon_{i,t-1} + \alpha_{i,3}\sqrt{h_{i,t}} \quad \text{and}$$

$$h_{i,t} = \beta_{i,0} + \beta_{i,1}h_{i,t-1} + \beta_{i,2}\varepsilon^2_{i,t-1} + \beta_{i,3}x_{i,t-1} \tag{1}$$

where $y_{i,t}$ is exchange rate return given by $y_{i,t}$ = 100 log($S_{i,t}$ / $S_{i,t-1}$), $S_{i,t}$ is the daily bilateral exchange rate of country i, and $h_{i,t}$ is the conditional variance while $\varepsilon_{i,t}$ is the residual.

The serial correlation in the data is represented by MA(1), while the GARCH(1,1)-M incorporates conditional heteroscedasticity and the relationship between returns and volatility of exchange rates (Engle, Lilien, and Robins 1987). The last term in the returns equation implements the addition of a location parameter suggested by Newey and Steigerwald (1997), when the conditional mean is identically nonzero or the symmetric condition does not exist.[10] The additional parameter, even with the existence of asymmetry, meets the identification condition necessary for the consistency of non-Gaussian QMLE.

The term x_{it}, which is equal to |100 log($S_{i,t}$ / $T_{i,t}$)|, specifies the band in the conditional variance equation, where $T_{i,t}$ is the "target" exchange rate. Therefore, the coefficient $\beta_{i,3}$ shows the impact of previous deviation of the exchange rate from its target rate on the magnitude of the volatility. A negative $\beta_{i,3}$ implies that the band holds, while a positive value means the band does not hold. In order to capture the intervention policies of East Asian central banks under managed-float regimes or undeclared-band regimes with interventions based on the previous day's exchange rate, the proxy for the "band" $x_{i,t-1}$=|100 log($S_{i,t-1}$ / $S_{i,t-2}$)| is used.[11] However, for countries with declared bands based on the previous day's exchange rate, the proxy $x_{i,t}$=|100 log($S_{i,t}$ / $T_{i,t}$)|−w_it is used, where w_{it} represents the bandwidth for the i^{th} currency at time t.[12] In addition, for Hong Kong, China, which has a one-sided band, the proxy $x_{i,t}$=|100 (log7.8−log $S_{i,t}$) is used. Unlike the EMS, East Asian central banks intervene unilaterally rather than in a coordinated manner to maintain the bands. It should be noted that the only policy-related term in the GARCH model is the band. As such, much caution should be taken in interpreting the results for countries with undeclared bands.

The specification of the univariate model employs the robust Wald test developed by Bollerslev and Wooldridge (1992). The test statistic only requires first derivatives of the conditional mean and variance functions with respect to the parameters of the model, and a specification of a constraint vector. Under the null hypothesis that the restrictions hold, the robust Wald statistic is asymptotically

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[10] The location parameter becomes relevant only if the εt is non-Gaussian and the conditional mean and conditional variance are not appropriately specified (E(εt|yt-1)ε0 and E(εt2|yt-1)ε1).

[11] Since the US dollar is the common intervention currency, the band specification only considers the bilateral rate versus the US dollar.

[12] Only Indonesia and Korea have specified bands. The IMF's Annual Report on Exchange Rate Arrangements and Exchange Restrictions (International Monetary Fund, various years) reports the specified bands.

a chi-square distribution, with degrees of freedom equal to the number of restrictions. The robust Wald statistic is valid regardless of the normality assumption.[13]

Initially, the no-GARCH(1,1) specification is tested against the GARCH(1,1) alternative hypothesis. If GARCH(1,1) is not rejected, then the alternative hypothesis of MA(1)-GARCH(1,1)-M-with-a-band specification in equation (1) is tested against various restricted null specifications: the MA(1)-GARCH(1,1)-M (β_3=0), MA(1)-GARCH(1,1) (α_1=β_3=0), GARCH(1,1)-M (α_2=β_3=0), and GARCH(1,1) (α_1=α_2=β_3=0) without the band; and the corresponding restrictions of models with the band, α_1=0, α_2=0, and α_1=α_2=0, respectively.

Table 2 reports the results of the robust-Wald specification tests. The most parsimonious specification of each series determines the choice of univariate model. The no-GARCH(1,1) restriction against the GARCH(1,1) alternative cannot be rejected at the 5 percent significance for Indonesia, Korea, and Philippines. Additional robust Wald tests for the remaining five cases indicate GARCH(1,1) to be the appropriate specification for the bilateral exchange rates of Malaysia and Singapore, while MA(1)-GARCH(1,1) characterize the Hong Kong, China and Taipei,China dollar vis-à-vis the US dollar. The GARCH(1,1)-with-band specification is suitable only for the Thai baht.

TABLE 2
ROBUST WALD TESTS[a] FOR MODELS OF BILATERAL RATES VERSUS US DOLLARS

$$y_{i,t} = 100\log(S_{i,t}/S_{i,t-1})\,;\; y_{i,t} = \alpha_{i,0} + \alpha_{i,1}h_{i,t} + \varepsilon_{i,t} + \alpha_{i,2}\varepsilon_{i,t-1} + \alpha_{i,3}\sqrt{h_t}\,;$$
$$h_{i,t} = \beta_{i,0} + \beta_{i,1}h_{i,t-1} + \beta_{i,2}\varepsilon^2_{i,t-1} + \beta_{i,3}x_{i,t-1}\,;\; x_{i,t} = \left|100\log(S_{i,t}/T_{i,t})\right|$$

NULL HYPOTHESIS	IN[c]	ML[d]	PH[c]	TH[f]	HK[e]	KR[c]	SG[d]	TW[e]
β_1=β_2=0 [b]	0.188	94.767**	0.108	7.027*	156.123**	0.429	120.890**	167.527**
α_1=α_2= β_3=0	–	0.281	–	7.128+	16.100**	–	2.110	290.528**
α_1=α_2=0	–	0.021	–	0.008	15.676**	–	0.992	290.059**
α_1=β_3=0	–	0.277	–	7.111*	0.086	–	2.110	0.722
α_2= β_3=0	–	0.276	–	7.119*	16.046**	–	1.101	288.795**
α_1=0	–	0.000	–	0.001	0.071	–	0.933	0.301
α_2= 0	–	0.021	–	0.007	15.652**	–	0.046	288.388**
β_3=0	–	0.271	–	7.105*	0.024	–	1.100	0.444

IN = Indonesia; ML = Malaysia; PH = Philippines; TH = Thailand; HK = Hong Kong, China; KR = Korea; SG = Singapore; TW = Taipei,China.

[a] The misspecification test initially considers the no-GARCH as the null hypothesis against the alternative of GARCH(1,1). If GARCH(1,1) is not rejected, then additional misspecification tests under various null hypotheses are conducted against the general model (MA(1)-GARCH(1,1)-M-with-a-band) as the alternative. The most parsimonious specification is chosen for each series. The computer programs are in RATS 4.3.

[b] To test for no-GARCH(1,1), the robust Wald test has a null hypothesis of no GARCH(1,1) against GARCH(1,1) as an alternative. (**), (*) and (+) indicate 1%, 5%, and 10% levels of significance, respectively.

[c] Exchange rates of Indonesia, Korea, and Philippines do not reject the no-GARCH(1,1) specification against the alternative of GARCH(1,1) at the 5% level of significance.

[d] Additional robust Wald tests indicate the nonrejection of the null against the general model for Malaysia and Singapore, therefore the GARCH(1,1) model is chosen.

[e] The rejection of the null α_1=α_2= β_3=0, α_1=α_2=0, α_2= β_3=0, α_2= 0 but the nonrejection of α_1=β_3=0, α_1=0, β_3=0 at the 5% level of significance for Hong Kong, China and Taipei,China denote an MA(1)-GARCH(1,1) model.

[f] The rejection of the null α_1=α_2=0, α_1=β_3=0, and β_3=0 at the 5% level and α_1=α_2= β_3=0 at the 10% level of significance; and the nonrejection of α_1=α_2=0, α_1=0, α_2=0 at the 10% and higher level of significance denote a GARCH(1,1)-with-a-band model.

[13] Baillie and Bollerslev (1990) highlight the importance of robust inference procedures in high-frequency data considering the typical findings of non-normality, skewness, and excess kurtosis.

Table 3 shows the QMLE estimates of the various models and their corresponding asymptotic robust standard errors. The calculation of the parameters uses the SIMPLEX algorithm to arrive at initial estimates and the Berndt et al. (1974) algorithm for the reported estimates, at a critical value of 10^{-4}. The computation of the asymptotic robust standard errors, as in Bollerslev and Wooldridge (1992), is from the first derivatives of the conditional mean and variance functions. QMLE is valid under non-normality provided the standardized-residual means and variances are 0 and 1, respectively (see Hamilton 1994, 663).

Table 3
Quasi-maximum Likelihood Estimates for Bilateral Rates versus US dollar

$$y_{i,t} = \alpha_{i,0} + \alpha_{i,1}h_t + \alpha_{i,2}\varepsilon_{i,t-1} + \alpha_{i,3}\sqrt{h_t}\ ;\ h_{i,t} = \beta_{i,0} + \beta_{i,1}h_{i,t-1} + \beta_{i,2}\varepsilon^2_{i,t-1} + \beta_{i,3}x_{i,t-1}\ ;\ x_{i,t} = \left|100\log(S_{i,t}/T_{i,t})\right|$$

Parameter	Indonesia	Malaysia	Philippines	Thailand	Hong Kong, China	Korea	Singapore	Taipei,China
α_0	0.0150	−0.0044	−0.0018	−0.0038	0.0004	0.0130	−0.0170	0.0047
	(0.0034)	(0.0043)	(0.0111)	(0.0032)	(0.0008)	(0.0068)	(0.0044)	(0.0116)
α_2	–	–	–	–	−0.1860	–	–	−0.5486
					(0.0568)			(0.0324)
β_0	0.0211	0.0106	0.2068	0.0124	0.0000	0.0832	0.0084	0.0503
	(0.0031)	(0.0025)	(0.0400)	(0.0024)	(0.0001)	(0.0171)	(0.0021)	(0.0110)
β_1	–	0.4477	–	0.3153	0.5900	–	0.5609	0.5557
		(0.0714)		(0.1056)	(0.0615)		(0.0720)	(0.0664)
β_2	–	0.3852	–	0.6743	0.6768	–	0.2841	0.3379
		(0.0768)		(0.0624)	(0.1739)		(0.0727)	(0.0685)
β_3	–	–	–	−0.1599	–	–	–	–
				(0.0184)				
$\mu(\varepsilon)$	−0.0001	−0.0024	0.0002	0.0340	−0.0284	−0.0006	0.0324	−0.0063
$\sigma^2(\varepsilon)$	0.9954	1.0006	1.0010	0.9995	1.0010	0.9800	0.9996	1.0005
$m_{i,3}$	2.7751	−0.4286	0.6019	−0.7434	−1.4436	1.2476	−0.1990	−0.1928
$m_{i,4}$	42.7014	6.2240	61.1550	14.1411	18.7482	75.5653	4.4438	6.6921
$Q_{i,5}$	35.3398	5.8914	3.9394	26.4084	6.0606	3.5827	15.3187	2.1687
$Q^2_{i,5}$	–	1.4270	–	31.8658	1.4287	–	2.3260	1.8758
LogL	2428.807	1881.463	486.661	2529.174	5009.296	1294.49	1863.969	219.102

Note: The asymptotic robust standard errors are in parenthesis. The critical value for a two-tailed t-test is 1.96 at the 5% significance level. $\mu(\varepsilon)$ and $\sigma^2(\varepsilon)$ are the standardized-residual mean and variance, respectively. $m_{i,3}$ and $m_{i,4}$ refer to the skewness and kurtosis, while $Qi_{,5}$ and $Q^2_{i,5}$ are the Ljung-Box Q-statistics for the standardized residuals and residuals square, respectively. $Qi_{,5}$ is significant only for Indonesia, Singapore, and Thailand, while $Q^2_{i,5}$ is significant only for Thailand at the 5% level. α_1 and α_3 are omitted since none of the models shows GARCH-in-mean and asymmetry ($\sigma^2(\varepsilon)\neq 1$), respectively. Also, since the variances of Indonesia, Korea, and Philippines are constant, $Q^2_{i,5}$ is omitted. The computer programs are written in RATS 4.3.

The coefficient estimates, as shown in Table 3, are significant in 23 out of 29 cases. Although Thailand had an undeclared band in practice, the coefficient of the assumed "band" (β_3) is negative and significant. Thus, using the "proxy" band, the pre-crisis undeclared exchange rate band regime of Thailand seems to be valid. The Ljung-Box Q and Q-squared statistics are insignificant in 10 out of 16 cases. However, the nonzero skewness and excess kurtosis for all of the series still indicate non-normality, thereby showing the importance of using robust inference procedures (Baillie and Bollerslev 1990).

V. SYSTEMATIC RELATIONSHIPS

The causality tests developed by Cheung and Ng (1996) are used in order to examine the existence of systematic relationships among East Asian exchange rates during the pre-crisis period.[14] The tests determine the causality in variance, in the Granger (1969) sense, of two series using a two-step procedure. The first step requires the identification of the appropriate model for the univariate time series, and the second step utilizes a cross-correlation function (CCF) of the standardized squared residuals to check for causality in variance. The tests are robust to distributional assumptions.

The standardized squared residuals of the models in Section III are defined as:

$$\bar{\varepsilon}^2_{i,t} = (y_{i,t} - \mu_{i,t})^2 / h_{i,t} \tag{2}$$

where $\mu_{I,t}$ is the mean of the exchange rate return for country *i*. The CCF of the sample is given by:

$$\hat{\rho}_k(\bar{\varepsilon}^2_i, \bar{\varepsilon}^2_j) = \frac{\sum(\bar{\varepsilon}^2_{j,t} - E(\bar{\varepsilon}^2_{j,t}))(\bar{\varepsilon}^2_{i,t-k} - E(\bar{\varepsilon}^2_{i,t}))}{\sqrt{\sum(\bar{\varepsilon}^2_{j,t} - E(\bar{\varepsilon}^2_{j,t}))^2 \sum(\bar{\varepsilon}^2_{i,t} - E(\bar{\varepsilon}^2_{i,t}))^2}} \tag{3}$$

where the subscripts j, I, and k refer to the first series, the second listed country, and the lead (k<0) or lag (k>0), respectively.

Cheung and Ng (1996) show that the square root of the number of observations multiplied by the sample cross-correlation has an asymptotic normal distribution. Therefore, the statistic ($\sqrt{T}\hat{\rho}_k(\bar{\varepsilon}^2_i, \bar{\varepsilon}^2_j)$) may be used to detect causality in variance. Given $\hat{\rho}_k(\bar{\varepsilon}^2_i, \bar{\varepsilon}^2_j) \neq 0$, if k>0, then the second series lags or "Granger causes" the first series; but if k<0, then the first series lags or "Granger causes" the second series. *k*=0 indicates instantaneous causality. Bidirectional causality may exist if the statistics are significant for both k>0 and k<0. The maximum value of leads and lags is 10.

The cross-correlation of the standardized residuals also shows causality in mean. However, Cheung and Ng (1996) cautioned that for some model specifications, the presence of both the causality-in-variance and causality-in-mean may affect the CCF because of the violation of the independence assumption in either test. In a GARCH(1,1) model, the causality in mean may have a large effect on the size of the causality in variance test since the conditional variance is a function of squared errors, although the reverse may not be true.

Table 4 shows the significant cross-correlation leads and lags of the standardized residuals and residual squares of the bilateral rates versus the US dollar. The results indicate systematic relationships among the exchange rates, but mainly in mean rather than in variance. Since only the causality in mean exists in most cases, and the specified model is either no-GARCH or GARCH(1,1), then the CCF is not likely to be biased. The results are also consistent with the observed rigidity of the East Asian exchange rates vis-à-vis the US dollar.

[14] Section II indicates all the series to be non-Gaussian so the "causality tests" cannot be used within the vector error correction model. Harris (1995) specifies that the residuals in a vector error correction model should be Gaussian.

Table 4
Significant Cross-Correlation Leads and Lags of Standardized Residuals in Levels (Upper Diagonal) and Squares (Lower Diagonal)
(Bilateral Rates vs US dollar)

Second Variable	First Variable IN	ML	PH	TH	HK	KR	SG	TW
IN		−5, 0	−5, 0, 2, 3, 5	−5, 0, 5	−5, −2, 5	−5, 0, 5	−5, 0	−2
ML	-		0, 3	0, 5	−5, 0	0, 1, 4,5	−1, 0, 1	-
PH	-	-		−5, −4, −3, −2, 0, 1, 2, 3, 5	-	−5, 0, 1	−5, −3, 0	2, 4, 5
TH	-	−4, 0 1	-		-	−5, 0, 5	−5, 0	−4
HK	-	5	-	−4		1, 4	−1, 0, 2, 5	−4, −2
KR	-	-	-	-	-		−3, 0	-
SG	-	−1, 0, 1	-	0	−5	-		−2, −1, 5
TW	-	-	-	-	−3,	-	−5, 1, 2	

IN = Indonesia; ML = Malaysia; PH = Philippines; TH = Thailand; HK = Hong Kong, China; KR = Korea; SG = Singapore; TW = Taipei,China.

Note: "Significant" refers to significance level of 5% or higher. The cross-correlation leads and lags of standardized residuals (causality in mean) are shown on the upper diagonal while the lower diagonal shows the significant lead or lags of the standardized residual squares (causality in variance). A positive value means that the second variable lags the first variable while a negative value connotes the first variable lags the second variable. Zero indicates significant instantaneous correlation. The computer programs are written in RATS 4.3.

Among the ASEAN-4, bidirectional causality in mean exists between Thailand and the Philippines, the first two countries involved in the contagion. There is also bidirectional causality in mean between Indonesia and both the Philippines and Thailand. The direction of causality in mean for Malaysia and the other ASEAN-4 countries is from Malaysia to Indonesia, Philippines, and Thailand, although causality in variance is bidirectional between Malaysia and Thailand. Table 4 also indicates the spillover to Korea, which shows bidirectional causality in mean with Indonesia, Philippines, and Thailand. There is also unidirectional causality in mean from Malaysia to Korea. For Hong Kong, China; Singapore; and Taipei,China, the causality is mainly from them to the ASEAN-4. Exceptions include bidirectional causality in both mean and variance between Malaysia and Singapore, and unidirectional causality in mean from the Philippines to Taipei,China.

For both mean and variance, systematic relationships also exist between Hong Kong, China and both Singapore and Taipei,China. The causality is bidirectional for mean and variance between Singapore and Taipei,China. Instantaneous causality in mean exists for most of the economies except Korea and Taipei,China on one hand and the other seven East Asian economies on the other. There is also no instantaneous causality in mean between Hong Kong, China and Indonesia. Instantaneous causality in variance is also present between Malaysia, and both Thailand and Singapore, as well as between Singapore and Thailand.

Therefore, the results indicate the presence of systematic relationships among East Asian bilateral exchange rates vis-à-vis the US dollar in the pre-crisis period, which are consistent with the contagious nature of the Asian currency crisis. Such systematic relationships can help explain how the currency crisis, which originated in Thailand, quickly spread to other Southeast Asian countries as well as Korea, resulting in a regionwide crisis.

Extensive trade links among the eight East Asian economies can help explain the pre-crisis systematic relationships among exchange rates. The magnitudes of those links are shown in Table 5. Although institutional economic integration along the lines of the European Union or North Atlantic Free Trade area has been slow to take hold in East Asia, de facto integration has reached substantial levels, especially in trade. Intraregional trade is now a major engine of growth for both the NIEs and ASEAN-4. To a lesser extent, investment flows are also contributing toward closer intraregional economic linkages. The NIEs and Malaysia have all emerged as significant sources of foreign direct investment in East Asia, as Table 6 shows.

Table 5
Trade among the NIEs and ASEAN-4, 1996 (millions of US$)

	Korea	Taipei,China	Hong Kong, China	Singapore	Malaysia	Thailand	Indonesia	Philippines
Korea	–	4,014	11,191	6,460	4,343	2,671	3,188	1,923
Taipei,China	2,655	–	26,718	4,562	2,946	2,782	1,950	1,926
Hong Kong, China	2,935	4,311	–	4,964	1,694	1,809	1,006	2,155
Singapore	4,715	4,872	10,208	–	22,511	7,096	2,875	2,297
Malaysia	2,386	3,212	4,607	16,018	–	3,207	1,219	939
Thailand	1,013	1,421	3,240	6,749	2,014	–	846	631
Indonesia	3,281	1,609	1,625	4,565	1,110	823	–	688
Philippines	371	661	868	1,224	687	780	90	–

Note: The economy on the vertical axis represents exporting country. For example, Korea exported US$4,014 million to Taipei,China and Taipei,China exported US$2,655 million to Korea.
Source: Australian Department of Foreign Affairs and Trade (2000).

Table 6
Foreign Direct Investment Flows among the NIEs and ASEAN-4, 1996 (millions of US$)

	Korea	Taipei,China	Hong Kong, China	Singapore	Malaysia	Thailand	Indonesia	Philippines
Korea	–	5	325	19	258	85	211	29
Taipei,China	2	–	303	95	310	254	167	47
Hong Kong, China	229	267	–	471	70	148	525	76
Singapore	47	86	769	–	1,141	169	505	136
Malaysia	673	204	328	845	–	100	165	151

Note: The economy on the vertical axis indicates investing country. For example, Korea invested in US$5 million in Taipei,China while Taipei,China invested US$2 million in Korea. Thailand, Indonesia, and the Philippines are not significant sources of foreign direct investment.
Sources: US State Department (2007).

Economic linkages are more limited among Indonesia, Philippines, and Thailand. However, the three countries compete with each other in export markets and in attracting foreign investment. Thus, devaluation of the Thai baht adversely affects the export sector in the Philippines and Indonesia. In addition, post-crisis analysis by Corsetti et al. (1998) point to common structural problems in East Asian countries just prior to the crisis. These include sizable current account deficits, unsustainable lending booms, and sharp increases in bad loans. Therefore, in principle, foreign investors' loss of confidence in Thailand can easily lead to loss of confidence in other East Asian countries.

VI. CONCLUSIONS AND POLICY IMPLICATIONS

Some distinctive characteristics of pre-crisis daily bilateral East Asian exchange rates vis-à-vis US dollar have been seen. In contrast with previous findings on major currencies, including exchange rates within the EMS, the bilateral exchange rates of Hong Kong, China; Indonesia; and Thailand show trend stationarity rather than unit roots. The Hong Kong dollar is pegged to the US dollar, while the rupiah and baht are underdeclared and undeclared bands, respectively. Furthermore, using robust inference procedures, returns to the Indonesian rupiah, Philippine peso, and Korean won vis-à-vis the US dollar are found to not exhibit volatility clustering.

Trend stationarity and lack of volatility clustering in pre-crisis daily exchange rates may be due to central bank policies. Hong Kong, China has an exchange rate mechanism to maintain the peg, backed by a large amount of foreign currency reserves. In the Philippines, commercial banks and the central bank suspend daily foreign exchange trading for up to two hours whenever the exchange rate deviates certain percentage points from the previous day's closing rate.[15] Furthermore, controls imposed on capital flows via licensing requirements may also limit volatility. Capital flow restrictions in Korea (see Black 1996) and Indonesia (see Nasution 1998, 270–2] may help explain the nonexistence of volatility clustering in the won–US dollar and the rupiah–US dollar rates.[16] Similarly, Thailand also imposed capital inflow restrictions from 1995 to 1997 (see Ariyoshi et al. 2000).

Systematic relationships are also found to exist among the pre-crisis East Asian exchange rates. Those relationships are consistent with the contagious nature of the crisis. In line with the preliminary findings of rigidities in the bilateral rates of East Asian countries, the systematic relationships show causality primarily in mean returns. This is especially true for Southeast Asian currencies. Extensive relationships in mean exist between Thailand and the Philippines, Philippines and Indonesia, and to a lesser degree between Thailand and Malaysia. In addition, Thailand and Malaysia affect each other in variance. The spillover from the ASEAN-4 to the NIEs, mainly in mean returns, is from the ASEAN-4 currencies to the Korean won, the Malaysian ringgit to Singapore dollar, and Philippine peso to the Taipei,China dollar, vis-à-vis the US dollar. Also observed are systematic relationships among the NIEs. It should be noted that such systematic relationships could simply reflect relationships among foreign exchange market conditions rather than relationships among exchange rates.

Despite the presence of systematic relationships and the vulnerability to external shocks they imply, East Asian central banks were able to maintain peg regimes mainly through capital controls.[17] However, starting in the mid-1990s, East Asian countries gradually liberalized their capital accounts, while still maintaining the peg regimes. Foreign investors seem to have ignored exchange rate risks

[15] Although the mechanism was discontinued after 1 March 1996, it partially explains the lack of volatility clustering in the daily peso-US dollar rate for most of the 1990s.
[16] Korea did pursue gradual capital account liberalization since 1993.
[17] The debate on capital controls focuses on its role in crisis management (see Corsetti et al. 1998), rather than its role in maintaining the peg regimes of East Asia.

(International Monetary Fund 1998) even after macroeconomic imbalances (Corsetti et al. 1998) emerged.[18] In addition, the capital account liberalization occurred under a weak prudential regime (Ariyoshi et al. 2000) in the financial sector. After capital account liberalization in East Asia, central bank intervention aimed at maintaining the peg regimes proved costly for East Asian countries.

The empirical findings provide some support for the notion that peg regimes contributed to East Asia's currency crisis. Not only did the rigidity of exchange rates help to attract unsustainably large capital inflows into the region and contribute to the development of macroeconomic imbalances such as sizable current account deficits, systematic relationships among those exchange rates may help to explain the contagion nature of the crisis. Therefore, if East Asian economies continue to pursue capital account liberalization, the region's central banks would do well to stick to their post-crisis policies of greater exchange rate flexibility, in addition to strengthened financial supervision.

Although the focus of this study was on the region's exchange rates in the pre-crisis period, another interesting topic for future research is the region's foreign exchange market conditions in the pre-crisis period. It is possible that contagion during the crisis may have occurred via linkages among the exchange market pressure of regional countries.[19]

Finally, it should be pointed out that this study is first and foremost a study of *behavior* of pre-crisis East Asian exchange rates. Therefore, the paper at best draws indirect policy implications from the results, which should be interpreted with a great deal of caution.

[18] Thailand imposed controls on capital flows from 1995. Although effective in the short run, they could not prevent the devaluation of the baht (see Ariyoshi et al. 2000).

[19] For example, van Horen, Jager, and Klaasen (2006) investigate the effect of EMP of Thailand, the epicenter of the Asian crisis, on the EMP of four other crisis-hit Asian countries—Indonesia, Korea, Malaysia, and Philippines. They find some evidence of contagion for Indonesia and Malaysia but not for Korea and Philippines.

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About the Paper

Joseph Dennis Alba and Donghyun Park write that the exchange rate peg to the United States dollar is widely believed to have been a major cause of the Asian financial crisis of 1997–1998. Rigid exchange rates may have invited massive capital inflows into East Asia by creating a false sense of security among investors. A substantial empirical literature examines the actual behavior of pre-crisis exchange rates in the region. The authors seek to contribute to this literature by using daily data compared to other studies that tend to use monthly data and other lower-frequency data.

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