Filters and GARCH Methods for Nonstationary Sequences and the Effect of New Exchange Rate Regime[1]

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Abstract

After a long period of pegged, crawling or managed rates, Indonesia, Korea and Thailand have switched *de jure* to floating regimes after the 1997 Asian crisis. We focus on two issues: the volatility of exchange rates and the regime effect. Filters and GARCH type volatility models are applied to ascertain *de facto* variability in exchange rates. Vector autoregressive and error correction models are also used to clarify the regime effect. The initial results suggest a "floating regime along with risk management" in currency exposure, rather than the "fear of floating" or "return to pegging".

JEL classification: F31; F33

Keywords: Exchange Rate Regimes, Filtering, Volatility, the Philips curve

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1. Introduction

The recent literature concludes that the East Asian currencies have returned to *de facto* pegging of the domestic currencies to the US dollar since late 1998. Using the data of Indonesia, Korea, and Thailand, three countries severely affected by the Asian currency crisis (the ACA3), this study performs preliminary tests on (i) the empirical evidence of *de facto* regime in terms of volatility; (ii) the hypothesis that a country with a more flexible regime is less connected with international financial impact on domestic interest rates, "the impossible trinity"; and (iii) economic results of the choice of exchange rate regime, by estimating the slope of "the Philips curve". Our approach is different from those of the existing literature in that the filtering methodology (Pollock (1997, 1999, 2000, 2001b, and 2003b)) combined with the microstructural framework (Lyons, Killeen, and Moor(2001), hereafter LKM) and the GARCH type volatility models are used to verify the *de facto* variability in exchange rates. We therefore disregard the variations in interest rates and official reserves as the criteria to classify exchange rate regime. Recursive and structural vector autoregressive (VAR) models and error correction models (ECM) are used for examining the regime effects. The inference from the results suggests that a flexible regime with an appropriate risk management in currency exposure is critically necessary for the Asian small open-capital economies.

We review several issues raised in relation to exchange regimes. The recent empirical studies over the effects of the decision of whether or not to fix the exchange rate on monetary policy are mixed and continually debated. For developing countries in East Asia, a number of economists consider a flexible exchange rate regime as more appropriate (Fischer 2001; Rogoff, 1999). In contrast, other economists argue that a fixed exchange rate regime is a better choice for those countries (McKinnon and Gunther (2004), and McKinnon and Pill (1999)). In a conventional view on the exchange rate regime, a fixed exchange rates regime can reduce exchange rate volatility. Under a fixed exchange rate regime with perfect capital mobility, domestic interest rates move closely with the interest rate of the country in which domestic currency is pegged. On the other hand, floating exchange rates allow countries to have an autonomous monetary policy so that the monetary authority can set domestic interest rates independently. Under a floating regime, exchange rates can be insulated from shocks to international financial markets. However, examining the results on the effects of external shocks on domestic interest rates and exchange rates, it may be argued that the theoretical monetary trilemma[2] no longer exists. Monetary autonomy may be not possible as a result of the "open economy dilemma" regardless of the regime. The short term effects of external shock are less inconclusive than those of domestic variables, in the long run, however, international interest rates are significant on the domestic interest rates in the ACA3 regardless of their regimes. Hence, both the domestic and international interest rates are apt to be cointegrated in the long run under capital mobility. The existing literature agrees that the macroeconomic performances of countries, in terms of growth, investment, trade, and inflation, suggest a strong link between the choice of the exchange rate regime and macroeconomic performance. We examine the alternative regime effects particularly on domestic inflation and growth by assessing the slope coefficient of the Phillips curve. The preliminary results support the existing findings.

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[2] The trilemma is from the impossible trinity being that countries cannot simultaneously achieve fixed exchange rates; perfect capital mobility; and monetary policy autonomy.

The selection of our samples is of particular interest, for several reasons. First, these countries employed a diverse array of exchange rate systems during the period under review. Second, they share common characteristics that facilitate a comparative analysis, such as sharing the historical tendency to relatively lower inflation, sound fiscal balance, and lower unemployment with moderately higher savings and investment. It is also characterised by the fact that their currencies have not been internationally traded as a vehicle currency for settlement, transaction, and payment. Third, the currency crises allow us to test the bi-polar regimes on economic performance for each country. Throughout this paper, the US dollar and the federal yield rates are employed as the proxy for an anchor currency and international interest rates.

The remainder of the paper is organized as follows. In the next section, the classification of exchange regimes is briefly discussed. In section 3, two empirical methodologies are used to verify whether the ACA3 has adopted a new floating regime. In particular the digital Butterworth filter has been introduced to detrending the micro-components in conditional variance. Various GARCH type volatility models are assessed to fit the variations of exchange rates. In section 3 the transmission of international financial rates on domestic exchange rate and interest rate are examined. The brief results of new floating regime on inflation and growth are discussed in the same section. Finally, concluding remarks are in section 5. Some preliminary results are presented in Appendices.

2. East Asian Dollar Standard: Fear of Floating?

This section focuses on the question of how the Asian Crisis Affected Countries moved towards more exchange rate flexibility. Calvo and Reinhart (2002) and McKinnon (2003) conclude that the east Asian currencies have returned to *de facto* pegging of the domestic currencies to the US dollar since the introduction of floating regime. Several benefits accrue to a country with floating exchange rates. A floating regime allows that external shocks and imbalances are continuously adjusted and reflected in exchange rate movements. It also allows the country to pursue its own monetary policy without having to be overly concerned about the balance of payments effects. These benefits however will not have been realized in the ACA3, unless *de facto* the exchange rate is allowed to fluctuate. In fact, many countries that have announced floats suffer from a fear of floating and actually pursue exchange rate stabilization to varying degrees (the "fear of floating" by Calvo and Reinhart (2002)). Potential reasons for this are high interest rate variability and procyclical interest change.

Several studies were carried out in the early 1990s with the aim of identifying "*de facto*" exchange rate regimes[3]. Calvo-Reinhart (2002) and Reinhart (2000) classify the "*de facto*" exchange rate regimes with several criteria to sample a large number of countries over the period 1973-1999. They take account of the variance of exchange rates, interest rates and official reserves. Calvo-Reinhart (2002), there is the empirical probability that the monthly percentage change in the exchange rate will fall within a band of ±1% and ±2.5%. The same calculation is done for official reserves and interest rates[4]. Bénassy-Quéré and Coeuré (2002) propose a method aimed at improving anchor determination to account for *de facto* pegs to

[3]Among others, Popper and Lowell (1994) examine the exchange rate regimes in the United States, Canada, Australia and Japan by measuring central bank interventions through changes in official reserves and interest rates. Frankel and Wei (1993) identify results of the exchange rate policy by examining changes in parities. Poirson (2001) introduces a continuous indicator to measure the degree of flexibility of exchange rate regimes. The indicator is the ratio of exchange rate volatility to reserves volatility.

[4] The results show, for example, that the probability of the percentage change in the exchange rate to fall within a band of ±1% is only 27% for the USD/DEM and is much higher for the emerging countries.

currency baskets, which are overlooked in other classifications by giving a symmetrical role to all key currencies. They confirm that a large number of the currencies in the sample are estimated to be pegged to the dollar, while only few of them declared a peg to the IMF. Bubula and Ötker-Robe (2002) set a threshold that fits the allowed change in the exchange rate of a pegged currency with fluctuation margins of ± 1%. This is the size of the fluctuation band retained by the IMF for its definition of a pegged exchange rate. Reinhart and Rogoff (1999) improve upon existing methods, by using exchange rates on parallel markets for countries with a dual currency market. Their classification, so called "natural classification", is based on the percentage change of the nominal exchange rate in absolute value, and on the probability of remaining in a band of fluctuation and the level of inflation, over 40% as "free falling" and over 50% as "hyper float"[5].

The informal Asian dollarization reflects the rationale to enhance macroeconomic stabilization and avoid competitive depreciation, with the dollar as the dominant vehicle for international settlement and the storage of currency. By regressing currency baskets, McKinnon and Gunther (2004) show that the dollar weights in the East Asian currency basket are in post-crisis period widely unchanged, except Korea (0.97 to 0.86) and Thailand(0.92 to 0.78) (Table 2.1). In practice, East Asian countries coped with yen/dollar depreciation by temporarily lower dollar weights in the baskets. Daily exchange rate volatility is larger than the pre-crisis level, except in Indonesia. However, if the smaller East Asian economies adhere to their dollar pegs they remain vulnerable to yen/dollar fluctuations, because on a monthly basis, there is more volatility against the dollar and less volatility against the yen (Williamson (2000), Kawai (2002), Frankel and Wei (1993))[6]. Table 2.1 shows the prominent role played by the bilateral dollar exchange rate in the determination of changes in the multilateral real exchange rate (MRER) of the selected Asian countries. In the regression, the coefficients on the dollar variable are well above 0.5 and statistically significant at the 1% level. This means that an important share of the variations in the bilateral real exchange rate translates into changes in the effective real exchange rate (ERER) of the countries selected. R^2 results convincingly show that the single most important determinant of changes in the ERER of most Asian countries is the bilateral dollar real exchange rate. The dollar as the main anchor currency in the Asian countries is also evidenced by the results of Bénassy-Quéré and Coeuré (2002).

Table 2.1 Weights of the Currency Baskets

$$\Delta \ln(\varepsilon_t^i) = a_1 + a_2 \Delta \ln\left(s_t^{iUS}\right) + a_2 \Delta \ln\left(Yen/\$\right) + a_2 \Delta \ln\left(DM/\$\right)$$

	α1	α2:Dollar	α3:Yen	α4:DM	R^2
IND-pre-crisis	0.00	1.00***	-0.01	0.01	0.97
	(3.19)	(144.93)	(-0.92)	(0.85)	
IND-crisis	0.00	0.48***	0.64**	-0.15	0.02
	(1.12)	(1.06)	(2.35)	(-0.25)	
IND-post-c	0.00	0.95***	0.22***	0.02	0.27
	(0.02)	(13.10)	(3.11)	(0.23)	
KRW-pre-c	0.00	0.97***	0.06***	0.01	0.93
	(1.42)	(66.27)	(3.31)	(0.29)	

[5] Their classification is composed of 7 possible regimes: "peg" , "band", "crawling peg", "crawling band", "moving band", "managed float" and " "freely floating". A drawback of this classification is to be based only on the behavior of exchange rate and to neglect the changes in reserves, which can reveal the interventions of the central bank.

[6] During the 2004 and the early 2005, I have interviews with the top 10 leading dealers in foreign exchange market for Korean won, including City bank, JP Morgan. The question was what do you think the determinants of foreign exchange rate movement in the medium term, they answered that the movement of Japanese yen/dollar rates and government intervention are the major factors among others.

KRW-crisis	0.00	1.22***	0.05***	0.11	0.20
	(0.62)	(5.86)	(0.41)	(0.59)	
KRW-post-c	-0.00	0.86***	0.18***	-0.00	0.70
	(-0.01)	(30.34)	(9.13)	(-0.07)	
THB-pre-c	-0.00	0.92***	0.08**	-.01	0.95
	(-0.61)	(81.25)	(5.17)	(-0.35)	
THB-crisis	0.00	0.64***	0.32***	0.21	0.14
	(1.04)	(4.11)	(3.46)	(0.95)	
THB-post-c	0.00	0.78***	0.18***	0.05***	0.73
	(0.99)	(32.30)	(9.45)	(1.64)	
MAR-pre-crisis	-0.00	0.88***	0.09***	0.01	0.90
	(-1.48)	(54.80)	(5.30)	(0.45)	
MAR-crisis	0.00	0.70***	0.33***	0.11	0.20
	(1.39)	(5.33)	(4.19)	(0.59)	
MAR-post-crisis	-0.00	1.00***	-0.00	0.00	1.00
	(-0.01)	(979.34)	(-0.93)	(0.71)	
SPD-pre-c	-.00	0.82*	0.14*	0.08***	0.86
	(-1.32)	(34.37)	(4.83)	(2.97)	
SPD-crisis	0.00	0.69***	0.33***	0.02***	0.49
	(1.01)	(10.74)	(8.48)	(0.19)	
SPD-post-c	0.00	0.75***	0.19***	0.03***	0.90
	(0.46)	(47.95)	(14.98)	(1.45)	

Notes: t-statistics in parentheses, *significant at the 10% level, ** significant at the 5% level, ***significant at the 1% level, 869 observations for the period of pre-crisis, 415 for the crisis, and 1158 observations for the post-period, daily rates. Gunther Schnabl, Tubingen University together with McKinnon (2003), The East Asian Dollar Standard, Fear of floating and original sin, Hong Kong institute for monetary research working paper, No.11/2003. Pre-crisis covers 02/01/94-05/30/97, crisis for 06/01/97-12/31/98, and post-crisis covers 01/01/99-06/10/03, respectively. IND, KRW, THB, MAR, and SPD indicates Indonesian rupia, Korean won, Thailand baht, Malaysian Ringgit, and Singapore dollar. Note: IMF, 2001 survey foreign exchange market organization, and IMF, 2001 annual report on exchange arrangements and exchange restrictions (AREAER). Daily exchange rate volatility across exchange rate regimes, 2001 (in%); No separate legal tender-0.3; Currency board arrangements 0.53; Pegs against single currency-0.5; Pegs against a basket-0.42; Pegs within horizontal bands-0.65; Crawling pegs-0.46; Crawling bands-0.44; Managed floating 0.54;Independently floating-0.98.

Next, reviewing the literature on the *de facto* classifications by Levy-Yeyati and Sturzenegger, (2002, 2003) and Calvo and Reihnart (2002), exchange rate regimes are identified as floating systems with features of a highly volatile nominal exchange rate and a low level of intervention by the central bank (Table 1.1-1.2 in Appendix 1 and Table 2.2, below). They agree that the new exchange rate regimes of the ACA3 are categorized as managed or dirty floats or dirty float. It is however noteworthy that the level and variation of international reserves as a proxy for intervention by the central banks needs some caution . Because the determinants of international reserves are primarily from the external balance including the external flows in foreign assets and liabilities and the balance between current and capital account. For example, synchronous capital flows do not alter the net level of international reserves. Also it is common practice that interventions through financial derivatives such as swaps among the central banks are usually not recorded in balance sheet. In this aspect, we take a different way of examining the alternative regime by relying entirely on the fundamental trend in volatility of exchange rates. This is tried by extracting noise process and fitting the volatility from the nonstationary sequences of the exchange rates.

Table 2.2 Classification (De facto) of Exchange Regime: Indonesia, Korea, and Thailand

	INDONESIA	KOREA	THAILAND
1990 LYS*	Dirty/CP2	Dirty/CP2	Dirty/CP2
IFS**	Dirty	Dirty	Fix
KMC	Dirty/CP	Dirty/CP	Dirty/CP
1991 LYS	Dirty/CP2	Dirty/CP2	Dirty/CP2
IFS	Dirty	Dirty	Fix
KMC	Fixes	Dirty/CP	Dirty/CP
1992 LYS	Dirty/CP2	Dirty/CP2	Dirty/CP2
IFS	Dirty	Dirty	Fix
KMC	Fixes	Dirty/CP	Dirty/CP
1993 LYS	Dirty/CP2	Dirty/CP2	Irrelevant

IFS	Dirty	Dirty	Fix
KMC	Inconclusive	Dirty/CP	Inconclusive
1994 LYS	Dirty/CP2	Dirty/CP2	Irrelevant
IFS	Dirty	Dirty	Fix
KMC	Inconclusive	Dirty/CP	Inconclusive
1995 LYS	Dirty/CP2	Dirty/CP2	Dirty/CP2
IFS	Dirty	Dirty	Fix
KMC	Inconclusive	Dirty/CP	Dirty/CP
1996 LYS	Dirty/CP2	Dirty2	Irrelevant
IFS	Dirty	Dirty	Fix
KMC	Dirties/CP	Dirties	Inconclusive
1997 LYS	Dirty/CP2	Dirty/CP	Dirty/CP
IFS	Dirty/Float	Dirty/Float	Fix/Dirty
KMC	Dirties/CP	Dirties/CP	Dirties/CP
1998 LYS	Dirty3	Dirty/CP	Dirty/CP
IFS	Float	Float	Fix/Dirty
KMC	Dirties	Dirties	Dirties/CP

Sources: Levy-Yeyati, Eduardo, and Federico Sturzenegger, "Classifying Exchange Rate Regimes: Deeds vs. Words," mimeo, Universidad Torcutato Di Tella, 2002b, **IMF. Note: CP refers crawling peg; Dirty or Dirties means dirty floating system.

3. Filtering and Modelling Volatility of Foreign Exchange Rate

3.1 Filtering Nonstationary Sequences of Exchange Rates

This section is based on the theory and methodology by Pollock (1997, 1999, 2000, 2001b, and 2003b) who implements bidirectional frequency-selective filters in cases where the data sequence is short and nonstationary. The method has a firm theoretical basis and it is computationally efficient. The filtering methodology provides three main insights: (i) it demonstrates the detrended volatility captures the non-microstructural components in the conditional variance of the sequence; (ii) it reveals the patterns of smoothing operation in the disorderly FX market intervened by the monetary authority; and (iii) it presents the persistence of currency sequence. This would be revealed by de-trending the series and by isolating their noise processes.

The trend or signal sequence (*t*) is generated by a nonstationary autoregressive integrated moving-average (ARIMA) process, and the residual component (*t*) is generated by an ordinary stationary autoregressive moving-average (ARMA) process. Thus the data sequence *y*(*t*), which is a function mapping from the set of integers I = {t=0, ±1, ±2... } onto the real line, may be represented by y(t) = $\lambda_1 \sum_{\tau=1}^{t} \Delta R_\tau + \lambda_2 \sum_{\tau=1}^{t} X_\tau$ under flexible rates, where X_t refers the interdealer order flow, $X_t = \sum_{i=1}^{N} T_{it}$ with T_{it} denote the net interdealer trade initiated by dealer i. The real line therefore describes a cointegrating relationship between the level of the exchange rate, cumulative macro fundamentals and cumulative order flows. It is assumed that the cointegrating vector is regime dependent (Killeen, Lyons, and Moore (2001)). According to KLM (2001), under flexible rates, the changes in the exchange rate from the end of day can be written as, $\Delta P_t = \upsilon_t \Delta R_t + \upsilon_2 X_\tau$. *(t)* is the stochastic trend which it is rewritten with sample mean of Δp_{t+1} and foreign exchange earn, R_t, as $(1\text{-p})\Delta P_t + R_t$, and (*t*) is the residual sequence which can be rewritten as $(1\text{-p})^2$ ^{2}. Exchange rate sequence now presents as y(t) is

(3.1) y(t) = (t) + (t) = [T(L)]/(1-L)dv(t) + R(L)ε(t)

where *v*(*t*) and *ε(t*) are statistically independent sequences generated by normal white noise processes with *$V\{v(t)\}=\sigma^2 v$ and $V\{\varepsilon(t)\}=\sigma^2 \varepsilon$* respectively. *ψT(L)* is $E[P_{t+1} + \Delta R_{t+1}|\Omega]$ where

Ω denotes the information operator and *R(L)ε(t)* are rational functions of the lag operator *L* yields the $\frac{1}{\alpha\gamma}\left(\sum_{t=1}^{T} X_t\right)$, where α is the optimal trading parameter from public order flows, and γ is the elasticity of the aggregate public order flows. From the KLM(2001), η(t) =ψR(L)ε(t) is given $\sigma_R^2 \leq \frac{1}{4\upsilon_1 N\theta^2(1-p)^4\sigma_c^2}$, and for the stable solution, $\lim_{\sigma_R^2 \to 0} \sigma^2 \to 0$,

where σ_R^2, σ_c^2 are the payoff innovation and public order flow, *N* denotes the number of dealers, is the coefficient of risk aversion. *P* denotes the probability of a regime change, and υ_1, is the payoff discount parameter. The lag operator has the effect that *Lx*(*t*) = *x*(*t*-1). The operator *(1-L)- d* is a summation or integration operator for the nonstationary character of *(t)*. Let *Q`* be the matrix counterpart of the operator *(1-L)d* which produces *ζ(t) = (1-L)d (t)* and *k(t) = (1-L)dη(t).* Then

(3.2) Q′y = Q′ + Q`η = ζ + k = g, where E(ζ) = 0 and D(k) = Q`D(η)Q = σ2vΩs.

The differenced signal *z* is

(3.3) E(|g) = E(ζ) + C(ζ,g)D^{-1}*(g){g-E(g)} = Ωs(Ωs+λΩN)*$^{-1}$*g, where λ = σ2ε/σ2v.*

The estimator *k* of the differenced noise vector *k* is

(3.4) E(k|g) = E(k) + C(k, g)D^{-1} *(g){g-E(g)} = λΩN(Ωs + λΩN)* $^{-1}$*g.*

By solving the equation, *g*= *(Ωs + λΩN)b* for the value of b and *z = Ωsb* and *k = λΩNb.*

Minimize *(y-x) -1(y-x)* subject to *Q′x = z,* where *x* is the trend of from the estimate *z* of the differenced vector *ζ*. Consider Langrangean function

(3.5) L(x,μ) = (y-x)′Σ^{-1}*(y-s) + sμ′(Q′x-z).*

Differentiating the function with respect to *x,* we obtain

(3.6) Σ-1(y-x)-Qμ = Q′ΣQμ.

(3.7) Q′(y-x) = g-z = λΩRb = λQ′ΣQb,

From (3.7) we get *μ = (Q′ΣQ)* $^{-1}$*Q′(y-x) = λb.* for *μ* into (3.6) gives *x= y-λΣQb* which is the equation for estimating the recovered trend from *(t)=y(t) - η(t).* The *(αγ)*$^{-1}$ is increasing in *η(t)* and *k(t)/(1-L)d,* where $\alpha\gamma = \frac{1}{2}\left[\left(\frac{1}{\theta(1-p)^2\sigma^2} + \frac{2}{\theta\sigma^2}\right) + \sqrt{\left(\frac{1}{\theta(1-p)^2\sigma^2} + \frac{2}{\theta\sigma^2}\right)}\right]$ (KLM (2001)). The objective in fitting a major fundamental fluctuation is to avoid incorporating any components of the noise ($\sigma_R^2 \to 0$) within the finite-sample implementation, from aggregate order flows by dealers, customer order realizations, and speculators as well as transitory foreign exchange market intervention by the central bank. The stable solution ($\sigma_R^2 \to 0$) and the (υ_2, σ^2) of KLM analogously indicates that the (υ_2, σ^2) covariance tends towards the lower variance equilibrium, so that the low variance solution is stable by increasing the rate of transition of λ in H-P filter and the filter order *n* and reducing the designated cut-off frequency ω_c. The trend should contain all cycles that have a period longer than one year, and filtering the nonstationary short term sequences of exchange rates under the new regime. The Butterworth filter isolates the gap between the noisy components, and is bound to contain mechanism that ought to be consigned to the volatility under the new regime. The examples of application of the filtering by the Butterworth filter and the H-P filter and the Butterworth filter, depict the data on the volatility of monthly exchange rates, Indonesian rupiah, Korean won, and Thai baht (figure 3.1 and 3.2-3 in appendix 4).

Figure 3.1a The polynomial order 12 of monthly series of the logarithms of differentiated Indonesian rupiah (left); Korean won (centre); Thailand (right) for the years: 1990-2003

(i) (ii) (iii)

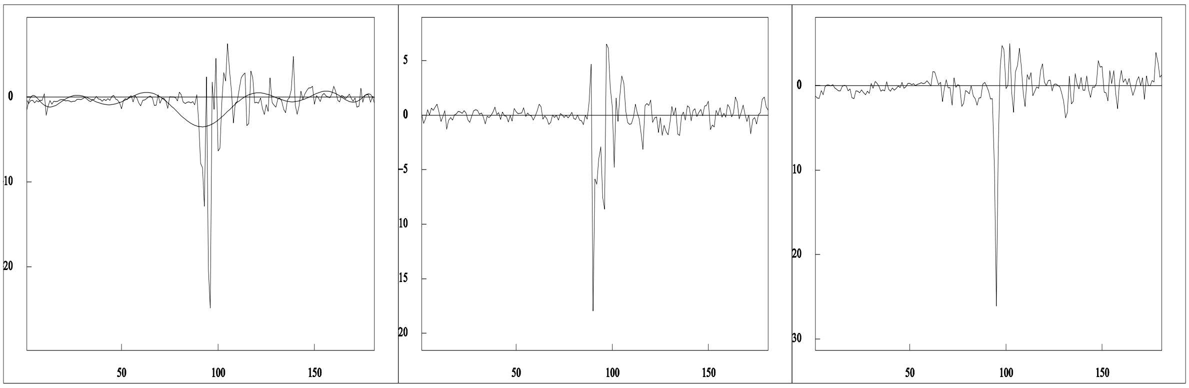
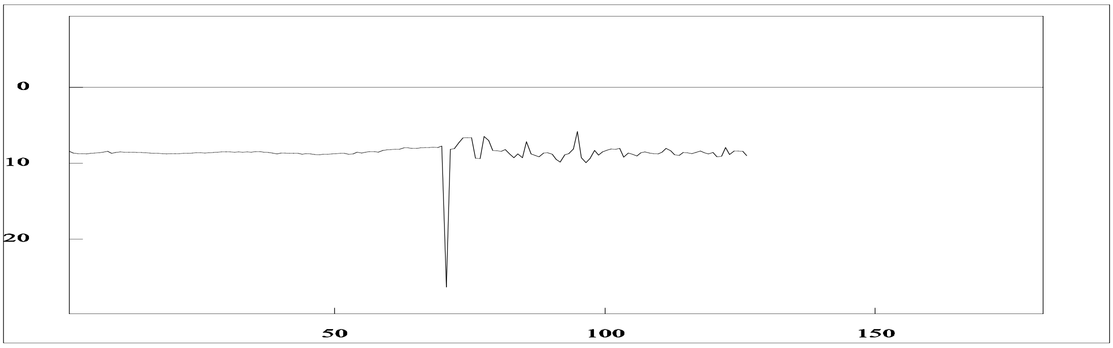
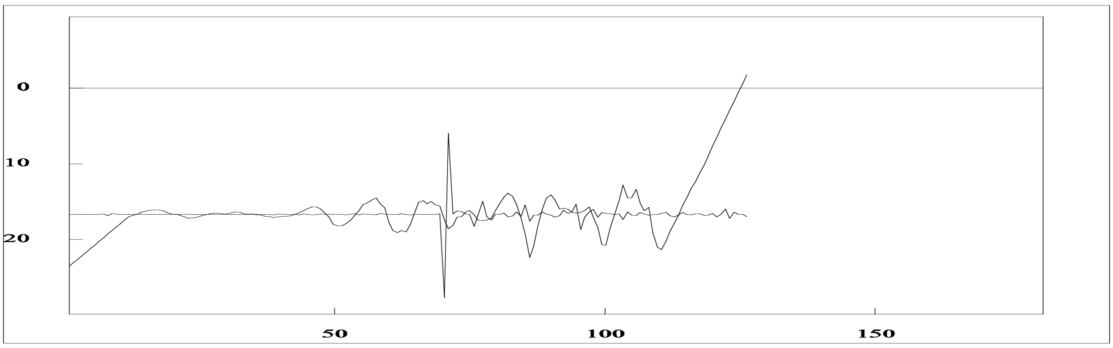


Figure3.1b The 12 order undirectional digital Butterworth filter with cut-off frequency of wu=π/4 and of 6th order lowpass with cut-off frequency of wL=π/16, superimposed on the same diagram. The monthly series of Indonesian rupiah; the 12 order lowpass (iv) and highpass of the twice differences of Indonesian rupiah (v); 15 order undirectional digital Butterworth filter with same specification of cut-off frequency of Thai baht: 1990M1-2004M12.

(iv) BWIWC



(v) BWIWC2D



(vi) 15BTW

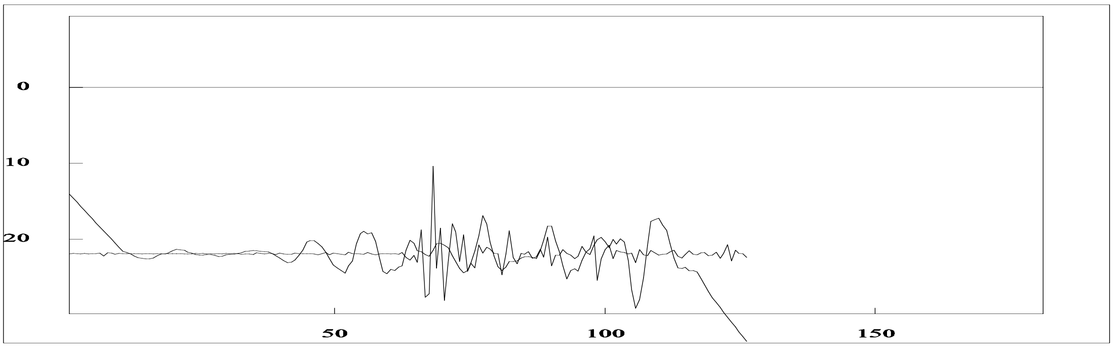
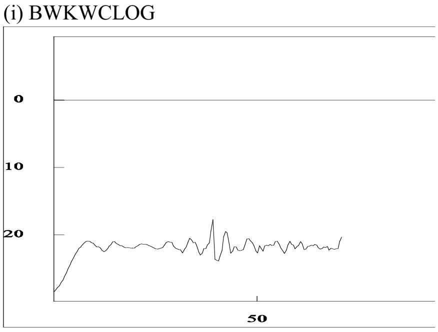
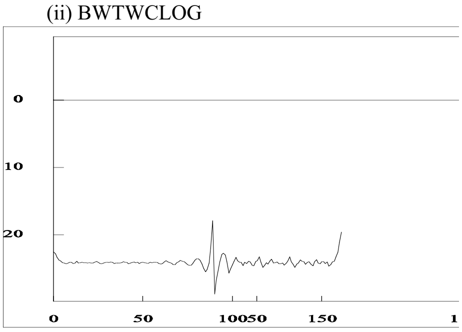


Figure 3.1c The 12 order high path digital Butterworth filter with cut-off frequency of wu=π/4 and of 6th order lowpass with cut-off frequency of wL=π/16, superimposed on the same diagram. The monthly series of the logarithmic Korean won (i); Thai baht (ii) and Indonesian rupiah (iii). Twice differenciated Korean won (iv) and Indonesian rupiah (v) and first differenciated Thai baht (vi), lowpass (vii) and highpass (viii) for Korean won: for the years 1999 to 2004.

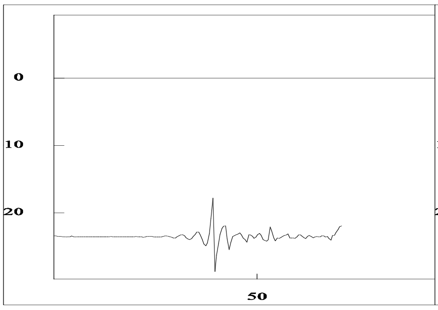
(i) BWKWCLOG



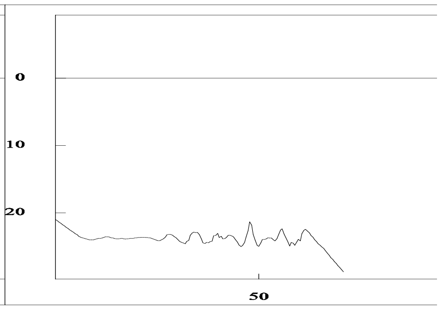
(ii) BWTWCLOG



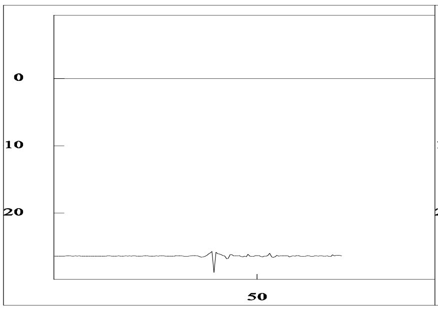
(iii) BWIWCLOG



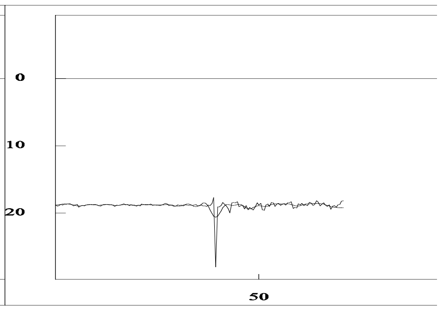
(iv) BWKWCFD2



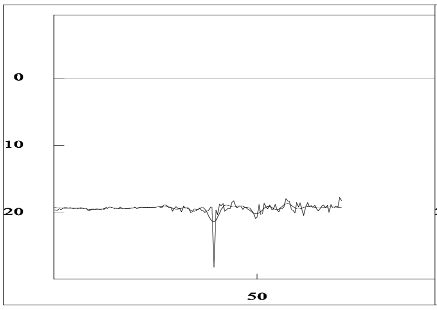
(v) BWIWCFD2



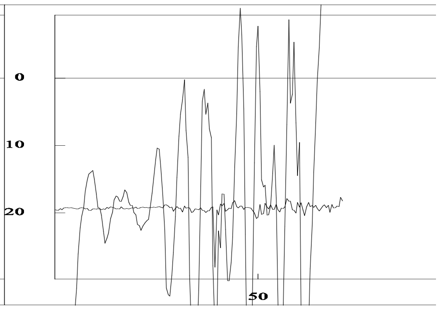
(vi) BWTWCFD1



(vii) BW40LKWC



(viii) BWKWCFD



A heuristic approach, applying two rational filters, Hodrick-Prescott (H-P) filter and Butterworth filter for the ACA exchange rate series, results in a signal-extraction filter has certain preconceived properties. For monthly data sequences, we impose the default values of λ=14400 for H-P filter[7] although λ can be determined via the ML estimation of the model. The results from fitting with H-P filter are presented in figure 3.2a and 3.2b of appendix 4. The broad trend clearly indicates the fluctuations for the post-crisis period for all the ACA currencies are higher and irregular, except for the Korean won. The fundamental fluctuations in Indonesian rupiah, Korean won, and Thai baht utilizing a Butterworth filter with an order of n = 12 and a differencing of degree d = 1, 2 are depicted in figure 3.1b and figure 3.1c. The results indicate that the rupiah, won and baht provide clear evidence of the floating regime effect, higher volatility, for the post-crisis period. The volatility trends for three currencies have irregular fluctuations and the amplitude of these fluctuations are not correlated among the three currencies. The results therefore ascertain that higher volatility and less FX market intervention by the monetary authorise under the new floating regime. In particular, the Butterworth filter captures the exchange rate market pressure (EMP) patterns at least one year earlier from the crisis in the data series of the rupiah and the baht. It may be therefore suggested that the filter may be applied detecting the early-warning signals of latent currency crisis.

3.2 Volatility Models and Estimating the Volatility under the New Currency Regime

The importance of estimating drift and volatility is that it is a central issue of modern financial analysis. The estimation of volatility is necessary as a proxy of risk in calculating valuation and portfolio selection of the security analysis and in pricing derivative securities, in computing value-at-risk, in determining the hedge ratio and a hedged position, and in market making for the bid-ask-spread. It is also critically important in setting monetary policy decision for central banks with estimating the volatility of price stability. We estimate the volatility models to ascertain the variation of foreign exchange rate under the new regime and assess which model is relatively unsurpassed fit for the selected Asian currencies.

Main empirical findings about volatility are in LeRoy and Porter (1981). They developed the volatility ratio test, and later, Fama and French (1988) ran the long horizon autoregressions. They conclude that drifts are time varying. Ghysels, Harvey and Renault (1996) report on stochastic volatility. The properties in financial volatility are identified as follows: fat-tails in the distribution, volatility clustering, highly persistent (long memory), leverage effect, and the states of volatility to be regime-switching (Hamilton, 1989).

[7] See Eview 3.1.

The GARCH class of models and the stochastic volatility (SV) class of models both estimate drift and volatility. These are conditional on past information and are the complements of each other in certain respects. For the SV models, the innovations to the volatility process are random and the volatility realization is therefore unobservable from data (Tauchen and Pitts, 1983). Melino and Turnbull (1990) use the GMM procedure (generalized method of moments) with the convergence of sample moments to their unconditional expectations, whilst QML (quasi maximum likelihood estimation) approach by Harvey, Ruiz and Shephard (1994) and its extension by Nelson (1991) rely on the nonlinear characteristic of Gaussian. SV model can be transformed into a linear non-Gaussian state space, so a Gaussian QML can be computed[8]. One of the most common measures of exchange rate volatility is the standard deviation of the growth rates of real exchange rates (V). This measure is approximated by a time-varying measure defined as follows:

$$V_{t+m} = \left[\frac{1}{m} \sum_{i=1}^{m} (R_{t+i-1} - R_{t+i-2})^2 \right]^{1/2} \tag{3.8}$$

where *R* is the natural log of the bilateral real exchange rate (*s*) and *m* is the order of the moving average.[9] An alternative measure of exchange rate volatility is defined as the time-varying coefficient of variation (CV) of the real exchange rate (this is, in fact, a measure of dispersion of the real exchange rate).

$$CV_{t+m} = \frac{\left[\frac{1}{m} \sum_{i=1}^{m} (\varepsilon_{t+i-1} - \bar{\varepsilon})^2 \right]^{1/2}}{\bar{\varepsilon}} \tag{3.9}$$

where $\bar{\varepsilon}$ is the mean of the bilateral real exchange rate between month *t* and *t+m-1.* Figures 1.2 - 1.4 show both measures of volatility[10] for the bilateral ACA3 exchange rates. A simple comparison between our two measures of exchange rate volatility highlights the main differences between them. For example, the CV measure indicates a large increase in the dispersion of the ACA3 real exchange rates in 1997 for the IDR/USD. In the 1980's for the KRW and IDR/USD, and in the late 1970s for the KRW/USD, the volatility measure (V) only suggests moderate increases in volatility in the late seventies and eighties for the three real exchange rates except during the crisis. From both measurements, the result shows that the volatilities for the three exchange rates have been increased approximately 15-25 % compared to the volatility of the pre-crisis fixed exchange regime.

The figures 1.2-1.5 suggests that the standard measure of volatility (V) misrepresents the real exchange rates and fails to identify periods of rapid but sustained change in the RER. On the other hand, the coefficient of variation measure is successful in capturing these events (as

8 The QML can implement and extend to nonstationary and multivariate and it allows for prediction. SMM (simulated method of moments) the simulation technique also can be used with the computer intensive Markov Chain Monte Carlo (MCMC) to draw from the distribution of volatilities conditional on the observations. The MCMC posterior mean exhibit root mean squared errors anywhere between half and a quarter of the size of the GMM and QML point estimates. The root-mean squared error of the posterior mean produced by the Bayesian filter is smaller than the point estimate produced by an approximate Kalman filter supplied with the true parameters. Recently, some papers suggest modelling drift and volatility together using a Vector Autoregressive (VAR) approach. VAR assumes both drift and volatility evolve jointly as a first order vector autoregressive process.

9 It is interesting to note that some authors use this indicator as if it were a measure of the standard deviation of the RER (and not of its growth rate).

10 Empirical regulates in the case of financial returns (S&P index for example), volatility clustering, fat/heavy tails (Kurtosis, k=3).

occurred, with the local/dollar rates in 1986, 1997)[11]. Both volatility measurements exhibit the volatility of exchange rates for the three countries which has increased to 2.6-6.6 times more under the floating regime since 1997 compared to the pre-crisis fixed regime. Coefficient of variation (CV) of real exchange rate captures the volatility more efficiently for the events than that of standard deviation of growth rate of real exchange rate.

3.2A Volatility Models and Estimating Volatilities

The general models of financial volatility include unconditional volatility; EWMA (exponentially weighted moving average); GARCH (generalized autoregressive conditional heteroscedasticity); and parameter estimation by ML (maximum likelihood methods)[12.] The Generalized Autoregressive Conditional Heteroscedasticity (GARCH) model was developed by Bollerslev and Engle (1986) based on the Autoregressive Conditional Heteroscedasticity (ARCH) model developed by Engle (1982). The advantage of the GARCH specification over the ARCH specification is its parsimony in identifying the conditional variance in GARCH. In empirical application, GARCH (1,1) model is the most popular structure for many financial time series. The main empirical features of exchange rate volatility contain long memory[13] and fat tail. Recent literature on volatility of exchange rates have applied Engle's (1982) autoregressive conditionally heteroskedastic (ARCH) model and stochastic volatility (SV) models. In the stochastic volatility model (SV) for foreign exchange rates, the two jump processes have a similar specification, but the correlation between the Brownian motion can be significantly different and the dependence between the jump processes can be different.

A SV model for foreign exchange rates[14] indicates that short term variation is dominated by the random disturbance terms. Adopting a variety of approaches, univariate GARCH, GARCH-M, EGARCH, and FIGARCH models are applied to find convincing evidence that does not result from overlooking a particular volatility model. Each GARCH-type model applied is briefly summarised below. The purpose to utilize Auto-Regressive Conditional Heteroscedasticity (ARCH) and Generalized ARCH Models is that they allow conditional variance ε_t to vary over time, $E[\varepsilon_t^2 \mid \Omega_{t\ 1}] = h_t$ for some nonnegative function. In the ARCH (q) model (Engle, 1982) with conditional heteroscedasticity, the variance of $u_t = Var(u_t \_\ u_{t-1}, u_{t-2},...) = E((u_t - E(u_t))^2 \_\ u_{t-1}, u_{t-2},...)$, the error variance depends on q lags of squared errors, h_t. The specification given by Engle (1982)

$$(3.10)\ y(t) = b1*x1(t)+b2*x2(t)+ ... +bk*xk(t)+u(t),\ \ \sigma\hat{}2= (u(t)|\Omega(t-1)),\ h(t)\hat{}2$$

$$= a0+a1*u(t-1)\hat{}2+ .... +aq*u(t-q)\hat{}2+f1*h(t-1)\hat{}2+ .... +fp*h(t-p)\hat{}2+d'w(t).$$

where $h(_t)^{\wedge 2}$ is the conditional variance of $u(_t) \sim N(0,h_t)$, $u_t = v_t\ \sigma_t$, with respect to the information set $\Omega(_{t-1})$. $w(_t)$ is a vector of pre-determined variables assumed to influence the

[11]The comparison between the V and VS circulation show that both measurements show the volatility has been increased for the floating regime since 1997. Thai Baht becomes much volatile (3-7) times more volatile for the post-crisis period), next is to Indonesian Rupiah which shows 2.7-6.6 times more volatile for the post-crisis period, finally Korean won, the volatility under the floating regime exhibits 1.8 to 4.3 times more volatile. The details of the results: * indicates VS and ** for V respectively.

KRW	fixed regime9.565834*	0.091106**floating regime	41.35716	432.34%	0.162358	178.21%
THB	fixed regime 0.207359*	0.017087**floating regime	1.511263	728.82%	0.060322	353.03%
INR	fixed regime50.66766*	0.539512**floating regime	335.2135	661.59%	1.446033	268.03%

[12] Uncertainty volatility models include implied volatility based on an option pricing, the σ parameter to plug into the Black-Scholes formula to match the corresponding market price. Implied volatilities vary with strike and maturity, namely σ=σ(K,T). implied volatility curves are usually skew-shaped or smile shaped.

[13] AR-type in the realized volatility and Heterogeneous Autoregressive model of the Realized Volatility (HAR-RV).

[14] $dlnS=(r_d-r_f-1/2v)dt + \sqrt{v}\, dZ_1 + dJ^*_1$ and $dv=(\kappa v'-\kappa v-\lambda v)dt + \sigma\sqrt{v}\, dZ_2 + dJ_2^*$. While long dated multi-currency options and derivatives are dlnS=(rd (y)-rf (y) - 1/2v(y))dt + √v(y) dZ0+ dJ*0 and d(yj) =(κjθj-κjyj-λjyj)dt + σj√yj dZj + dJ2* where j=1, …,κ.

conditional error variances in addition to the past squared errors. In the GARCH (generalized ARCH) process, the conditional variance of *u/* at time *t* is dependent not only on past squared disturbances but also on past conditional variances. GARCH (p, q) model is an extension to a generalized ARCH model to describe time varying variances where *d = 0*. The sum *a1*u*$_{(t-1)}$*)^*^{2}*+ .... +aq*u*$_{(t-q)}$*)^*^{2} will be referred to as the MA part and *f1*h*$_{(t-1)}$*)^*^{2}*+.... +fp*h(*$_{t-q}$*)^*^{2} as the AR part of the GARCH model[15]. GARCH (1,1) process *σt^2=α+γσ^2*$_{t-1}$*+δy^2* $_{t-1}$, is of the form which requires *γ+δ<1*. The underlying specification given by the equation (3.10), the conditional variance of $h_t = \omega + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$, is a function of an intercept *( ),* a shock from the prior period α_1 and the variance from last period β_1. The conditional variance, h_t^2 of u_t, with respect to the information set ω_{t-1}, and w_t is a vector of predetermined variables assumed to influence the conditional error variances in addition to the past squared errors.

GARCH(p,q)-M (GARCH (p,q) in Mean) model enables us to computer ML estimates that the estimation of GARCH-M models generally provides testing the rates' efficiency (Engle, Lilien and Robins (1987)). GARCH-M estimation indicates some function of the conditional variance, usually the standard deviation in the mean equation where there is a relationship between risk and return of an asset. This allows the mean of a series to depend on the conditional variance of the series. Maximum log likelihood (ML) estimates of the GARCH-M model can be expressed as

*(3.11) y(t) = b1*x1(t)+b2*x2(t)+ ... +bk*xk(t)+γ* h(t)^2+ u(t)*

*and, σ^2 (u(t)|ω(t-1)) = h(t)^2 = a0+a1*u(t-1)^2+ .... +*

*aq*u(tq)^2+ f1*h(t-1)^2+ ....+fp*h(t-p)^2 +d'w(t).*

where *h(t)^2* is the conditional variance of *u(t)* with respect to the information set *ω(t-1),* and *w(t)* is a vector of pre-determined variables assumed to influence the conditional error variances in addition to the past squared errors. The variables *x1(t), ... , xk(t)* and *w(t)* must be in the variables and can include lagged values of *y(t).*

EGARCH (p,q) model is introduced by Nelson (1991), and is used to compute ML estimates of the exponential GARCH(p,q) model where the logarithm is the conditional variance of the errors. The value of μ depends on the density function assumed for the standardized disturbances, *εt=ut/ht.* This model allows for the possible asymmetric effects of past errors on the conditional error variances[16]. GMM yields consistent and asymptotically normal estimators for EGARCH.

$$(3.12)\quad y_t = \beta' x_t + u_t ,$$

$$V(u_t \mid \Omega_{t-1}) = h_t^2 = \varepsilon_0 + \sum_{i=1}^{q} \alpha_i u_{t-i}^2 + \sum_{i=1}^{p} \phi_i h_{t-i}^2 + \ 'w_t$$

$$(3.13)\quad y_t = \beta' x_t + \gamma h_t^2 + u_t, \ where\ h_t^2 = V(u_t \mid \omega_{t-1}) ,$$

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[15] Since the conditional variance needs to be nonnegative, if α1=0, then the conditional variance is constant and $_t$ is conditionally homoscedastic.

[16] The sum of (u($_{t-1}$)/h($_{t-1}$))+ ... +aq*(u($_{t-q}$)/h($_{t-q}$))+c1*(ABS(u($_{t-1}$)/h $_{t-1}$)-μ)+ ... +cq*(ABS(u($_{t-q}$)/h($_{t-q}$))-μ) is referred to as the MA part and f1*logh($_{t-1}$)^2+ ... +fp*logh($_{t-p}$)^2, as the AR part of the EGARCH(p, q) model, respectively.

$$h_t = \sqrt{V(u_t \mid \Omega_{t-1})} = {}_0 + \sum_{i=1}^{q} \alpha_i \left(|u_{t-i}|\right) + \sum_{i=1}^{p} \phi_i h_{t-i} + \delta' w_t$$

while

$$\sum_{i=1}^{q} \alpha_i \left(|u_{t-i}|\right) \log h_{i=1}^{q} = \varepsilon_0 + \sum_{i=1}^{q} \alpha_i \left(\frac{u_{t-i}}{h_{t-i}}\right) + \sum^{q} \left(\left|\frac{u_{t-i}}{h_{t-i}}\right| - \mu\right) + \sum_{i=1}^{p} \phi_i \log h_{t-i}^2 + \delta\, w_t$$

With $\mu = E\left(\left|\frac{u_{t-i}}{h_{t-i}}\right|\right)$.

EGARCH-M, Exponential GARCH in Mean enables us to estimate the EGARCH (p,q)-in-mean model.

It may can be generalized that the long memory form of $\sigma_t^2 = c_0 + \sum_{j=1}^{\infty} c_j y_{t-j}^2$ has $c_j = \gamma^{j-1}$, while strong GARCH is ${}_t = (y_t\text{-}\mu_t)/\sigma_t$ *is i.i.d.,* semi-strong GARCH is $E({}_t|F_{t-1})=0$ and $E(\hat{}_t|F_{t-1})=1$; weak GARCH if it is mean zero and uncorrelated, $\sigma_t^2 = E_L[\,{}_t^2 \mid I_{t-1}]\ where\ i_{t-1} = \{1, y_{t-1}, y_{t-2}, \ldots y_{t-1}^2, y_{t-2}^2, \ldots\}$. FIGARCH (Fractional integration) leads to an expansion of higher order GARCH models with a conditional variance, $h_t = \omega + \sum_{j=1}^{p} \alpha_j \varepsilon_{t-j}^2 \sum_{k=1}^{q} \beta_k h_{t-k}^2$. The FIGARCH model is considered under the IGARCH model in the conjecture of Baillie, Bollerslev and Mikkelsen (1996). The memory property of the series $(1-L)^d \sigma_t^2 = \alpha + y_{t-1}^2$, where *d* determines the property, if *d* equals *1,* it is the standard IGARCH model[17], for $d \neq 1$ can define the binomial expansion of $(1\text{-}L)^{-d}$ in the form (Robinson et al, 1997). The FIGARCH case corresponds to $\alpha=1$, and stable GARCH case corresponds to $\alpha=0$. When $d=1$, the parameter α reduces to an autoregressive root. Hence the model becomes either a GARCH or IGARCH, depending on whether $\alpha<1$ or $\alpha=1$. When the model reduces with $d=1$, $\delta(L)=1$ and $\beta(L)=1\text{-}\beta 1L$, *it* becomes to the covariance nonstationary GARCH (1,1). In *(3.10)*, $d>1$ gives rise to negative coefficients and so is not permitted. The motivation for using *(3.11)* must be rooted from the FIGARCH and IGARCH cases. If the GARCH component observes the usual covariance stationarity restrictions, this implies $\delta(1)/\beta(1)>0$, with $\alpha<1$ where these processes are covariance stationary. L_T is the function for estimation in the Student's t log-likelihood. Small u_t in this criterion works in a similar way to the Gaussian log-likelihood, but large innovations such that $u/(v\text{-}2)\ ht>>1$ seems smaller to the aggregate than in the Gaussian case, depending on the size of v[18].

3.2B Estimating GARCH Models for the selected Asian Currencies

The purpose of the GARCH type models applied in this paper is to clarify the volatility of the exchange rates under the new floating regime introduced in 1997. The volatility of exchange rates is estimated with various GARCH type models from the data of daily currency prices

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[17] IGARCH (Engle and Bollerslev, 1986) which is a linear GARCH variation that similar to integrated series in regular (ARMA-type) time-series. This occurs when $\alpha+\beta=1$. When this is the case it means that there is a unit root in the conditional variance; past shocks do not dissipate but persist for very long periods of time.

[18] The results of HYGARCH model show a remarkable degree of uniformity. The point estimates of each parameter seem to be different in the sampling error to be expected from data generation processes as exchange rates are determined in related markets. We conjecture that the similarity of these structures in their movements in levels. While the estimates are small, they are generally significant. On the other hand the hyperbolic memory in variance, measured different from 1, a little below 1.

both spot and forwards rates of 6 and 9 months for Indonesian rupiah (IDR), 6 month and 1 year for Korean won (KRW) and Thai baht (THB). All data are obtained from DataStream and the Asian Development Bank (ADB). The nominal effective exchange rate (EER) indicates $log(e_t)\text{-}log(e_{t-1})$, and the volatility of EER is computed from the daily return. The descriptive statistics for the whole period and two sub sample periods, pre- and post- crisis for the three currencies are summarized in Table 1.3 in Appendix 1. The standard deviation and standard error are higher and the curvature in kurtosis is flatter under the new exchange rate regime and in all ACA samples. The degree of volatility tends to increase with the frequency with which observations are sampled. This can be seen in the series of spot daily exchange rates covering the period from January 3 1995 to January 31 2003 that in total contain 2387 daily observations show higher than for monthly or yearly. The currency rate of its own lags, is the one, two, five, and ten order that shows after one lag, it decays with large magnitude toward zero base line with mixed sign. For the estimation of GARCH type models, the conditional variance of daily and high frequency returns in bilateral exchange rates will be used as a proxy for the anticipated volatility of exchange rate. The Wald statistic for testing the joint hypothesis, states that $\alpha 1= \ 1=0$ is equal to 4.93, which is significant at the 95 per cent level. From the estimation of GARCH models for h^2_t, the conditional variance of log, the mean equation for the exchange rate changes follows the AR (1) process $(x_t = \beta 0 + \beta 1 x_{t+1} + u_t)$ where $x_t = \ log(.)$ and u_t is a white noise process $(E(u_t)=0; E(u_t, u_{t-j})=\{\ ^2 \text{ for } j=0, \text{ and } 0 \text{ for } j\neq 0\}$ and $|\beta 1|<1$). Therefore, x_t is a covariance-stationary process. While the conditional mean of x_t changes over time, if the process is covariance stationary, the unconditional mean of x_t is constant. That is $E(x_t)= \beta 0/1\text{-}\beta 1 x_{t-1} +u_t E(x_t) = \beta 0/1\text{-}\beta 1$, where Ex_t denotes $E(x_t| x_{t-1}, x_{t-2},...)$, the linear regression of x_t on constant and the past observations. Its conditional variance $(\ ^2)$ could be time varying. Such a process is the autoregressive conditional heteroscedastic process of order 1 denoted u_t ~ARCH(1) or GARCH (0,1), $V(x_t|\Omega_{t-1})=V(ut|\Omega_{t-1})=h^2_t =\alpha 0+\alpha 1 u^2_{t-1}$, if $\alpha<1$, then the unconditional variance of ut is given by $\ ^2=\alpha 0/1\text{-}\alpha 1$. If h^2_t evolves according to $h^2_t =\alpha 0 + \alpha 1\ u^2_{t-1} +...+\alpha q\ u^2_{t-q} + \ 1h^2_{t-1}+...+\ p\ h^2_{t-q}$, the generalized autoregressive conditional heteroscedastic model, denoted by u_t ~GARCH (p, q). The well established GARCH (1, 1) process, $V(u_t|\Omega_{t-1})=h_t t^2 = \alpha 0 + \alpha 1 u^2_{t-1} + \ 1h^2_{t-1}$ where $\alpha 0>0$, $|\ 1|<1$, and $1\text{-}\alpha 1\text{-}\ >0$. These restrictions ensure that the unconditional variance of u_t given by $V(u_t)=\alpha 0/(1\text{-}\alpha 1\text{-}\ 1)$ is positive.

We specify the AR(2) for the conditional mean of log of exchange rates, to fit a GARCH (1, 2) model to the errors of the regression equation with 0.1 for MA lag 1, 0.2 for MA lag 2, and 0.6 for AR lag 1. The process converges after 71 iterations. However the second order coefficient of the MA part of the process is not statistically significant. Therefore, a GARCH (1,1) model seems to fit the observations reasonably well. When the coefficients of the GARCH add up to unity, the model is known as the integrated GARCH (IGARCH) model, and this implies that the shocks to the conditional variance are persistent. This model can be estimated using the GARCH-in-mean with the values of 0.5, 0.1, and 0.2 for the initial estimates of , α_1, and $\ 1$. According to the results, the ML estimate of has the correct sign that it is the evidence of a GARCH-in-mean effect in this model. The evidence on the volatility of the conditional variance of u_t or h^2_t is rather mixed. The ML estimate of Ω_1 is highly significant, but the ML estimate of α_1 is near zero and is not statistically significant. Estimating the GARCH-in-mean model with initial values 0.1 in mean, we find 0.2 in MA lag 1, and 0.4 AR lag 1 for the initial estimates of , α1, and $\ 1$. Comparing these results with those in Appendix 2 suggests that the other specification fits the data much better than the other. Even the Schwarz Bayesian criterion unambiguously selects the model with conditional errors, ' equals 8.72 and hence the hypothesis that =0 cannot be rejected. The non-normal errors can be assessed graphically. The ‘outliers’ of the graph, refers to the 1997 currency crisis. EGARCH-in-mean models however didn’t fit for both entire period and pre-crisis period with the specification of 0.1 in γ, -0.2 in $\alpha 1$, 0.1 in α’, and 0.3 in   1, except the 1 year forward rates of Korean won for the post-crisis period.

Estimation results for the GARCH models are summarized in the tables of Appendix 2. The top half gives the ML estimates of the regression coefficients, β, and the estimate of γ when a GARCH-in-mean model is estimated, which indicates asymptotic standard errors and t-ratios, as well as a number of summary statistics and model selection criteria. The ML estimates of the parameters of the conditional variance model together with their asymptotic standard errors are presented in the note section below of the table. The plot estimates of the conditional standard errors and plot of histogram of the standardized residuals $(e_t'=u_t'/h_t')$ are illustrated in Appendix 2. Plot residuals provide a plot of adjusted residuals, together with a standard error band that represents $\pm 2\sigma'$ as the estimated standard error of the regression.

The GARCH type models for fitting the volatility of the bilateral exchange rates in the ACA, the estimators and forecasting errors enable us to generally conclude that GARCH(p,q) models fit well for both spot and forward currencies in the ACA (Table 3.2 and Appendix 2). The shocks coefficient, $\alpha_1+\beta_1=\varphi$, the φ for the spot rates exhibits higher than that of the 6 and 9 month forward rates for all ACA currencies in particular the 6 month Thai baht forwards (0.95 and 0.42 for spot and forward respectively). The highest φ is to the Korean won spot rates with EGARCH indicates the effect of a floating regime in Korean won compared to those of the rupiah and the baht. Also the coefficient α_1 increases in the forward rates of all three currencies, suggesting that the fluctuations of exchange rates result in information being impounded into the forward rates more rapidly and more slowly in the spot rates for all currencies in the ACA.

Other volatility models, AGARCH-M (Absolute value GARCH(p,q) in Mean)[19], the HYGARCH (hyperbolic GARCH’ model, Davidson, 2003), ‘nonparametric and semiparametric models[20]’ (Engle and Gonzalez-Rivera, 1991, Linton, 1993 and Drost and Klaassen, 1997) and multivariate models[21] are considered although they are left for further studies.

The limitations of the GARCH estimation may be due to the estimation errors from: 1) inappropriate initial values, 2) large damping factor, 3) existence of serial correlation in the residuals of the regression[22], 4) no ARCH effects in the model[23] and, 5) inappropriate assumptions about the conditional distribution of errors[24].

Instead of using the usual criteria, likelihood, AIC, BIC, we can choose the best model using the Black-Scholes pricing error, this generates results by investigating the structure in mean and variance and model performance using in- and out- sample volatility forecast. An outstanding research topic may be able to be extended by establishing an optimal choice of

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[19] It enables us to compute ML estimates of the following the GARCH model where the conditional standard error disturbances, u_t.

[20] It can be used to obtain more efficient estimates of parameters when the error distribution to be of unknown functional form, are considered although they are left for further studies non-Gaussian. Yt=β’xt+εtσt, σt^2=α+γσt-1^2+δ(yt-1-β’xt-1)^2, where εt is i.i.d. with density of the standardized residuals εt=yt-μt/σt

[21] They are defined as $\sum_t = E(ytyt' \mid Ft-1)$ for some *(n x 1)* vector of mean zero series.

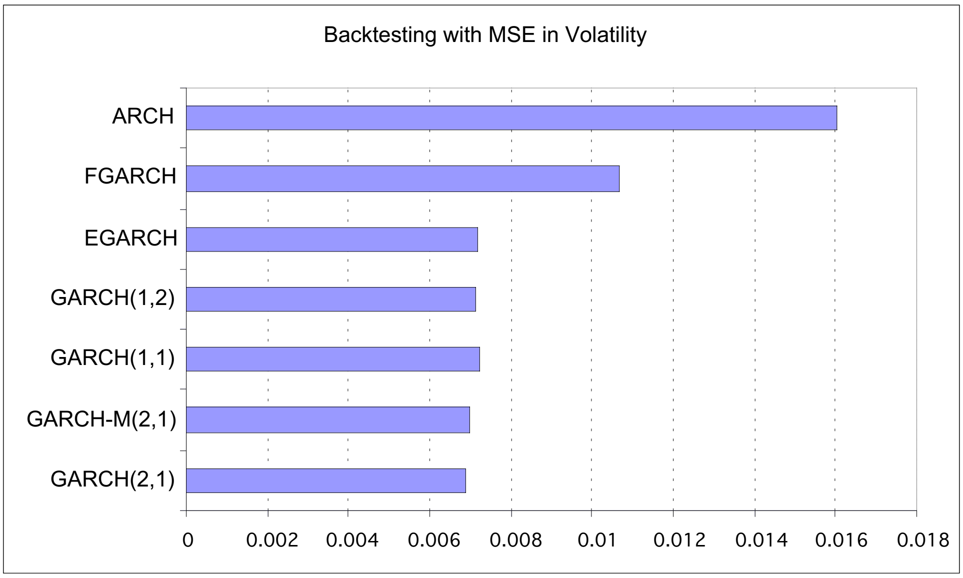
[22] This problem may be solved by adding lagged values of the dependent variable to the model.

[23] This problem can be solved by hypothesis testing after an OLS regression.

[24] In the case of this, the conditional distribution of the errors from normal to student t distribution can be alternatively considered.

delta which depends only on observable data to solve the problems in pricing with path-dependent options and volatility swaps in currencies. Examining the "mixture distribution hypothesis" and "sequential information model" on the dynamics from the realized volatility of exchange rate, order flows, and FX EBS (electronic broking system) are also interesting areas for further research.

Figure 3.2 Results: GARCH models Backtesting in Volatility



Notes: The averaged results are based on the three countries, ARCH(1): σ^2t+1=w+α t^2, GARCH(1,1): σ^2t+1=w+α t^2 +βσt^2; GARCH(2,1): σ^2t+1=w+α1 t^2 +α2 t^2 +βσt^2 ; GARCH(2,1) WITH MEAN AR(1): σ^2t+1=w+α1 t^2 +α2 t^2 +βσt^2 and rt+1=θ0+θ1rt ; GARCH(1,2): σ^2t+1=w+α t^2 +β1σt^2 + β2σt-1^2 ; FGARCH(1,1): σt+1^2=w+α ^2t+β1σt^2+y t2*1{ <0} ; if negative correlation between the volatility changes and returns, run aj regression on diff(log(sigma))(t)=α+β*returns(t-1)+ t αestimate: -0.0001; β-estimate=-0.8574; R^2=0.0177; F-stat=40.7679(t-test), p-value=0.0000, there exists a negative correlation between returns and changes in volatility.; Exponential MA (RISKMETRICS): σt+1^2=λσt^2+(1-λ)rt^2, λ=0.94(0.96); the best model can be selected by maximizing the likelihood or minimizing MSE(choose the largest), AIC(between MSE and BIS), and BIC(the smallest model). MSE equals 1/TΣt=1>T.e^2_t; AIC equals -2L+2/T.k, BIC equals -2L+log(T)/T.k.

4. The Role of Exchange Rates under New Floating Exchange Regime

The basic proposition of international macroeconomics theory is the notion of the open economy trilemma whereby countries can pursue two of three options, fixed exchange rates, domestic monetary autonomy, and capital mobility. A number of recent empirical studies on emerging economies investigated domestic financial markets and their response to international financial shocks. These studies include Borensztein et al. (2001), Habib (2002), Frankel et al. (2002), Shambaugh (2004), and Obstfeld et al. (2004). Borensztein et al. (2001) show evidence that the implication of the exchange rate regimes on the effects of external factors on domestic interest rates is inconclusive. Their results are consistent with the findings of Habib (2002). While using a large number of country panel data with a longer sample period, Frankel et al. (2002) find that the countries under a floating exchange rate regime appear to have a degree of temporary monetary independence. Shambaugh (2004) finds that domestic interest rates in countries under a pegged exchange rate regime follow the interest rate movements in the country to which the currency is pegged. Similarly, Obstfeld et al. (2004) have extended Shambaugh's (2004) paper, and find that a non-pegged exchange rate regime gives more room for monetary policy autonomy.

Using a simple recursive and structural VAR and error-correction model (ECM) on the data of ACA3 countries, we examine whether the alternative exchange rate regime responds in a different way to the base country interest rate. The results from our empirical evidence suggest that there is no trade-off between adopting a floating exchange rate and formulating autonomous adjustments in monetary policy with unrestricted open economy variables. The domestic interest rates of the ACA samples have been increasingly cointegrated with the US interest rates. The results are associated with the openness of the capital account of the countries rather than their exchange rate regimes. The results also imply that the monetary trilemma is unrealistic, since decisions to reject pegging and maintain a floating regime did not bring about the expected monetary autonomy in the small open economies selected. We therefore argue that although a country floats its exchange rate, it may not be able to pursue domestic goals with its monetary policy under liberalized capital mobility[25]. It therefore may be argued that a fixed regime may have more room for autonomy than the floating regime.

A short-run dynamic specification follows the methodology of Pesaran, Shin, and Smith (PSS) (2001), Frankel et al (2002) and Shambaugh (2003). The specification provides estimation in the error correction form and examines the statistical significance of the coefficient on the levels relationship. We examine whether the levels of domestic interest rates (R_i/MILR) and base country interest rates, the US federal rates (R_{bi}/UDCR) are cointegrated when the two nonstationary series become in fact stationary when combined. Both domestic interest rates and expected exchange rate are endogenous under the assumption of an open capital market with a floating exchange rate regime. External variable, domestic interest rates and an exchange rate system of equations are estimated by employing simultaneous Vector Autoregression (VAR) models that will provide dynamic relationships between the variables. The orders of the variables are the external variables, the domestic interest rates, and the exchange rates. Next, a recursive structural model is imposed as

(4.1)
$$\begin{bmatrix} 1 & 0 & 0 \\ a_{21} & 1 & 0 \\ a_{31} & a_{32} & 0 \end{bmatrix} \begin{bmatrix} {}_{t} \\ r_t \\ e_{t+1} \end{bmatrix} = \; ' + \sum_{K=1}^{K} \begin{bmatrix} {}_{t\ k} \\ r_{t\ k} \\ e_{t+1\ k} \end{bmatrix} + u_t$$

Where $_t$ denotes the variables for the domestic monetary and external financial factors, a structural VAR model consisting of five series for a vector of endogenous and semi-exogenous variables given by $_t$ = (UDCR, IR, FDTD)'. *r* is the first differentiated local 3 month interbank lending rate as a proxy of market interest rate. UDCR is the first differentiated US discount rate as a proxy for the foreign interest rate or base rate, Δe is the variation of normalized bilateral exchange rate vs. US dollar, IR is international reserves excluding gold and measured by billions of dollars, a proxy of capital openness and regime classification. FDTD is the foreign debt of the banking sector divided by total debt proxy for open-economy financing that indicates the credit and balance channel and the cost of finance from the supply side effect. IR and FDTD are measured by foreign currency hence directly involve the interest rate and the exchange rate channels.

The joint dynamics of X_t are notated by the following structural VAR:

(4.2) $A\Pi t = A(L)\ \Pi t\text{-}k + BZt\text{-}k + Ut\text{-}k$

(4.3) $\Pi t = F(L)\ \Pi t\text{-}k + GZt\text{-}k + Xt\text{-}k$

[25] Shambaugh (2003) and Borenzstein et al (2001) conclude that "the impossible trinity" is supported by his result that the more firmly floating, the smaller the reaction to changes in the U.S. interest rate.

(4.4) $E(UU')=D=A\Sigma A'$

Where *A* is a (5 x 5) matrix of structural contemporaneous coefficients, *A(L)* is a polynomial of order *p* in the lag operator *L*. *Z* is a vector of deterministic terms with associated coefficients matrix B. The vector of structural shocks $U_t=\Sigma($ *MILR, UDCR, UD, IR, FDTD* $)'$ consists of the idiosyncratic shocks associated with each of the endogenous variables that drive aggregate changes[26]. Being independent sources of fluctuations in the endogenous variables, it is assumed that the shocks are serially and mutually uncorrelated so that their variance-covariance matrix $E(UU')=D$ is diagonal. A reduced form VAR is derived from equation (4.2), where, $F(L)=A^{-1}A(L)$ and $G=A^{-1}B$. X_t is the vector of reduced form innovations of *(MILR, UDCR, UD, IR, FDTD)'*, with variance-covariance matrix, $E(X_t X_t')=\Sigma$, from equation (4.2) and (4.3): $\Pi_t=A^{-1}U_t$ or $U_t=A\Pi_t$. Equation (4.4) shows that the structural shocks U_t and their variances in *D* are related to the reduced form innovations and covariance respectively through the coefficient matrix A. Given estimation of X_t and Σ, identification of the VAR shows the recovery of the structural shocks and variances through the imposition of a sufficient number of restrictions on the A matrix to organise the instantaneous correlations among the endogenous variables. Based on the theoretical model developed by Kamin and Rogers (2000), a long run equilibrium relationship can be described as below:

(4.5) *(MILR, UDCR, UD, IR, FDTD)'*

(4.6) *MILR =* *a10+a11UDCR+a12UD-A13IR-a14FDTD+UMILR*

(4.7) *UDCR =* *a20 +UDCR*

(4.8) *UD =* *a30-a31 MILR -a32 UDCR +A33IR+a34FDTD+UD*

(4.9) *IR =* *a40+a41MILR -a42 UDCR +A43 UD +a44FDTD+UIR*

(4.10) *FDTD =* *a50-a51 MILR - a52 UDCR +A53 UD + a54IR + UFDDT*

Where in matrix notations associated with the equation (4.5), the structural shocks and the reduced from innovations are:

(4.11)

$$\begin{bmatrix} 1 & a21 & a13 & a14 & a15 \\ 21 & 1 & a23 & a24 & a25 \\ a31 & a32 & 1 & a34 & a35 \\ a41 & a42 & a43 & 1 & a45 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} MILR \\ UDCR \\ UD \\ IR \\ FDTD \end{bmatrix} + \begin{bmatrix} a10 \\ a20 \\ a30 \\ a40 \\ a50 \end{bmatrix} = \begin{bmatrix} UMILR \\ UUDCR \\ UUD \\ UIR \\ UFEDT \end{bmatrix}$$

With a0=(a10, a20, a30, a40, a50)', U_t=A $_t$ -a0 that a0 matrix consists of intercept terms is one part of the Z matrix. Ordinary Least Square (OLS) is used to assess direct relationship between both home and base country interest rates under different regimes on a simple specification of equation:

(4.12) Δ *MILR* $= a + B^*\Delta$ *UDCR*$+ u_{it}$.

(4.13) Z_t $= a + b_t +$ *MILR* - _*UDCR*$_{it}$.

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[26] See Sims (1986)

$$(4.14)\quad MILR_{it} = a + gUDCR + \_\ Bt\Delta UDCR_{t-1} + e_{it}.$$

$$(4.15)\quad \Delta\ MILR = \_(c + MILR_{it-1} - \_UDCR_{t-1}) + B\Delta UDCR + u_{it}$$

$$(4.16)\quad \Delta MILR_t = \alpha + \Sigma I{=}1{>}4\ \ i\Delta MILR_{t-1} + \Sigma I = 1{>}4\gamma i\Delta UDCR_{t-1} +$$

$$\Sigma I = 1{>}4\ \ i\Delta UD_{t-1}\Sigma I = 1{>}4\ \ i\Delta IR_{t-1} - \Sigma I{=}1{>}4\ \ i\Delta FDTD_{t-1} + \ _t$$

Z is a stationary variable where both MILR and $UDCR_{it}$ are nonstationary. To avoid loss of power in the PSS tests as well as to pertain the assumption of nonstationarity of covariating vector, the cointegrating vector, _, is prespecified at 1[27] then test whether the differential of R_i *and* R_{bi} is stationary from the residuals. *Z* in cointegration tests, nonstationarity is assumed in covariate vector and is analytically presented and then checked the stationarity of *Z* [28]. The stable movement of *R* compared to *Rb* for both the fixed and floating periods implies the reaction of policy rules for exchange rate targeting in the ACA central banks. The different regime effect is tested with open capital and inflation target variables in modified equation (4.12) into equation (4.14) $\Delta\ MILR = a + B1{*}\Delta UDCR + B2{*}\Delta UD + B3{*}IR{*}\Delta FDTD + u_{it},$ where IR or FDTD indicates no capital controls. If *B2* and *B3* are different from zero, that is evidence supporting the predictions of the trilemma. For the fixed, the *B2* coefficient is significant, but the *B1* is not, demonstrating less common shocks or the effect of fear of floating.

In the equation (4.15), the _ parameter describes the speed of the adjustment back from any deviation from the long-run relationship. This represents the short run reactions of the *B* coefficients as well as long run equilibrium under the assumption that UDCR is exogenous to MILR. Estimating equation (4.16) and interest rates appear close to unit roots in finite samples as nominal interest rates are bounded below by zero. The tests applied to both 'levels' and the 'first differences' of all the series for both fixed and floating period. The results show that the series and the gap, $(R_{\ it-1} - \_Rb_{t-1})$ is non-stationary processes, *I(1)* at the 5 percent significant level. Serial correlation seems to be absent with DW statistics being slightly above 2 for the dependant variable of CBDR(the central bank discount rate) and below 2 for MIBLR(interbank lending market rate)[29]. If the *B* coefficient and R^2 for floating are significantly greater than zero that is the evidence of fear of floating or common shocks. If the coefficient and R^2 are higher for the fixed sample that is supportive of the exchange rate constraint given by the trilemma. As can be seen from the positive and statistically significant *B* coefficients, both fixed and floating interest rates are correlated with the base country interest rate. In contrast to the trilemma theory, floating period has a higher coefficient and higher R^2 than fixed one(.46 vs. .27 and .19 vs. .01). The sample period for the post-crisis is more tightly integrated since the currency crisis in 1997-1998. When the money market rates and central bank discount rates are examined, the money market sample has higher coefficients for both fixed and floating data[30]. This may indicates a function of capital markets being more open, consequently money market rates respond more closely to the

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[27] Investment theory is necessarily held the long run correlation between R and R_b.

[28] See Zivot (2001), pp6-8 and pp13 – 15 of Eric Zivot's, "The power of single equation tests for cointegration when the cointegrating vector is prespecified" 1996, analytically presents the stationarity and alternative derivation of the asymptotic cointegration with a single prespecified cointegrating vector that is valid for weakly exogenous regressors and depends on a consistently estimable residual parameter than on values in the unit interval.

[29] Korea 3 month interbank lending rates are associated with banking sector financing source and closer relationship with US discount rate. The results imply the discount rates are monetary policy instrument rather than market rates that have higher volatility and lower mean value compared to discount rates. See the Figure 1-5 the gap and changes among interest rates, reserves and exchange rates as well as the gap, i.e. covered interest rate parity between the home and base interest rates.

[30] (B = .59, std error = .04, R^2 = .26) for money market rates and for discount rates (B = .28, std error = .08, R^2 = .01).

market and international arbitrage. Within each interest rate data sample, the gap between fixed and not is over .2 and the R^2 gap is .12 for the discount rates sample and .31 for the money market rates sample. The results hold the theoretical assumption that the shorter maturity on money market rates should have larger spreads[31] in small open economies. In all samples, the coefficients *B1, B2* and *B3* tended to be insignificantly different from zero with mixed signs. Considering the levels of the relationship based on _, and the speed of adjustment of the based on _, the larger in absolute value _ is, the faster the adjustment. The results show that compared to the fixed and the floating is more likely to reject the null of no levels relationship, have a coefficient close to one, and have a high adjustment speed, hence it is appeared to have less autonomy for the floating. The R^2 of fixed samples is generally extremely low, demonstrating that the domestic and capital factors other than the base interest rate are the significant drivers of fixed countries' monetary policies. This implies these countries may have a reasonable amount of monetary autonomy in general.

4.1 Interest Rate Volatility versus Exchange Rate Volatility

It can be considered that reducing exchange rate volatility may come at the cost of higher interest rate variability, which in turn may translate into higher variability on debt servicing costs for developing countries (Reinhart and Reinhart, 2000a). To investigate whether such a trade-off between the interest rates and exchange rates variability has occurred, we compute annual versions of the CV for these two variables for the case of the Korean won. This approach is undoubtedly simplistic and rough, but might provide an initial approximation to the issue (Figure 1.5 and 1.6). Each graph also shows an adjusted log-linear regression, its corresponding equation and the R^2 of the estimated equation is provided as below[32]. The graphs show that there is no apparent trade-off between interest rate and exchange rate variability and there exists evidence of a positive simple correlation between these two types of variability. The trade-off between these two indicators may be hidden as result of movements in other variables that influence them (see the footnote below). In the case of Korean won, the average volatility of exchange rates for the post-crisis floating regime has increased to about 14% increased compare to that of the fixed regime. On the other hand, under floating regime, the volatility of short-term interest rate has a 65% increased and a 58% for the long-term interest rates. Based on Bubula and Otker-Robe (2003), an exchange market pressure (EMP) is simply computed by letting α_1 and α_2 be the weights for exchange rate and interest rate variations, an EMP index is then calculated as a weighted average of percentage change in bilateral exchange rate and variation in percentage in the domestic interest rate, $= \alpha_1 A + \alpha_2 B$, where $A = \%\Delta e$ and $B = \Delta I$, the volatility of each weighted-component is equal and their sum equals to unity.

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[31] Taylor [1999] discusses the fact that banks are willing to commit arbitrage funds in short-term markets making covered interest parity (CIP) hold.

[32]In the logarithm of equation for the volatility of interest rate that is explained by the volatility of exchange rate has a weak explanatory power under both regimes, indicates the omission of other variables in the equation as well as the uip for those developing countries are usually do not hold for less developed bond market and transitory risk premium or credit rating. However, the equations and R squares show the post-crisis floating regime for the relationship in volatility between the short term interest rate and exchange rate has largest among others ((i) y = 0.0043Ln(x) + 0.0389 (pre-crisis with short term interest rate with fx volatility), R2 = 0.0333; (ii)y = -0.0241Ln(x) + 0.0261 (pre-crisis, with long term interest rate and fx volatility), R2 = 0.0573; (iii)y = 0.0103Ln(x) + 0.0685 (post crisis with short term interest rate with FX volatility), R2 = 0.1007; and (iv)y = 0.0486Ln(x) + 0.3114 (post-crisis with long term interest rate volatility), R2 = 0.0553).

$$\beta_1^{2^*} \operatorname{var}(A) = \beta_2^{2^*} \operatorname{var}(B),\ \beta_1 + \beta_2 = 1,\ \textit{rearranging gives}$$

$$\beta_1 = \frac{\sigma_B}{\sigma_A + \sigma_B}, and\, \beta_2 = \frac{\sigma_A}{\sigma_A + \sigma_B} \qquad (4.17)$$

Where σ_A and σ_B are the sample standard deviations of A and B. the EMP is rewritten as

$$\iota = \frac{\sigma_B}{\sigma_A + \sigma_B} * A + \frac{\sigma_A}{\sigma_A + \sigma_B} * B, or\, \iota = A + \frac{\sigma_A}{\sigma_B} * B \qquad (4.18)$$

The is therefore increases in both the depreciation of the exchange rate and a rise in interest rate changes. The broad indicators of the exchange market pressure (EMP) for the post-crisis samples exceed the pre-crisis threshold value. We view that the ACA3 may accordingly need some attention on risk management in the exposures of volatility in exchange rates and interest rates[33].

4.2 The Economic Results under New Exchange Rate Regimes

The macroeconomic performances of countries in terms of growth, investment, trade, and inflation, suggest a strong link between the choice of the exchange rate regime and macroeconomic performance. The econometric results by Levy-Yeyati and Sturzenegger (LYS, 2002) and Ghosh et all (1997, 2000) observe that the float regime appears to achieve the highest growth over the sample, while pegs yield the best inflationary performances. While, intermediate regimes, with the managed float and the crawling peg, yield higher GDP growth and are associated with intermediate results in terms of inflation (Williamson, 2000).

We provide evidence as to the new exchange rate regime has an impact on economic results, in particular on inflation and growth. A simplified Phillips curve (PC) has been adopted with sets of assumptions based on a stochastic dynamic Mundell-Fleming model (Appendix 3)[34]. To collaborate with the regime effect on the Phillips curve equation, it can be rewritten as

(4.1) $\pi t fix - \pi_t = -(1/\sigma+\eta)u_t$, where $\pi_t fix = Et(p_{t+1} - p_t) =$

$\pi_t - (1-\theta/\eta)(y_t - d_t)$ with the slope of PC, $fix =$

$- d(\pi_t - \pi_t)/du_t \mid fix = 1/\sigma+\eta$.

(4.2) $\pi_t fl - \pi_t = -(\lambda+\sigma+\eta/(1+\lambda)(\sigma+\eta))u_t$, *where* $\pi_t fl =$

$Et(p_{t+1}\, p_t) = Et1(p_{t+1}\, p_t) + (1-\theta)(m_t - y_t) = \pi_t + (1+\theta)(m_t - y_t)$,

where the slop of PC, $fl = - d(\pi_t - \pi_t)/du_t \mid flx =$

$\lambda+\sigma+\eta/(1+\lambda)(\sigma+\eta) = (1/\sigma+\eta)(1+\sigma+\eta-1/1+\lambda)$.

Where $\pi_t = 0$. The slope of the Phillips curve is decreasing as the aggregate demand elasticity, σ and η, increases, and $\pi_t = m^g = y^g$, u_t equals $y_t - y_t$, the output gap. The equation (4.1) indicates that the Phillips curve is flatter with the aggregate demand elasticity of the two parameters, σ to the domestic real rate of interest rate, and η to the real exchange rate. As

[33] EMPI, EMPK, and EMPT values 2842.940001, 1039.869862, 28.71375337 for the pre-crisis samples; 10893.41009, 1898.1116, 48.09681071 for the post-crisis sample. EMP of Indonesia, Korea, and Thailand shows an increased EMP for the period of post-crisis (1999-2004) to 3.8, 1.8, and 1.7 times more than those of the pre-crisis.

[34] A linear relation is also assumed as an approximation for possible nonlinear functions.

flexible exchange rates are assumed to provide an adjustment of prices in buffering the real economy against external shocks, a change in unanticipated inflation would be associated with a smaller change in the output gap, i.e., a steeper Phillips curve unless the elasticity of aggregate demand to the real interest rate and the real exchange rate are large (Loungani, Razin, and Yuen (2002)).

Time-variations in the expected inflation and natural output growth rates can shift the expectations-augmented Philips curve. The expectation-augmented Phillips curve is written as

(4.4) $\pi_t = \ 0 + \ 1u_t + \ _t$, where $\pi_t - \pi_t\hat{}= - \ u_t$.

(4.5) $\pi_t = \pi''_t + ax_t + e$, where $x = -(ak/\phi)(\pi-\pi'') + u_t$.

or $\pi = \pi'' - \alpha(x-u)$ with $x=x0 - b(i-\pi) + u$.

Where $\beta 0$, the intercept equals π_t', and β1, the slope, equals ϕ. ε is the residual errors. It is assumed that the changing inflationary expectations and the changing tradeoff between inflation and output under different exchange rate arrangements, $\beta 0$ and $\beta 1$ will vary across exchange rate regimes. Where π_t and $\pi\hat{}$ are the actual and expected rates of inflation at time *t*, u_t [35]the output gap at time *t*, and ϕ $(\phi>0)$ is the slope of the Phillips curve. In theory, the null hypothesis is that $\beta 1flx>\beta 1fix$ and $\beta 0flx>\beta 0fix$. x equals the percent output gap, $((y-y^n)/y^n)$, and ($''$) indicates a target inflation. $\alpha=\phi/ak$ and ϕ/a describe the trade-off between inflation variability and the volatility of the output gap. If the relative weight, ϕ of the output goal in the loss function of the central bank increases, the variance of the output gap decreases, whereas the variance of the inflation rate increases. The marginal cost of fluctuations in output is ϕx and the marginal cost of fluctuations in inflation is $k(\pi-\pi'')$. ϕ (k) is a measure of the cost of fluctuations in output for the central bank. If $x<0$, the economy operates below the natural level of output, a slight increase in x and δx, leads to $-\lambda x\delta x$. The effect of inflation is $a\delta x$, and the resulting cost is $ak(\pi-\pi\hat{})\delta x$. Equalizing marginal cost and marginal utility of monetary policy actions gives, $x = -(ak/\phi)(\pi-\pi'')$.

Monetary policy rules (Taylor 1993) specify the policy instrument, the nominal interest rate, as a function of the inflation rate and the output gap. As the inflation rate and the output gap are endogenous variables, when solving the Phillips curve equation and the monetary policy rule for output gap x, when substituting the result for the output gap, $-b(i = \pi'-r^*) + u$, a short-run equilibrium is solved as $i = r^* + \pi' - (\pi-\pi'-e)/b(a+\alpha) = i + (1+(1/(b(a+\alpha)))(\pi'-\pi) + e/b(a+\alpha)$, and the long run equilibrium nominal interest rate ($i\hat{}$) is ($r^* + \pi\hat{}$). This implies a positive inflation shock that reduces the output gap.

The monthly data series for production index, consumer price index (CPI), inflation, real effective exchange rate (REER), central bank discount interest rate, 3 month bank lending rate are obtained from the Asian Development Bank (ADB) Database. The production index is used for a proxy of output, and both CPI and Inflation rate (INF) are used for the price level with the base year equals to 1996. The transitional period (1997M4–1999M12) which may be contaminated by the mixture of the two exchange rate regimes, is therefore dropped from the whole sample periods (1990M1-2004M3). Unit root tests using Dickey Fuller regressions

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[35] The excess capacity, u_t is estimated from the regression: $ln(y_t) = \alpha 0 + \alpha 1t + u_t$, where α0 and α1 are the estimates of $ln(y0)$ and g. u_t is therefore measured by the logarithm of detrended output.

include an intercept with and without a linear trend for all variables and residuals for regressions rejected the null hypothesis of unit root at 95% critical value for the ADF statistic. The average level and variability measured by both standard deviations and coefficients of variations of inflation, output, real effective exchange rate (REER), and consumer price index (CPI) of the ACA3 are summarized in Table 1.4 (appendix 1). The inflation is more volatile under the flexible exchange rate regime than that of the fixed rate period while output volatility is about the same in both periods in all ACA samples. In case of CPI, the index shows a lower level during the fixed period while the variability is higher for the same period. In the Phillips curve framework, a weak explanatory power and the cross correlation plots show no clear relationship in the ACA3 for both fixed and flexible periods (Table 1.9 and 1.10). The simple plots of inflation between these innovations are negative autocorrelation that indicates the information implies the 'inflation targeting' monetary policy in level. Based on the stochastic dynamic Mundell-Fleming model adopted, the preliminary empirical findings based on the ACA3 data support some evidence for the theory that the slope coefficient of the Phillips curve under the floating rate regime has a bigger intercept and slope than its fixed rate counterpart and the cost of output expansion in terms of unanticipated inflation is found to be smaller under the fixed rate system.

5. Conclusion and Policy Consideration

This paper provides some preliminary evidence on the important issues on the newly floated exchange regimes in the countries severely affected by the Asian Currency Crisis in 1997, Indonesia, Korea and Thailand. We focus on the *de facto* variation in exchange rates and the impact of new currency regimes on monetary autonomy and economic performance. By utilizing the filtering methodology for nonstationary sequences of exchange rates and the simultaneously expanded conditional variance estimation with the GARCH class models, we confirm that the variation of exchange rates has been increased to a similar level as the floating standard. Next, by assessing the "impossible trinity" and the slope of the Philips curve, we broadly conclude that the open-economy capital and financial impacts are significantly more than the regime impact. Empirical evidence suggests that fixed regimes eventually collapse due to the EMP, overvaluation, or speculative attack under liberalized capital flows (Obsfeld and Rogoff (1995)), while Reinhart (2000) says that floating regimes do not exist. Williamson (2000) therefore recommends making fixed exchange rate regimes more flexible by introducing soft crawling bands pegged to currency baskets. However, intermediate regimes (the soft pegs and managed floating regimes) have been more crisis prone than both hard pegs and floating regimes (Bubula and Rtker-Robe, 2003). The IMF recently has turned from "corner" solutions[36] to its current doctrine of floating arrangements with inflation targeting (Rogoff, 1999), IMF, Fischer, 2001).

Considering long term policy implications under a floating regime, a monetary policy may be set based on the trinity: (1) a flexible exchange rate (Obstfeld and Rogoff (1996)), (2) an inflation target (Bernanke et al (1999)), and (3) a monetary policy rule (Taylor (1999)). As the exchange rate and the interest rate are interdependent under a floating regime, the responsibilities for exchange rate policy and interest rate policy may not be divided, hence, in the short- and medium term, we may recommend a 'Floating Regime with Macro-Financial-

[36] For the evidence based on Markov chains of regime transitions against the "bi-polar view", see A. Bubula and I. Otker-Robe (2000), the evolution of exchange rate regimes since 1990: evidence form de facto policies.

Micro Risk Management' on the exposure of both exchange rates and interest rates for the small-open-capital-Asian countries, rather than adopting a simple 'inflation targeting'.

Further research based on macro-financial- microstructural framework is essential for providing a theoretical foundation to facilitate the development of Asian financial markets and the extension with international money and capital markets. An important question, whether the effectiveness of monetary and fiscal policy is valid regardless of exchange regimes, once a small open economy encompasses a developed financial system, may need some attention for further studies. These are the subjects of my ongoing research.

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APPENDIX 1 Exchange Rate Classification, Descriptive statistics, and Results

Table 1.1 Exchange Rate Regime Changes since the Asian Currency Crisis

Changes of Exchange Rate System: **(Year of Floating)**

Independent float	Fixed (Managed Float)	Currency Board
1991	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
92	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
93	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
94	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
95	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
96	Indonesia/Korea/Thailand/Malaysia/China	Hong Kong
97 Korea	Indonesia/Thailand/Malaysia/China	Hong Kong
98 Indonesia/Korea	Thailand/Malaysia/China	Hong Kong
99Indonesia/Korea/Thailand	Malaysia/China	Hong Kong

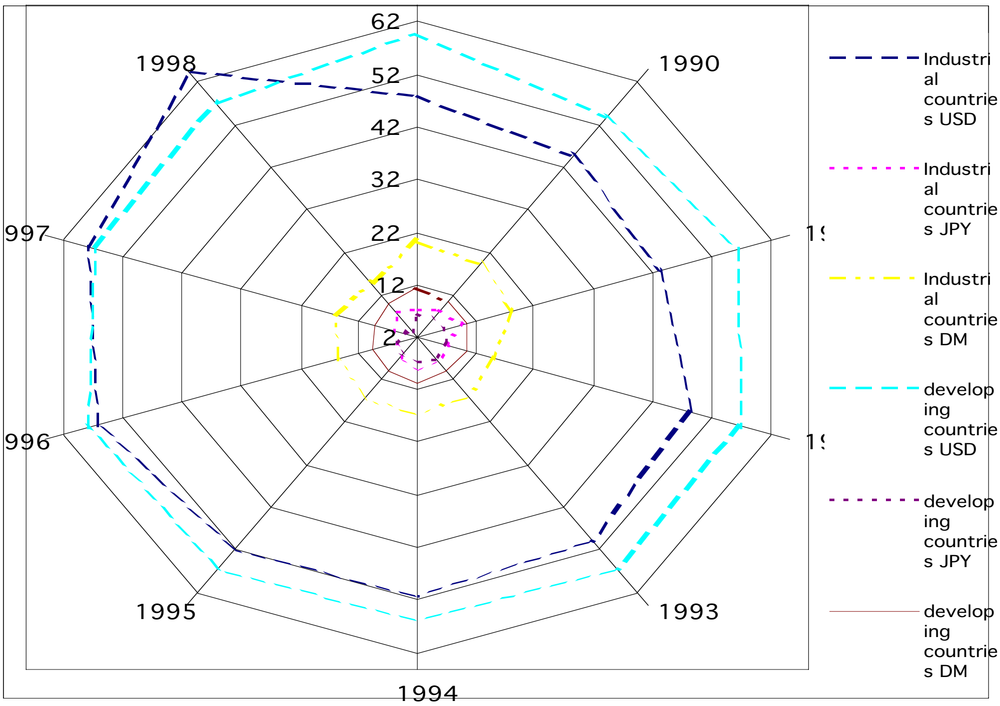
Source: Exchange rate regimes are classified according to the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions; other various resources also are referenced. Re-categorized anything other than an 'independent float' or a currency board as a managed float or fixed regime including exchange rate regimes, pegs, crawling pegs, crawling bands.

Table 1.2 Date Indicating the Switch towards a More Liberal Exchange Regime

Asian Currency Crisis Affected Countries (ACA5)	DOMESTIC	EXTERNAL	IMF Classification
INDONESIA	1983	1985	Float 1997
MALAYSIA	1978-85, 1987	1987	
PHILIPPINES	1981	1992	Float 1997
KOREA	1991	1991	Float 1997
THAILAND	1989	1991	Float 1997

Source: IMF, various including the central bank of the ACA countries.

Figure 1: Currency Composition: Total Identified Official Holdings of Foreign Exchange



Source: Data are from Eichengreen and Mathieson, WP/00131 and 1999 IMF annual report, compiled by the author.

Table 1.3 Descriptive Statistics and Contemporaneous Correlations

Pre-crisis-fixed regime	INR	KRW	THB	MYL	PHP
Mean	2161.1625	26.345474	25.836845	785.7818	2.6174976
SE	29.304891	0.2020999	0.2319197	5.7507617	0.015339
Median	2103.5	26.18935	25.35522	787.14286	2.56539
SD	285.62875	1.9698263	2.2604734	56.051492	0.1495057
SV	81583.782	3.8802156	5.1097398	3141.7697	0.022352
Kurtosis	9.8282531	5.7527836	21.635059	3.0866816	10.065754
Skewness	2.5184189	1.4539952	4.6181917	1.1120471	2.5929698
Range	1770.4428	12.11597	14.4766	345.30186	0.93914
Min	1803.7311	22.35678	24.5419	683.50814	2.44172
Max	3574.1739	34.47275	39.0185	1028.81	3.38086
Confidence Level(95.0%)	58.185547	0.4012741	0.4604819	11.418272	0.0304559
Post-crisis-floating regime	INR	KRW	THB	MYL	PHP
Mean	8885.561	49.268655	41.248322	1201.5093	3.7999318
SE	115.36499	0.7292697	0.2984618	7.3755259	1.554E-05
Median	8922.1429	51.2875	41.49955	1191.595	3.8
SD	972.08272	6.1449358	2.5148837	62.147286	0.0001309
SV	944944.81	37.760236	6.32464	3862.2852	1.715E-08
Kurtosis	0.4861109	-0.9921878	-1.0545957	-0.7095924	0.4302981
Skewness	0.3487659	-0.6916927	-0.1692205	0.4494054	-1.2805098
Range	4546.5281	18.505	9.02983	239.73087	0.00047
Min	6767.0909	37.82595	36.59381	1086.3182	3.79961
Max	11313.619	56.33095	45.62364	1326.0491	3.80008
Confidence Level(95.0%)	230.08822	1.4544826	0.5952632	14.710022	3.099E-05

Source: Data are obtained from Asian development bank, ADB data base for Indonesia, Korea and Thailand.

Tale 1.4 Correlation among Policy Variables: Fixed Pre-crisis and Floating Post-crisis period.

PRE-CRISIS	INR	ICPI	IPI	IGIR	IGIRM	POST-CRISIS	INR	ICPI	IPI	IGIR	IGIRM
ICPI	-0.848	1.000				ICPI	-0.911	1.000			
IPI	-0.642	0.836	1.000			IPI	0.102	0.088	1.000		
IGIR	-0.407	0.448	0.137	1.000		IGIR	0.000	-0.270	-0.363	1.000	
IGIRM	-0.831	0.948	0.732	0.674	1.000	IGIRM	-0.824	0.846	0.123	-0.101	1.000
ICBR	0.546	-0.683	-0.579	-0.342	-0.606	ICBR	-0.806	0.627	-0.207	0.060	0.557

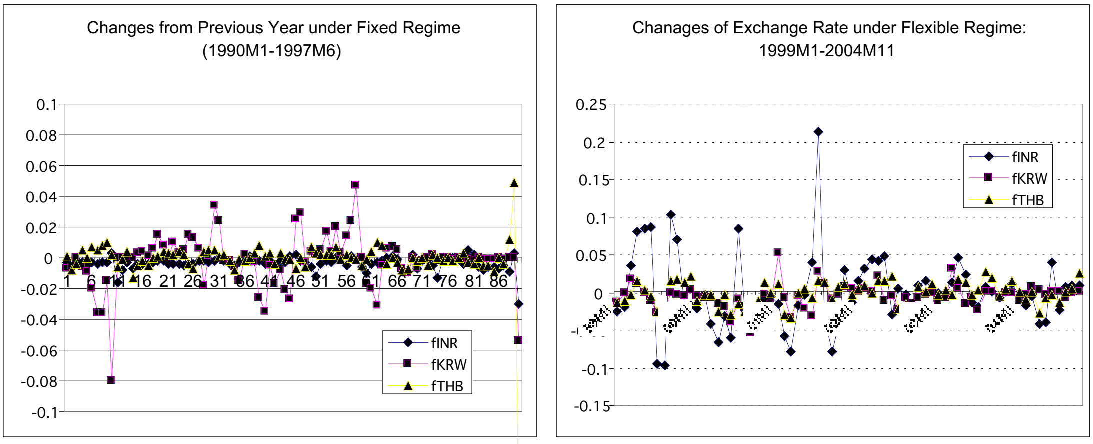
PRE-CRISIS	KRW	KCPI	KGDP	KGIRM	KGDP	KGIR	KPI
KCPI	-0.883	1.000					
KGDP	0.312	-0.406	1.000				
KGIRM	-0.445	0.537	0.504	1.000			
KGDP	-0.652	0.994	-0.332	-0.704	1.000		
KGIR	-0.744	0.922	0.329	0.626	0.406	1.000	
KPI	-0.824	0.969	-0.184	0.484	0.893	0.956	1.000
KCBR	0.591	-0.629	0.252	-0.246	-0.602	-0.630	-0.635
POST-CRISIS	KRW	KCPI	KGDP	KGIRM	KGDP	KGIR	KPI
KCPI	-0.855	1.000					
KGDP	0.359	-0.416	1.000				
KGIRM	-0.898	0.914	-0.222	1.000			
KGDP	-0.858	0.978	-0.412	0.920	1.000		
KGIR	-0.758	0.968	-0.407	0.881	0.959	1.000	
KPI	-0.739	0.920	-0.444	0.848	0.954	0.956	1.000
KCBR	0.645	-0.846	0.312	-0.751	-0.778	-0.856	-0.753

PRE-CRISIS	THB	TCBR	TCPI	TGDP	TGDP	TGIR
TCBR	0.104	1.000				
TCPI	0.415	-0.102	1.000			
TGDP	0.051	0.468	-0.470	1.000		
TGDP	0.428	-0.570	0.920	-0.348	1.000	

TGIR	0.456	-0.109	0.931	-0.310	0.880	1.000
TPI	0.419	-0.092	0.949	-0.417	0.914	0.942
POST-CRISIS	THB	TCBR	TCPI	TGDP	TGDP	TGIR
TCBR	-0.042	1.000				
TCPI	-0.755	-0.323	1.000			
TGDP	-0.279	-0.075	0.120	1.000		
TGDP	-0.751	-0.400	0.911	0.244	1.000	
TGIR	-0.765	-0.427	0.916	0.230	0.923	1.000
TPI	-0.778	-0.416	0.957	0.192	0.969	0.962

Source: Data from ADB data base, pre crisis period covers 1990Q1-1997Q3; post crisis sample covers 1991Q1-2004Q3, THB, TCBR, TCPI, TGDP*,TGDP, and TGIR represents Thailand central bank discount rate, consumer price index, gross domestic production, and gross international foreign currency reserves respectively.

Figure 1.2 The Variations of Exchange Rates: sub-sample (pre-and post-crisis) periods



Source: Data are obtained from Asian development bank, ADB database for Indonesia, Korea and Thailand.

Figure 1.3a Standard Deviation of Growth rate of Real Exchange Rate: 1971-2003

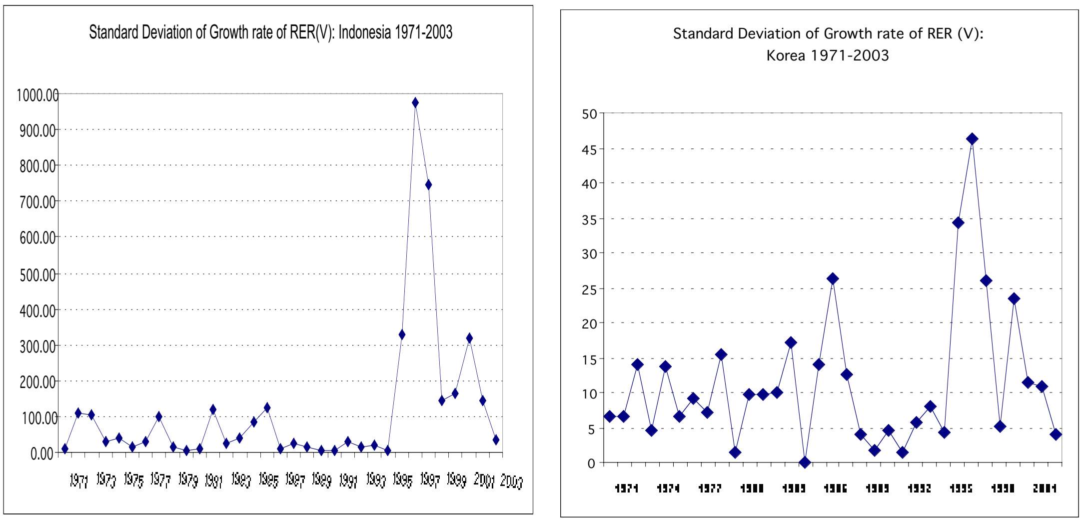
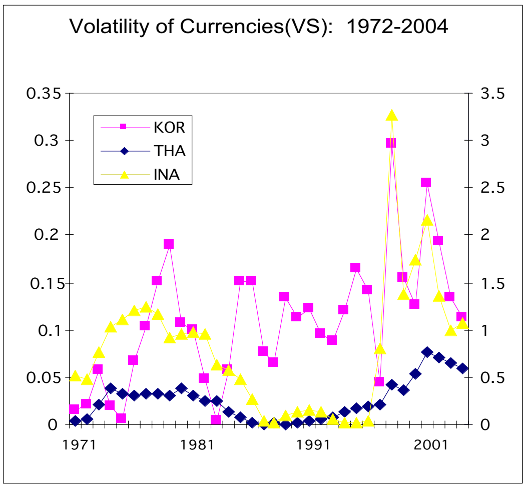


Figure 1.3b Standard Deviation of Δ % of Real Exchange Rate: 1971-2003, Indonesia



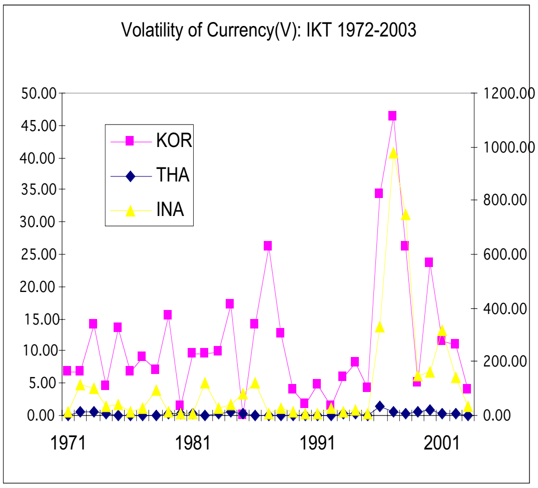


Figure 1.4 Coefficient of Variation of Real Exchange Rate (CV)

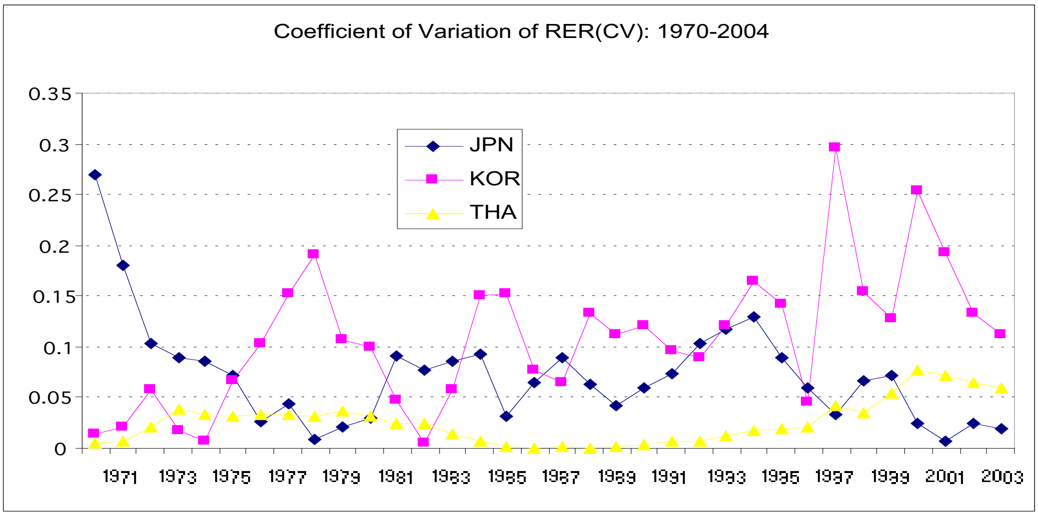
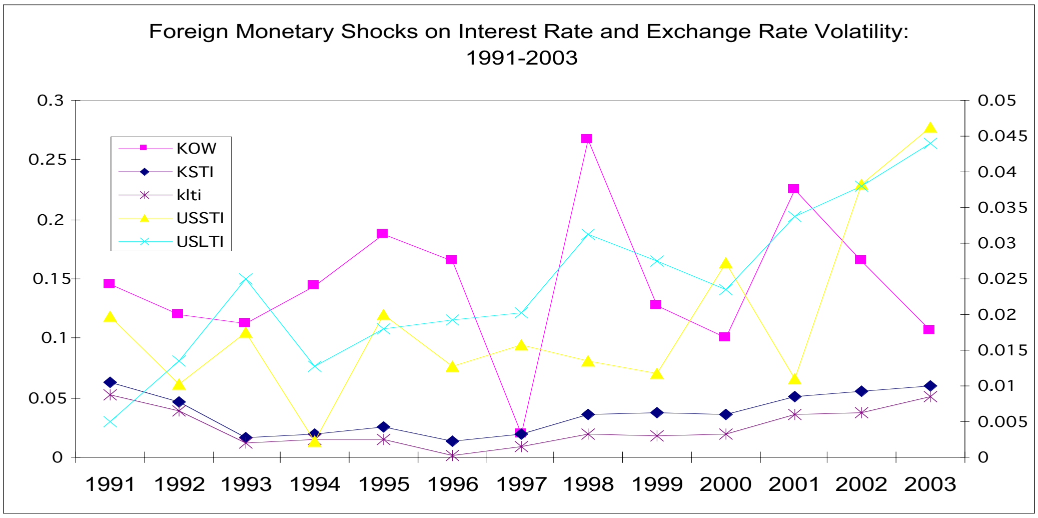


Figure 1.5 Interest Rates and Exchange Rate Volatility under the different Exchange Regime



Source: Data are obtained from the Asian development bank, ADB database for Korea, and the US data is obtained from the US Department of Commerce, Bureau of Economic Analysis

Figure 1.6 Exchange Rate Market Pressure Index (1993-2003):

the variations of exchange rate and interest rates: Case-Korea: AER(left), REER(Right)

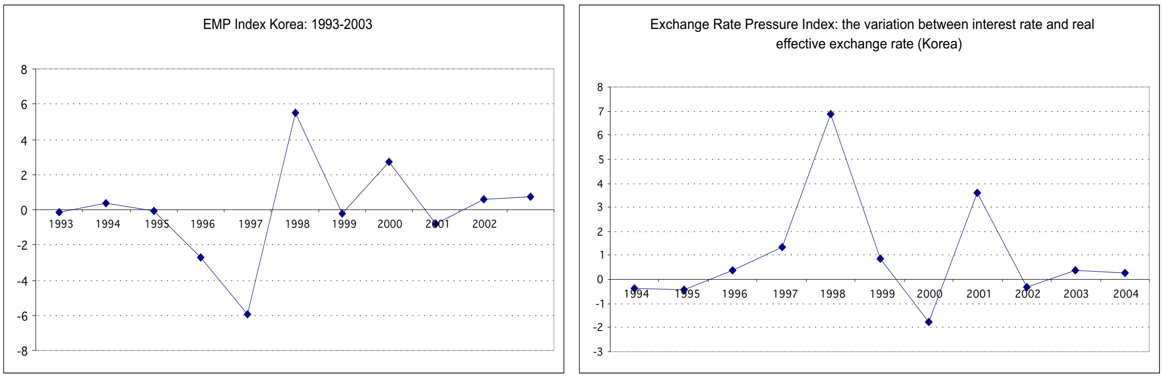
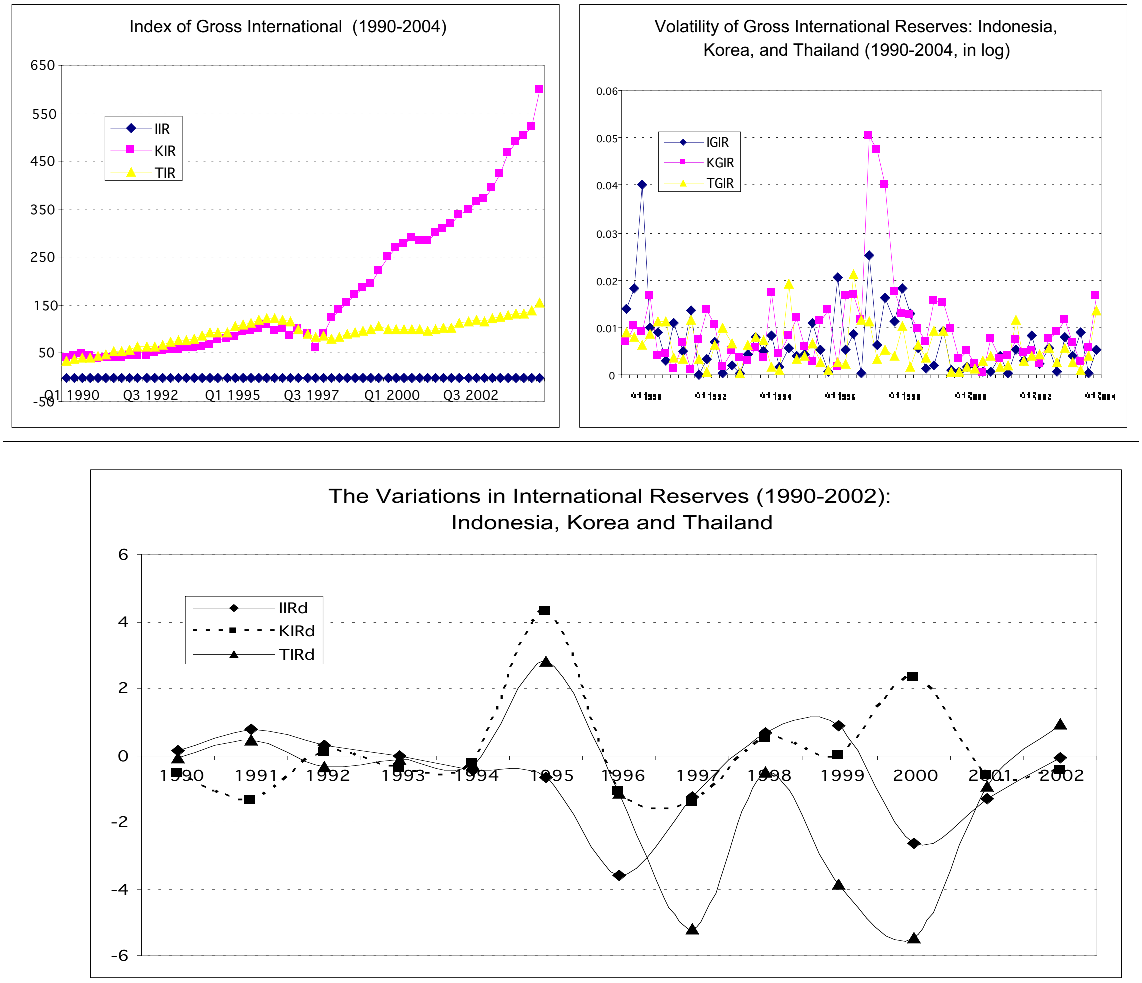


Figure 1.6 Gross International Reserves: Proxy of FX Market Intervention (90Q1-04Q4)



Source: Data are obtained from the ADB, IMF, IFS. Note. The top left and right panel: The standard deviation of (V)for the pre-crisis period for Indonesia, Korea, and Thailand is 1.29045, 2.948, and 1.53405 respectively, and for the post-crisis period, the standard deviation for the three countries is 2.21358, 9.6905, and 3.24085 respectively. The centre panel: Standard deviation of (VS) for the period of fixed regime for Indonesia, Korea, and Thailand is 0.008549, 0.009137, and 0.005065; Standard deviation for the period of flexible regime for Indonesia, Korea, and Thailand is 0.268346, 0.312371, and 0.276682. The standard deviation of the level of official reserves for the pre/post crisis period for Indonesia, Korea and Thai is 3483.018/1976.148; 8153.655/30499.23 and 9256.388/4260.284 respectively.

Table 1.5 Unit Root Tests

The table reports Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests without a time trend for volatility series. The lag length for the unit root tests is decided based on the Schwarz information criterion. The critical value for the tests at the 1% significance level is -3.453. It displays the 99 per cent critical values, using the results in MacKinnon (1991), *DX(t) = a0 + a1*t + a2*X(t-1)+b1*DX(t-1) + bp*DX(t-p), where DX(t) = X(t) - X(t-1).* All data are first differentiated after the logarithmic is replaced.

	ADF	p-value	PP	p-value
IDR	-3.672	0.005	-3.724	0.004
KRW	-3.918	0.000	-5.204	0.000
THB (4th lag)	-3.504	0.000	-10.466	0.000

Table 1.6 Summary Statistics of the VAR model

Fixed regime	Δe^k_t	Δi_t	Δui_t
C	-.089(.204)	-.044 (.0662)	-.070(.712)
Δui_{t-1}	.180(1.127)	.095***(2.601)	.136*(1.797)
Δui_{t-2}	-.041(.120)	.004(.259)	-.023(.296)
Δi_{t-1}	-.647***(9.118)	.127**(2.448)	.007(.462)
Δi_{t-2}	-.328***(.194)	.018(.992)	-.009(.532)
Δe^k_{t-1}	.313(.432)	.236***(3.105)	.025(.228)
Δe^k_{t-2}	-.083(.312)	.051(0.697)	-.125(1.167)
R squared	.341	.136	.008
D.W.	2.155	2.010	2.001
Floating regime	Δe^k_t	Δi_t	Δui_t
C	-.0914 (.891)	-.292(.36)	-.050(.412)
Δuit-1	.129***(4.601)	.215(3.127)	.197*(1.294)
Δuit-2	-.009(.153)	-.0181(.890)	-.039(.191)
Δit-1	-.027**(2.448)	-.214***(9.118)	.001(.268)
Δit-2	.011(.992)	-.032***(.194)	-.003(.312)
Δekt-1	.136***(5.105)	.213(.432)	.005(.268)
Δekt-2	-.021(0.863)	-.148(.715)	-.095(1.117)
R squared	.292	.358	.004
D.W.	2.159	2.017	2.013

Notes: Δekt, Δit, and Δuit denotes exchange rate, domestic interest rate, and US interest rate, respectively. Parenthesis indicates t statistics, *, **, and *** is 10%, 5%, and 1% significance level.

Table 1.7 Testing the Trilemma (Adjustment speeds to shocks in base international rate)

Variable	β1 UDi	β2, IR	β3, MILR	β4, FDTD	R^2
Fixed R.	.13 (.26)	.28(.13)	.1(.12)	.14(.15)	.09
Flexible R.	.28(.34)	.42(.26)	.29(.12)	.22(.36)	.23
Adjustment speeds to shocks in base international rate					
1-3 months; 4-12 months; 1-12 months					
Fixed R.	13	28	9		
Flexible R.	32	47	11		

Table 1.8 Domestic Interest Rate Volatility

(External Variable: US Interest Rate Volatility)

Variables	Indonesia	Korea	Thailand
vUS t-1	-401.93 (-1.38)	64.95(-9.58)	118.19 (3.27)
D t	-0.917 (-9.47)	9.268(-13.79)	-3.10 (-39.63)
D_vUS t-1	235.81 (0.77)	-1138.16(-1.8)	-88.22 (-2.44)
Constant	-0.556 (-5.72)	-23.4(-38.67)	-3.86 (-52.39)
_t-1	1.27 (21.00)	1.57(-128.31)	0.25 (30.62)
_t-2	-	-	-
ht-1	0.45 (41.68)	1.76(-933.42)	0.83 (236.80)
Wald _2	1014.37	2153.58	102.50
Log-Likelihood	-4481.3	807.8	-479.68

Table 1.9 The Phillips Curve in Indonesia, Korea, and Thailand

	Fixed ER (1990M1-97M3)	Restricted regime change(90-97)	Flexible ER(2000M1-2004M3)	Restricted regime change(00-04)
Indonesia				
Intercept	-9.9(5.89)	-10.68(6.08)	-2.34(7.60)	2.27(10.15)
Slope	2.2 (1.34)	1.85(1.41)	.59(1.73)	-.21 (2.16)
Korea				
Intercept	4.90(2.50)	2.14(3.49)	11.15(9.40)	17.31(13.13)
Slope	-1.38(.58)*	-1.61(.61)*	-2.49(1.89)	-3.9(3.46)
Thailand				
Intercept	-1.96(6.81)	3.36(2.85)	-3.56(7.28)	-1.13(8.70)
Slope	.21(1.52)	-.05(.039)	.44(1.54)	-.079(1.7)

Note: standard errors of the parameter estimates are given in parentheses. * indicates the coefficient has statistically significant t-ratio and p value at 95% confidence interval

Table 1.10 Cointegration estimation on Inflation and Output:

Sample Fixed (1990:M1-1997M3); Flexible (2000M1-2004M3) period

	Indonesia		Korea		Thailand	
On Inf	Fixed Er.	Flex. Er.	Fixed Er.	Flex. Er.	Fixed Er.	Flex. Er.
C(-1)	.34	-.008	.22	-.3E7	-1.7	.054
Slop(-1)	-.007(.02) -.26(.79)	0(.003) 300(.00)	-.18(.035) -5.21(.00)	-.24(.02) -8.5(.00)	.27(.04) -5.9(.0)	-.29(.05) -5.47(.00)
R squared	.12(.015)	-.45	.059(.024)	.01 (.46)	.028(.24)	.7E-4(.95)
Restricted						
C(-1)	.33	-.23(.08)	.16	.34(.08)	-.21	-.04(.14)
Slope(-1)	-.65(.64) (.315)	1.60(.27) 5.7(.00)	-1.27(.58) (.033)	-.09(.08) -1.1(.25) (.033)	1.2(.85)1.4(.18) .000	.15(.24) .6(.53)
REER	.031(.31)	-.07(.01)	.047(.12)	.007(.004)	.007(.04)	-.007(.01)
CHSQ*	1.06(.30)	12.9(.00)	3.45(.067)	.12 (.00)	2.9(.08)	.008(.82)
On Output	Fixed Er.	Flex. Er.	Fixed Er.	Flex. Er.	Fixed Er.	Flex. Er.
C(-1)	-.02(.01)	-.008(.01)	.21E3(.004)	-.32E-9	.006(.004)	.66E-3
slope(-1)	.99(.002) 432(.00)	1.0(.003)	1.001(.0012)(.000)	1.0(.000) .000	4(.001) 1.4(.16)	1.0(.002) 637((.00)
R sqrd.	.41(.00)	-.45(.00)	.95(.00)	.70(.00)	.88(.00)	.90(.00)
Restrct.						
C(-1)	-.022(.01)	.97(.06)	-.001	.97(.04)	.006(.004)	.98(.02)
Slope(-1)	.97(.055) 17.5(.00)	-.004(.0) .6(.52)	.96(.02) 44.9(.00)	.009(.01) (.57)	.99(.03) 33.9(.00)	-.01(.01) -.93(.35)
REER	432(.00)	-.001(.003)	.001(.92E-3)(.076)	.002(.001)	.52E-3(.001)	.81E-3
CHSQ*	.001(.002)	.37 (.00)	3.28(.070)	.69(.000)	.13(.71)	.89(.00)

Note: * indicates Likelihood ratio statistic of joint test of zero restrictions on the coefficients of additional variables. Standard errors are in (), while square bracket refers p-value at 95 percent confidence intervals.

APPENDIX 2 The Results of GARCH Applications on the ACA3 Currencies

A2-I

Spot rates (1529)	? 0	? 1	? 0	? 1(esq)	? 1(h-sq)	h^2	? 1-meu**	Log-L	? *	R^2
KRW										
GARCH	-.12692	.027656	.44681	.07767	.90694			-4532.2	18623	.6993
	(-1.1733)	(1.0016)	(.14136)	(.14136)	(.011960)					
GARCH -M	-.24317	.026647	.44802	.077223	.90724	.006165		-4531.9	18585.5	.9648
	(-1.2441)	(.96417)	(.14178)	(.011911)	(.013530)	(.71661)				
EGARCH	-.069819	.03929	.14331	.18641	.95868		5.1363	-4538.9	14555.6	.8350
	(-.96747)	(2.2093)	(.038173)	(.013536)	(.011663)		(.31459)			
IDR (478)										
GARCH	.87358	.028068	190.3151	.19679	.71166(.057154)			-2410.6	677.1866	.0014393
	(.54551)	(.67201)	(49.8493)	(.052397)						
GARCH -M	2.1166	.027756	192.5288	.20114	.70721	-.9595E-3		-2410.5	671.3494	.0015071
	(.68612)	(.66594)	(50.5458)	(.054376)	(.058601)					
EGARCH	.27809	.010627	1.3	-.053341	.82866		.39667	-2413.8	671.3494	.9436E-3
	(.16690)	(.17713)	(.31396)	(.042250)	(.042370)		(.073270)			
THB										
GARCH	-.0024910	,967498	.0011385	.20100	.75319			1011.3	3545.8	-.0016698
	(1.88499)	(2.6960)	(.2282E-3)	(.027450)	(.029991)					
GARCH -M	-.0048386	.077372	.0011283	.20041	.75427	.17654		1011.6	3564.6	-.0016097
	(-1.1295)	(2.6580)	(.2280E-3)	(.027426)	(.030022)	(.72390)				
EGARCH	-.0055142	.077458	-.26337	.043616	.93167		.31449	1010.7	10528.3	-0.0034785
	(-1.9899)	(2.7475)	(.048207)	(0.018067)	(.011722)		(.032115)			

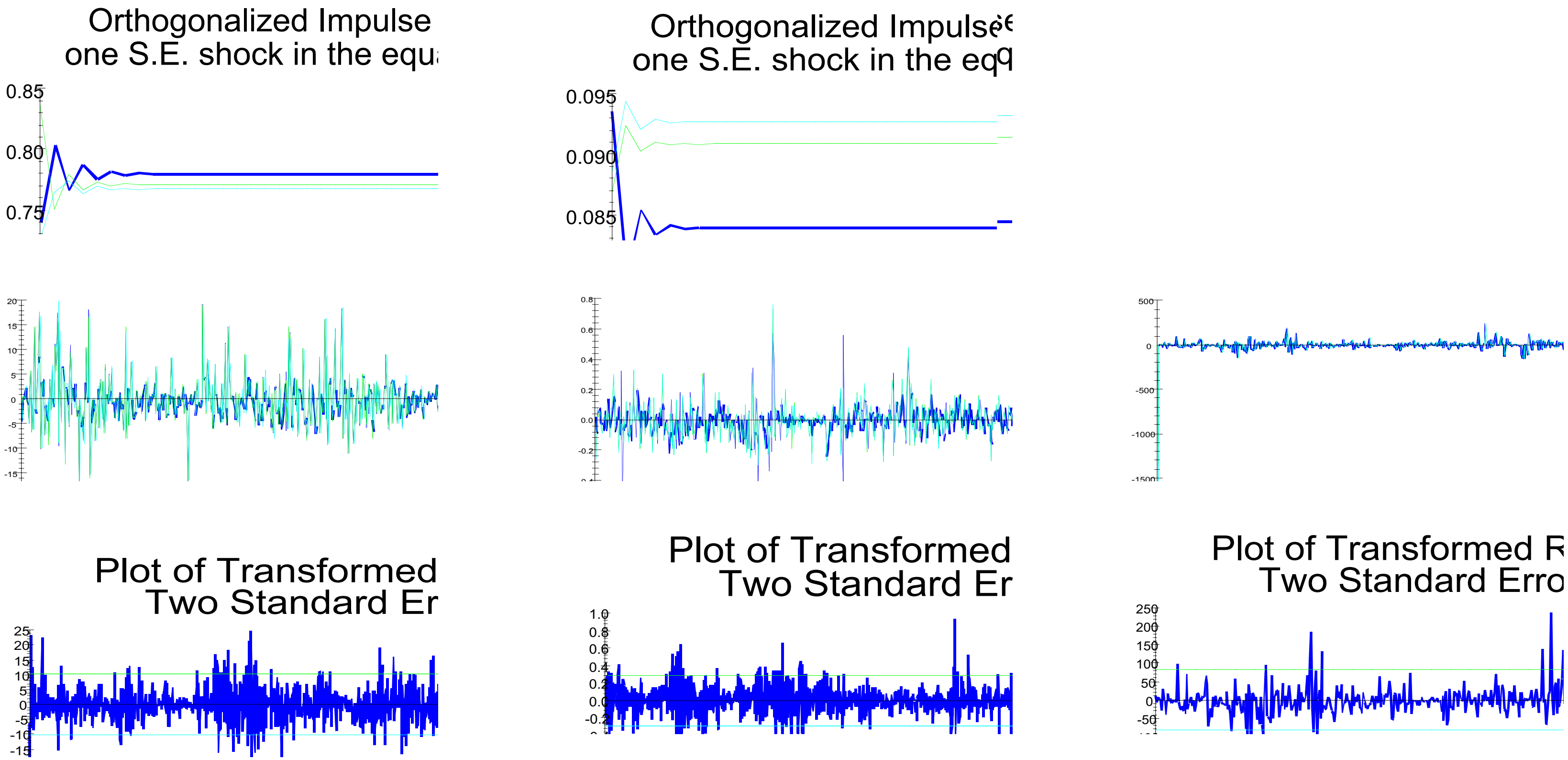
Notes: In-sample Likelihood Ratio, Test Results: GARCH Models for Volatility, t-ratio and asymptotic i.e. are parentheses, Log-L denotes the maximized value of the log-likelihood function in each case. *$\chi 2(2)$ and $\chi 2(3)$, **MEU equals SQRT(2/3.14159)=.79788, stands for the expectation of the absolute value of the disturbance term. All test statistics are estimated converged after 25 to 67 iterations.

Forwards (474)	0	1	β0	β1	β1	h^2	β1-meu**	Log-L	R^2	χ
DKRW9403										
GARCH	-.28042	-.063277	.48391	.049147	.92919			-1387.9	.0042873	9159.6
	(-1.4307)	(-1.2715)	(.19975)	(.013527)	(.015975)					
GARCH-M	-.58249	-.066277	.47863	.048491	.92992	.016260		-1387.6	.005171	9362.2
	(-1.2554)	(-1.3287)	(.19751)	(.013279)	(.015723)	(.71827)				
EGARCH	-.27233	-.063401	.092810	-.020762	.97308		.11739	-1389.9	.0043427	8929.0
	(-1.3722)	(-1.3122)	(.038302)	(.028120)	(.012336)		(.031273)			
DKRW1Y03										
GARCH	-.30041	-.066438	.47002	.045791	.93277			-1386.4	.0048462	9415.6
	(-1.5304)	(-1.3395)	9.19680)	(.013053)	(.015440)					
EGARCH	-.24526	-.075186	.085201	-.021741	.97535		.11262	-1387.8	.0050791	9422.6
	(-4.2294)	(-1.5518)	(.037010)	(.026878)	(.011953)		(.030073)			
DIDR6F03										
GARCH	.84433	.046827	190.5713	.19110	.71554			-2391.9	.0021983	672.2060
	(.52355)	(.81855)	(50.5650)	(.050859)	(.056812)					
GARCH-M	1.8289	.046812	192.2748	.19446	.71214	-7.576e-3		-2391.8	.0022303	658.5721
	(.57797)	(.81692)	(51.1548)	(.052620)	(.058110)	(-.36008)				
DIDR9F03										
GARCH	.84433	.046827	190.5713	.19110	.71554			-2391.9	.0021983	672.2060
	(.52355)	(.81855)	(50.5650)	(.050859)	(.056812)					
GARCH-M	1.8289	.046812	192.2749	.19446	.71214	-.7576e-		-2391.8	.0022303	658.5669
	(.57802)	(.81693)	(51.1557)	(.052620)	(.058111)	3				
EGARCH	.35148	.043970	1.2826	-.050657	.83119		.38227	-2395.1	0.0023029	
	(.20858)	(.75315)	(.31908)	(.042143)	(.043114)		(.070618)			656.9222
DTHB6F03										
GARCH	-.0070471	.057315	.0066037	.22618	.20309			400.1219	.0056663	18.8415
	(-1.5134)	(1.0242)	(.0019100)	(.063212)	(.17863)					
GARCH-M	-.015500	.050075	.0069280	.22604	.17276	.78876		400.2599	.0052277	16.3281

Notes: t-ratio and asymptotic s.e. are parentheses, Log-L denotes the maximized value of the log-likelihood function in each case. *χ2(2) and χ2(3), **MEU equals SQRT(2/3.14159)=.79788, stands for the expectation of the absolute value of the disturbance term. All test statistics are estimated converged after 25 to 67 iterations.

A2-II Figures:

Impulse response functions are shown for the Korean won and Thai baht for the spot and forwards rates. The two graphs present the impact of a generalized one standard deviation innovation in volatility of one exchange rate on itself and on the other volatilities in the system. Two standard error confidence bounds are presented around each impulse response function. The first column represents Korean won, the second column is Thai baht, and the third column shows Indonesian rupiah.



APPENDIX 3 Stochastic Dynamic Model for Phillips Curve Under Different Regimes[1]

To drives the dynamics of the equilibrium system (\`), a sets of assumptions are considered based on a stochastic dynamic Mundell-Fleming model

(4.a) $yd_t = d_t - \; r_t \; + \eta q_t$, and rewritten as $d_t \; = yg_t + d_{t-1} + d_t$.

(4.b) $m_t - p_t = y_t - \; i_t$. or $lnm_t d = ln(mt/pt) = \alpha + \alpha 1 lny_t - \alpha 2 i_t$, it may rewritten as $m_t = mgt + m_{t-1} + mt$.

(4.c) $p_t = \; p_t\text{\`} + (1- \;)E_{t-1}p_t\text{\`}$; under fixed regime, $p_t fx = \; p_t\text{\`} + (1+ \;)E_{t-1} (p_t\text{\`})^2$; under flexible regime, $ptfl = \; pt\text{\`}+(1- \;)E_{t+1}(pt\text{\`})$.

Where y^d represents aggregate demand in period t, specified as a function of an exogenous demand component, d_t. r_t refers the domestic real interest rate, and qt is real exchange rate. and η is positive elasticity. m_t is the money supply at time t, and *(>0)* denotes the elasticity of the demand for money. p_t is the domestic price level (equations 4.c). s_t is the spot exchange rate(equation 4.d). p_t\` is the market-clearing price and $E_{t-1}p_t$\` represents the price in the long term is set one period in advance to expectations of the future market clearing price. Where *(0, 1)* is the share of the market in domestic output as well as an element of price rigidity into the system i^* is the international interest rate, assumed to be constant over time for simplicity, $i_t = i^* + Et(s_{t+1} - s_t)$, the uncovered interest parity (*uip*). The expected rate of inflation then implies, $i_t\text{\`} = i^* + m^g - y^g$. Derivation of r_t for fixed regime is

(4.d^{fx}) $i_t - i_t\text{\`} = (i^* + E_t(s_{t+1} - s_t)) - i^* = 0$ and $r_t - r_t\text{\`} = i_t - i_t^*) - (\pi - \pi\text{\`}) = (1- \;/\eta)(\; y_t - \; d_t)$.

For float regime,

(4.dfl) $i_t - i_t\text{\`} = (i^* + E_t (s_{t+1} - s_t)) - (i^* + m^g - y^g)^3$ *and* $r_t - r_t\text{\`} = (i_t \; - E_t (p_{t+} - p_t)) \; (i_t^* - E_t (p_{t+1}\text{\`} - p_t t\text{\`}))^4$.

The equilibrium system consists of the four equations at each given time. The shock processes that drive the dynamics of the equilibrium system are

(4.e) $ys_t = y^g + ys_{t-1} + \; y_t$..

(4.ffx) $y_t - y_t\text{\`} = \eta(q_t \; - q_t\text{\`}) - \; (r_t - r_t\text{\`}) = -(1+ \;/\eta)(1- \;)(\; y_t - \; d_t)^{5.}$

(4.gfl) $y_t - y_t\text{\`} = (d_t - d_t\text{\`}) + \eta(q_t - q_t\text{\`}) - \; (r_t - r_t\text{\`}) = ((\; +\eta)(1+ \;)/ \; + \; +\eta)(1- \;)$

$(\; m_t - \; y_t)$ *and* $y_t = y_t\text{\`} + (( \; +\eta)(1+ \;)/ \; + \; +\eta)(1- \;)( \; m_t - \; y_t)$, *where* $y_t\text{\`} = y_t s$, where $q_t\text{\`} = 1/\eta(y_t s - d_t + \; i^*)$; $i_t = i_t{}^* = i\text{\`}$; $r_t = r_t\text{\`} + (1- \;/\eta)( \; y_t - \; d_t)$, $r_t\text{\`} = i^{*6}$; $s_t = s_t\text{\`} = s$; and $y_t = y_t\text{\`} - (1+ \;/\eta)(1- \;)(\; y_t - \; d_t)$; where $y_t\text{\`} = y_{ts}$.

---

[1] For the three selected countries, it is reasonable to assume that the loss function of central bank synchronously aims to a stability among inflation (π), economic growth(y), and exchange rate(e). That is $L = -1/2 \; E_t\Sigma_{s=1>\sigma} \; {}^s[\; {}_\pi(\pi_{t+s} - \pi^*)^2 + \; {}_y(y_{t+s} - y^*)^2 \; {}_e(e_{t+s} - e^*)^2]$, where $_{\pi,\; y,\; e}$ indicates the weight on the gap of inflation, growth, and exchange rate.

[2] [s + p*-1/η(y$_t$s-dt + i*)]+(1-){s+ p*-1/η[(y$_t$s- y$_t$)-(d$_t$- d$_t$)+ i*]}=p$_t$ \`+(1- /η)(y$_t$- d$_t$).

[3] et [(q t+1– q t) + (1-)(mt - yt) = Et (qt+1– qt) + (1-)(mt - yt)= (+η-1/ + +η)(1-)(mt - yt).

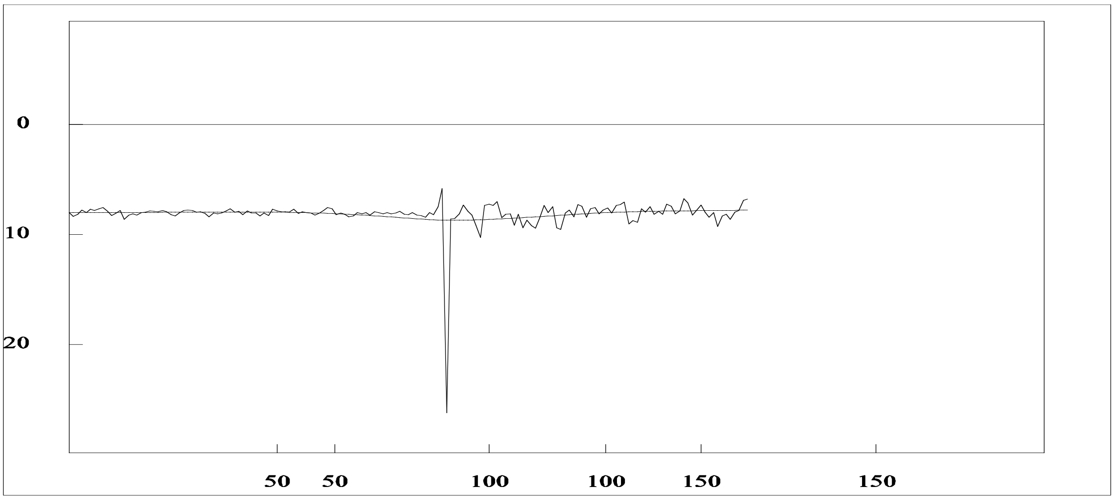
[4] =(it - it \`) – (1- )( mt - yt) =Et (qt+1 – qt) = rt \` + Et(qt+1– q)=rt\` - (1+ / + +η)(1-)(mt - yt)

[5] pt =p t\`-(1- /η)( yt- dt), where πt\`=0; qt = qt\`- (1- /η)( yt- dt)., pt = pt \` - (1-)(mt - yt), where pt \` = m t – yts + (i* + gm – gy); πt = π t \` + (1 mt - yt), where qi\` = 1/η(yts – d t + i*); i t = i t \` + (+η-1/ + +η)(1-)(mt - yt), where s t \` = mt + (1/η - 1)yts – (1/η)d t /η +)i* - p* + (mt - yt).

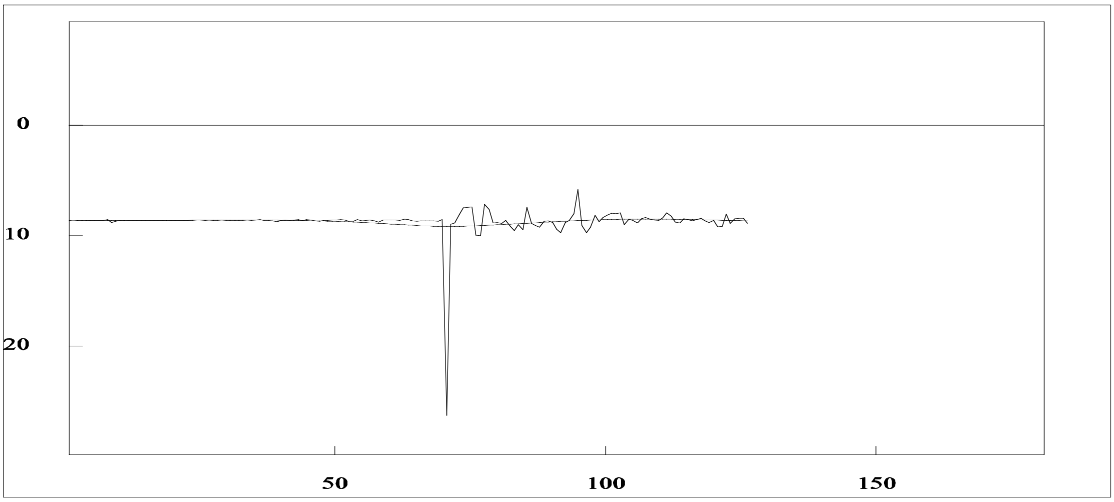
yg and mg are the deterministic growth rates of output and money. yε, dε, and mε are independently and identically distributed shocks of supply, demand, and money. Under flexible regime and exogenous supply of goods, output is supply determined, hence yt`=yts. Under fixed exchange rates, st = s =Et(st+1), so that it = i* by uip.

APPENDIX 4 Figure 3.2aThe monthly series of the H-P filter (=14400)of the first differences of the logarithmic Thai baht (i); Indonesian rupiah (ii); Korean won (iii) for the years 1999 to 2003.

(i) HPTW

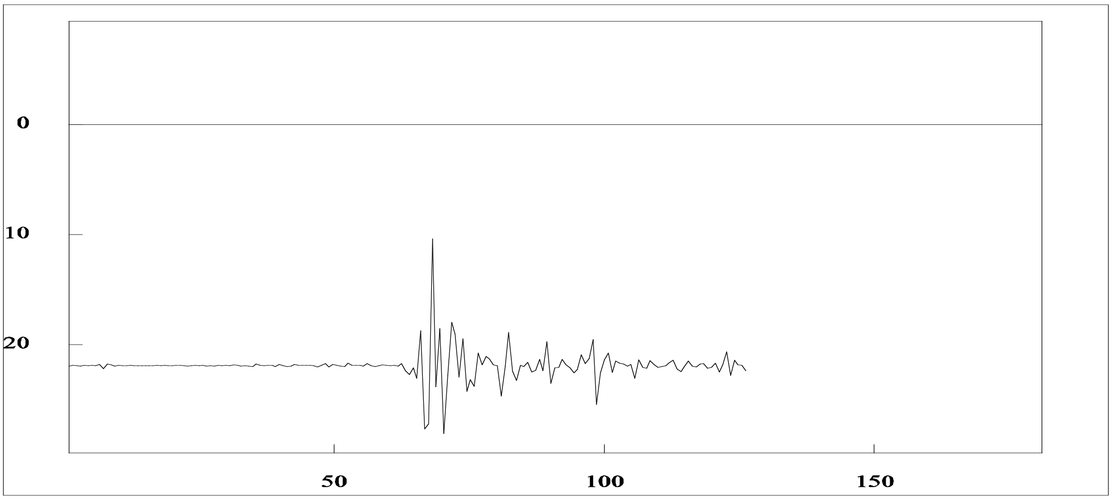


(ii) IWCHP



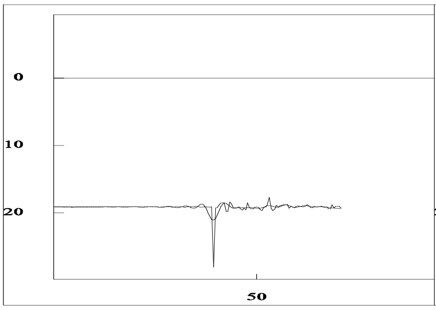
[6] Derivation of q t for fixed rate: qt -qt ` = (st + p*-pt)-(st `+p*-pt `) = -(1/θ/η)(εyt-εdt). For Flexible, qt is: η(qt – qt`) = (yt –yt`)+σ(rt -rt `), given yt =dt +ηqt -σrt =-(pt -pt`) + λ(it -it `)+σ(rt-rt `) using the money market equilibrium condition, = (λ+σ)E t [(q t+1– q t+1`) – (qt – qt`)] + (1+λ)(1-θ)(εmt - εyt), from Et(qt+1`) = 1/η(yst – d t + σi*) = q t `, the real exchange rate is from the assumption qt = qt` + κ(εmt - εyt) into the η(q t -q t `), solve for κ=(1+λ)(1-θ)/(λ+ση): qt = qt` + (1+λ/λ+σ+η)(1-θ)(εmt - εyt), so that Et (q t+1– q t) = - (1+λ/λ+σ+η)(1εmt - εyt), deviation of st: st – st` = (qt – qt`)+ (pt – pt`)-(p* - p*`)= - (σ+η-1/λ+σ+η)(1-θ)(εmt - εyt).

(iii) HPK

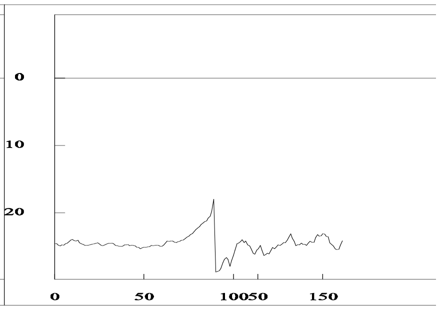


3.2b The monthly series of the H-P filter (λ=14400) of the logarithmic Thai baht (ii); Korean won (i); the twice difference of Thai baht (iv); Indonesian rupiah (v); Korean won (iii), and the first difference of the logarithmic of Korean won (vi), Indonesian rupiah (vii) and the top panel for Thai baht (viii): sample 1990M1-2004M12.

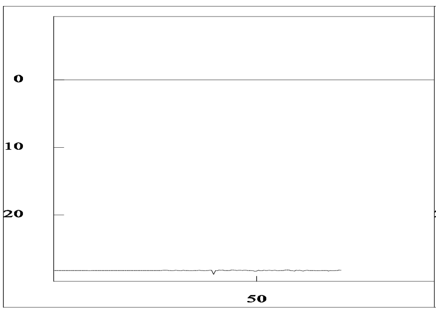
(i) HPKWCLOG



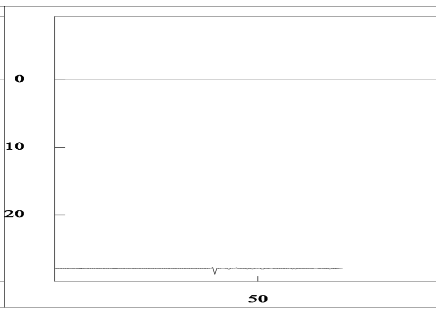
(ii) HPTWCLOG



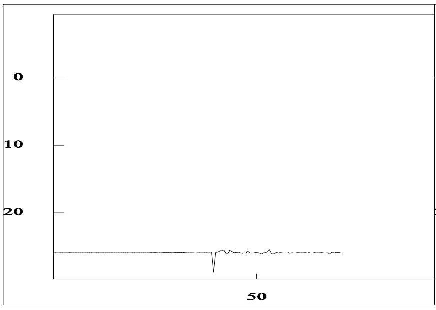
(iii) HPKWCFD2



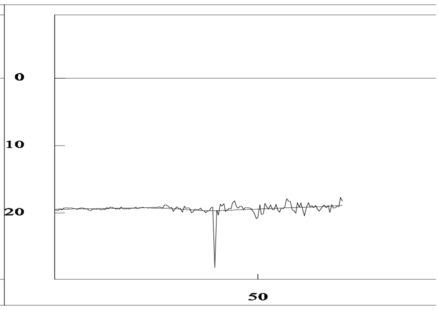
(iv) HPTWCFD2



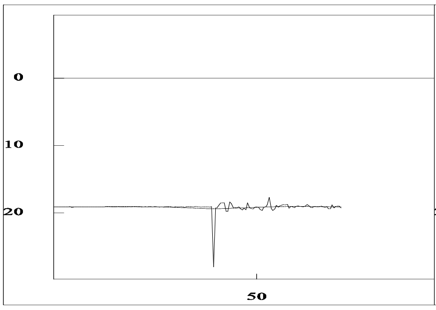
(v) HPIWCFD2



(vi) HPKWCFD



(vii) HPIWCFD



(viii) HPTWCFD

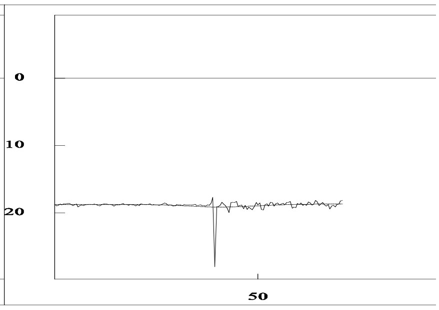
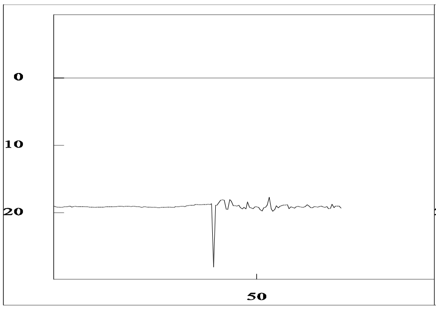
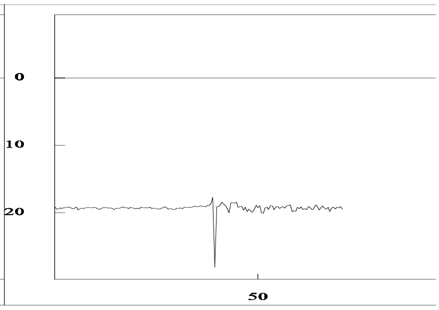


Figure 3.3 The polynomial order 12 of monthly series of the first differentiated Indonesian rupiah (i); Korean won (iii); Thailand (ii), and the logarithmic transferred polynomial order 12 monthly sequences of Korean won (iv), Indonesian rupiah (v), and Thai baht (vi) for the years: 1990-2004. (vii) presents the periodogram of the logarithmic won and (viii) the residual of periodogram obtained by fitting a linear trend through the logarithmic data of won. (ix) and (x) show the superimposed on the Butterworth filter, the gain of the 12^{th} order lowpass with cut-off frequency of wu=π/4 and wu=π/16, respectively.

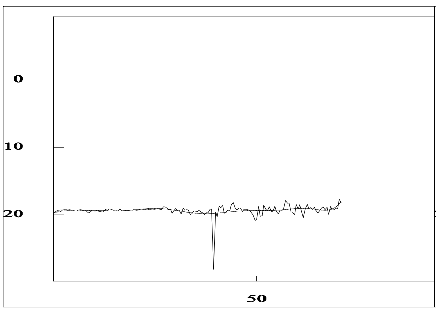
(i) POIWCFD



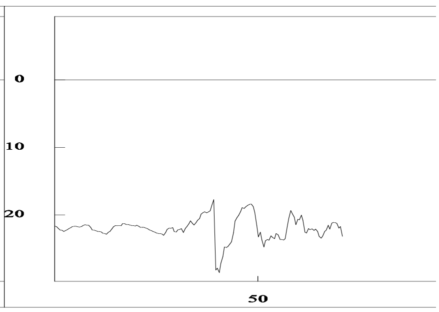
(ii) POTWCFD



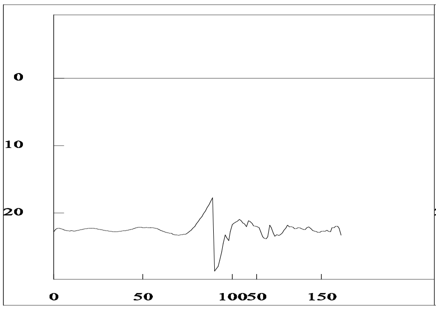
(iii) PKWCFD



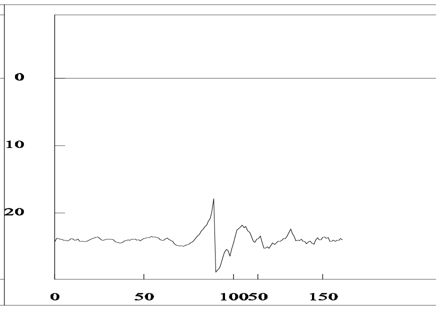
(iv) POKWCLO



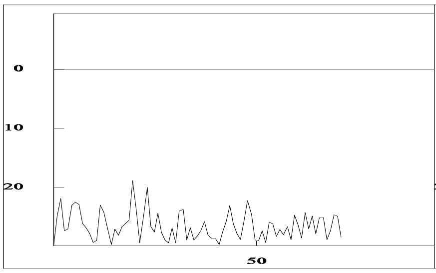
(v) POIWCLO



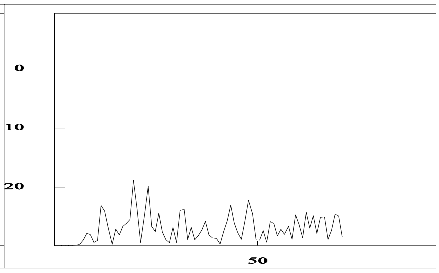
(vi) POTWCLO



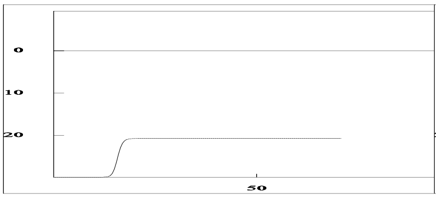
(vii) PGKWC



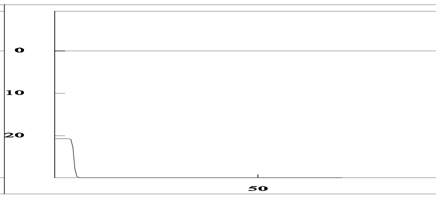
(viii) PGRKWC



(ix) F40



(x) FL12



Source: The monthly series of the ACA currencies are obtained from the Asian Development bank (ADB), all data are first or twice differenced or taking logarithmic transformation. The computer program is provided by Stephen Pollock and partially modified by the author for locating graphical production.

3.2 and 3.3d 3.3 in appendix 4).esents the H-P filter and the Butterworth filter.