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Market efficiency and cointegration of spot exchange rates during periods of economic turmoil: Another look at European and Asian currency crises

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Abstract

This study extends the study of foreign exchange market efficiency. It employs several verification testing procedures, rather than using only standard Johansen tests, to re-examine if cointegration among different spot exchange rates is actually present during the 1992–1993 European currency crisis and during the 1997–1998 Asian currency crisis. In contrast to the findings in prior studies, the test results collectively cast strong doubts on the presence of cointegration. Therefore, a cointegration test may not be an appropriate technique to detect and reveal market inefficiency if it in fact transpires during these two crises. Further, this study strongly corroborates empirical evidence that the reliance on Johansen tests can result in spurious findings of cointegration and thus incorrect inferences about efficiency.

JEL classification: C32; G14; F31

Keywords: Currency crisis; Spot exchange rates; Market efficiency

1. Introduction

Whether or not foreign exchange markets are efficient is of considerable interest to researchers and market participants. Among other econometric techniques, a cointegration analysis has been employed by several recent studies to examine foreign exchange market efficiency. The majority of prior empirical work (e.g., Coleman, 1990; Copeland, 1991; Lajaunie, McManis, & Naka,

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1996; Lajaunie & Naka, 1992; MacDonald & Taylor, 1989; Rapp & Sharma, 1999) has found that spot exchange rates for various major currencies generally are not cointegrated during the modern float. The absence of cointegration and thus a cointegrating vector and the error correction model (ECM) (e.g., Engle & Granger, 1987) implies that the current value of one currency cannot be predicted by past values of other currencies. This unpredictability is commonly interpreted by prior studies as evidence of weak-form efficiency in foreign exchange markets.

Interpretational ambiguity mainly arises when cointegration among spot exchange rates is detected. Baillie and Bollerslev (1989) interpret the predictability implied by cointegration as a violation of weak-form efficiency or as indirect evidence of a time-varying risk premium. Crowder (1994) finds that a cointegrating vector which is stationary or $I(0)$ by definition does not appear to be a proxy for a time-varying risk premium. This is because the forward premium used to represent the risk premium has a different time series property in that it appears non-stationary or $I(1)$. Baillie and Bollerslev (1994) contend that the forward premium is not a pure $I(1)$ process but fractionally integrated and mean reverting with finite cumulative impulse response weights. On the other hand, Wu and Chen (1998) employ a more powerful unit root test, find that the forward premium is in fact stationary and conclude that foreign exchange markets are efficient even though the presence (or lack thereof) of cointegration is not examined.

Jeon and Lee (2002) find that the G-7 countries' exchange rates are cointegrated during the period between the Plaza Agreement in 1985 and the Louvre Accord in 1987. They conclude that market inefficiency transpires during this period of international policy cooperation to stabilize exchange rates. Further, Haug, Mackinnon, and Michelis (2000) and Rangvid and Sorensen (2002) detect cointegrating relations among exchange rates of the European Union (EU) countries over extended time periods prior to the inception of the European Monetary Union (EMU) in 1999. They however interpret this result as an indication of stability and credibility of the EU exchange rate policy coordination through the Exchange Rate Mechanism (ERM) rather than as evidence of market inefficiency. The Maastricht Treaty (1992) which requires convergence of key economic variables, including exchange rates, among EU nations prior to becoming EMU members can further explain such finding.

In relation to voluminous studies using long spans of data, a few studies have performed cointegration tests of spot exchange rates during periods of economic turmoil. Aroskar, Sarkar, and Swanson (2004) find a cointegrating relation among daily spot exchange rates of EU currencies during the European currency crisis of 1992 and 1993. They suggest that weak-form inefficiency exists during the crisis partially because the ECM provides better predictive power for some included currencies than does the random walk model. Further, Aroskar and Swanson (2002) and Jeon and Seo (2003) evidence a cointegrating relation among daily spot exchange rates of Asian currencies during the Asian currency crisis of 1997 and 1998. These two studies conclude that weak-form inefficiency occurs during the crisis as well.

Similar to the analyses using long time spans, whether or not cointegration really indicates market inefficiency during these two crises is open to debate. On the one hand, cointegration and its embedded predictability can emerge and entail arbitrage opportunities if foreign exchange markets are truly inefficient during turbulent times. On the other hand, Dwyer and Wallace (1992), Baffes (1994), Engel (1996), Masih and Masih (2001) and Ferre and Hall (2002) indicate that cointegration does not necessarily imply market inefficiency, implying further that cointegration tests may not be appropriate tests of market efficiency for any period. This is because whether or not the predictability derived from cointegration can truly lead to arbitrage opportunities and/or the ability of market participants to earn risk-adjusted excess returns has not been confirmed nor verified. Further, Lence and Falk (2005) indicate that cointegration test results have no implications about market efficiency without additional restrictions on the economy or economies.

The upshot is that if cointegration among spot exchange rates during crises (such as the European and Asian crises) is in fact absent in the first place, the possible connection and relevancy of cointegration to inefficiency are mitigated. These include, for instance, whether or not the resultant ECM would provide better predictive ability in relation to the random walk model, whether or not the cointegrating vector could be a proxy for a risk premium, and whether or not cointegration would result in arbitrage opportunities and thus truly imply weak-form inefficiency. Conceptually, in this case, the exchange rates of currencies affected by the crises may simply exhibit considerable volatilities without being cointegrated with one another. It is also possible that the EU exchange rates are cointegrated during the non-crisis period due to the ERM and Maastricht Treaty (1992), but are not cointegrated during the crisis period due to the abandonment of relevant currencies from such exchange rate mandates.

Thus, the question is whether or not evidence of cointegration during these two crises has been identified correctly. Hakkio and Rush (1991) indicate that cointegration is a long-run property, and hence extensive time spans, rather than high data frequency used in a short-term crisis period analysis, should be employed to appropriately detect the cointegration evidence. Further, previous studies of spot exchange rates during currency crises (i.e., Aroskar & Swanson, 2002; Aroskar et al., 2004; Jeon & Seo, 2003) base their cointegration findings mainly on the Johansen cointegration methodology (e.g., Johansen, 1988). Sephton and Larson (1991) and Crowder (1996), however, indicate that the statistical power of cointegration tests can be highly suspect. Particularly, the Johansen methodology can result in indeterminate and/or incorrect inferences regarding the presence of cointegration and the number of cointegrating vectors in the system (e.g., Gonzalo & Lee, 1998; Hall, 1991). Given this possibility, the reliance on conventional Johansen tests may have caused spurious findings of cointegration and thus incorrect inferences of inefficiency during the two crises. These misleading findings and inferences can be harmful. For instance, investors in foreign exchange markets might take unnecessarily risky positions in affected currencies and hope to exploit arbitrage opportunities based on the estimate of a cointegrating relation which in fact does not exist. Thus, expanded methodologies are needed to verify the results from any one specific cointegration methodology typically relied upon by past researchers.

Given the importance of accuracy in measuring foreign exchange market efficiency, this study re-examines whether or not a cointegrating relation is really present among spot exchange rates of the affected currencies during the 1992–1993 European currency crisis and during the 1997–1998 Asian currency crisis. Several verification testing procedures which have not been considered in conjunction with one another in prior studies are included. First, the conventional Johansen cointegration test (e.g., Johansen, 1988) is performed. Then, its variant which is based on the partial VAR system (e.g., Pesaran, Shin, & Smith, 2000) and provides some econometric advantages over the conventional procedure is implemented. Further, the recursive test of cointegration parameter stability (Hansen & Johansen, 1999); the unit root test of cointegrating vector and common trend estimates based on the Gonzalo–Granger decomposition (Gonzalo & Granger, 1995); the Harris–Inder (HI) cointegration test (Harris & Inder, 1994) under the reversed null hypothesis of cointegration; and the Gregory–Hansen (GH) cointegration test (Gregory & Hansen, 1996) which considers the possibility of an endogenous structural shift in a cointegrating relation are conducted. This battery of tests can potentially add clarity and a new perspective to researchers and foreign exchange market participants concerning evidence of cointegration during the two currency crises which might have been detected spuriously in prior studies based on standard Johansen procedures.

This study finds that European exchange rates and Asian exchange rates are cointegrated during the European currency crisis and during the Asian currency crisis, respectively, based on the Johansen tests for the full VAR system and/or for the partial VAR system. However, the cointegration parameters obtained show evidence of instability, especially in the group of Asian exchange rates. The unit root tests also reveal that the estimates of cointegrating vectors and common trends, which conceptually should be stationary and non-stationary, respectively, appear to be identically non-stationary. Further, the null hypothesis of cointegration in the HI tests can be clearly rejected in both European and Asian exchange rate groups. Finally, the null hypothesis of no cointegration in the GH tests cannot be rejected in any exchange rate group. This is true irrespective of the specification for an endogenous structural shift in a cointegrating relation under the alternative hypothesis. These findings collectively cast strong doubts on the validity of cointegration findings detected through Johansen tests and consequently on the usefulness of cointegration tests to reveal market inefficiency even if foreign exchange rate markets are truly inefficient during these two crises.

The remainder of this study is organized as follows. Section 2 describes the data and explains the econometric methodology used, with estimation results and related findings set forth in Section 3. Section 4 provides conclusions.

2. Data and methodology

Daily exchange rates for the British pound, French franc, German mark and Italian lira are obtained from the Federal Reserve Bank of St. Louis. Consistent with those in Aroskar et al. (2004), the data cover the European currency crisis period from September 16, 1992 to March 31, 1993. Further, daily exchange rates for the Indonesian rupiah, Korean won, Malaysian ringgit and Thai baht are acquired from the Datastream International Databank. Also consistent with those in Aroskar and Swanson (2002), the data cover the Asian currency crisis period from September 1, 1997 to August 31, 1998. Each exchange rate is expressed as the US dollar price of one unit of a respective currency. All exchange rates are transformed into natural logarithms for further analyses.[1]

Cointegration presupposes that variables in the system are non-stationary and integrated of the same order. Therefore, the Augmented Dickey–Fuller (ADF) unit root test (Dickey & Fuller, 1979, 1981) is employed to examine the univariate property of each exchange rate series under the null hypothesis that the series is non-stationary and integrated of order one or $I(1)$. Further, the KPSS stationarity test (Kwiatkowski, Phillips, Schmidt, & Shin, 1992) based on the reversed null hypothesis of stationarity or $I(0)$ is performed to ensure the consistency of ADF unit root test results.[2] Finally, the Zivot–Andrew (ZA) unit root test (Zivot & Andrews, 1992) which allows for an endogenous structural break in a series is conducted to mitigate the bias towards non-rejection of the unit root null hypothesis if the series is in fact stationary but subjected to a structural break. This unit root test is advisable, in contrast to other similar tests (e.g., Perron, 1989), because the

[1] Selection of the crisis periods and included currencies is based on past studies (Aroskar & Swanson, 2002; Aroskar et al., 2004) that allow direct comparisons of cointegration test results. Based on the two studies, these currencies are chosen because, unlike the less obvious others, they showed significant losses or considerable depreciation in values during the crises. Further, the beginning of samples closely match the date on which Thailand let its baht float and the date on which the UK withdrew its pound from the EMS. Readers are referred to these two studies for detailed explanations on the selection process.

[2] The ADF and KPSS test equations are well known and thus are omitted to conserve space.

break date is not assumed a priori but rather is endogenously determined by the test equation. For each possible break date, T_{PB}, which runs from the observation $0.15T$ to the observation $0.85T$ where T is the sample size, the unit root test equation is estimated for each exchange rate series X_i as in (1):

$$\Delta X_{\mathrm{it}} = \mu_i + \theta \mathrm{DU}_{\mathrm{it}} + \alpha X_{\mathrm{it}-1} + \sum_{j=1}^{k} \phi_i \Delta X_{\mathrm{it}-j} + \varepsilon_{\mathrm{it}} \tag{1}$$

where $\mathrm{DU}_t = 1$ for $t > T_{\mathrm{PB}}$, and zero otherwise, and k is the number of augmented lags which is determined by the general-to-simple procedure.[3] The t-statistic for testing $\alpha = 0$ or t_α is computed for each T_{PB} iteration. Finally, the smallest value of t_α's calculated for all T_{PB} iterations becomes the ZA test statistic under the null hypothesis that the X_{it} is $I(1)$ against the alternative hypothesis that it is $I(0)$ with a structural break. If the null hypothesis is rejected, the T_{PB} associated with the ZA statistic becomes T_{B} or the date on which a structural break in a series transpires.

For the European or Asian group of p non-stationary exchange rates, the Johansen cointegration procedure (e.g., Johansen, 1988) is implemented by constructing a VAR(k) process as in (2):

$$X_t = \Phi_1 X_{t-1} + \cdots + \Phi_k X_{t-k} + \mu + \varepsilon_t \tag{2}$$

where X_t is a p-dimensional vector of exchange rates, Φ_j the coefficient matrices, μ the vector of constants, ε_t the white noise error vector with non-diagonal covariance matrix Ω, and k is the minimum lag length that reduces serial correlation in residuals in each equation in the VAR to zero statistically based on the Ljung–Box (L–B) Q-statistics.[4]

The VAR in Eq. (2) can be transformed into its error correction model representation as in (3):

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-1} + \mu + \varepsilon_t \tag{3}$$

where the matrix of interest is the long-run multiplier matrix $\Pi = \Phi(1) - I$. The Π matrix can be decomposed into two $(p \times r)$ matrices such that $\alpha\beta' = \Pi$. The β matrix contains parameters for r cointegrating vectors (CIVs) or long-run stationary equilibria which imply the presence of $(p–r)$ non-stationary common trends (CTs) driving the system of exchange rates, while the α matrix contains error correction coefficients which measure the extent to which each exchange rate responds to deviations from the long-run equilibria. The test for cointegration is the rank test for r non-zero eigenvalues (λ_i). The test statistic for the null hypothesis of at most r CIVs against the alternative of p CIVs is the λ_{trace} statistic given in (4):

$$\lambda_{\mathrm{trace}} = -T \sum_{i=r+1}^{p} \ln(1 - \lambda_i) \tag{4}$$

[3] The test equation with $k_{\max}$ lags is estimated and $k = k_{\max}$ is chosen for that T_{PB} iteration if the coefficient on the last augmented lag is statistically significant at the 10% level. Otherwise, $k_{\max}$ is reduced to k which marginally enables such coefficient to be statistically significant at the 10% level. In this study, $k_{\max}$ is set equal to 12.

[4] According to Crowder and Wohar (1998), a crucial criterion to choose k in the VAR is that serial correlation in residuals is eliminated. Some studies (e.g., Kasa, 1992) suggest the selection of k based on normality of residuals. However, this selection criterion may not be appropriate because the Johansen estimation procedure is invariant to non-normal errors (Gonzalo & Granger, 1994).

The test statistic for the null hypothesis of r against the alternative of $r+1$ CIVs is the $\lambda_{\max}$ statistic given in (5):

$$\lambda_{\max} = -T \ln(1 - \lambda_{r+1}) \tag{5}$$

The critical values for λ_{trace} and $\lambda_{\max}$ statistics are obtained from MacKinnon, Haug, and Michelis (1999).[5]

Further, the partial or conditional VAR system (e.g., Pesaran et al., 2000) is constructed based on the assumption that some variables in the full VAR system are weakly exogenous with respect to α and β. These variables are thus the sources of common trends and do not respond to deviations from long-run equilibria. Hansen and Juselius (1995) suggest that conditioning on weakly exogenous variables can improve stochastic properties of the model. Specifically, Harbo, Johansen, Nielsen, and Rahbek (1999) indicate that the partial VAR enables a reduction in the system dimension (which in turn should increase the test power) and the weak exogeneity assumption is essential in achieving the distributions of test statistics free of nuisance parameters. MacKinnon et al. (1999) further suggest that the partial VAR is beneficial because, among others, it explicitly allows for the effects of weakly exogenous variables on the distributions of λ_{trace} and $\lambda_{\max}$ statistics.

Hence, weakly exogenous exchange rates must be identified through the weak exogeneity test (e.g., Johansen, 1992). Conditional on the presence of r CIVs in the full VAR, testing the null hypothesis that one exchange rate is weakly exogenous can be done by restricting the relevant row in the α matrix (i.e., the row associated with the exchange rate being tested for weak exogeneity) to zero. The resultant likelihood ratio test statistic is distributed as χ^2 with r degrees of freedom. Due to the possibility that the presence of cointegration and the number of CIVs may not be determined correctly by Johansen tests, the weakly exogenous exchange rates used in the partial VAR construction are ones which consistently allow non-rejection of the weak exogeneity null hypothesis across all potential r's in the full VAR system.

Then, let X_t in Eq. (2) be partitioned into an n-vector Y_t and an m-vector Z_t where $n = p - m$; and Y_t and Z_t are endogenous and weakly exogenous exchange rates, respectively. Thus, $X_t = (Y_t', \; Z_t')'$ and $\Pi = (\Pi_{Y'}, \; P_{Z'})'$ so that the conditional ECM for Y_t can be expressed as in (6):

$$\Delta Y_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi_Y X_{t-1} + \Lambda \Delta Z_t + \mu_Y + \varepsilon_{Y_t} \tag{6}$$

Similar to the full system, the test for cointegration in the partial system can be achieved by decomposing Π_Y into α and β matrices and calculating the λ_{trace} and $\lambda_{\max}$ statistics. However, these two statistics have different distributions from those in the full VAR and the critical values are thus tabulated separately in MacKinnon et al. (1999).

While the λ_{trace} and $\lambda_{\max}$ statistics generally have reasonable size and power properties (e.g., Haug, 1996), empirical evidence suggests that total reliance on the Johansen test may not result in correct inferences concerning the presence of cointegration and the number of CIVs, especially in large systems. Hall (1991) and Ahlgren and Antell (2002) show that Johansen test results are quite sensitive to the lag length included in the VAR. Gregory (1994), Richards (1996) and Gonzalo and Lee (1998, 2000) further demonstrate that the Johansen test tends to spuriously reject the null

[5] MacKinnon et al. (1999) compute numerical distribution functions of these two statistics. They find that the resultant critical values are more accurate than those in Osterwald-Lenum (1992) which generally have been used by prior empirical work on cointegration.

hypothesis of no cointegration under certain conditions.[6] Hence, four additional tests to validate the Johansen test results are conducted.

First, the recursive stability test of the cointegration space or cointegration parameters (Hansen & Johansen, 1999) is performed. Based on the notion that the parameters should be stable if the model is to be valid and useful, the test is implemented by: (1) holding short-run dynamics constant at their full-sample estimates but allowing long-run relations to change over time; (2) estimating the ECM over the base period; and (3) keeping initial observations in the base period fixed and increasing one additional observation at each iteration to re-estimate the ECM such that the last estimation window is equal to the full sample. Conditional on the r CIVs over the entire sample, the relevant null hypothesis is that the β estimate at one recursive iteration does not differ statistically from the full-sample β estimate which has the lowest sample variance. The test statistic is distributed as χ^2 with $(p-r)r$ degrees of freedom.

Second, the ADF unit root tests are performed on the CIV and CT estimates from the Johansen test.[7] Gonzalo and Granger (1995) demonstrate that in the cointegrated system of p variables with r CIVs and thus $(p-r)$ CTs, the vector of variables X_t can be decomposed into stationary and permanent or non-stationary components as in (7):

$$\begin{aligned} X_t &= \text{stationary component} + \text{permanent or non-stationary component} \\ &= \alpha(\beta'\alpha)^{-1}\beta' X_t + \beta_\perp(\alpha'_\perp\beta_\perp)^{-1}\alpha'_\perp X_t \end{aligned} \tag{7}$$

where $\alpha_\perp$ and $\beta_\perp$ are the $(p\times(p-r))$ matrices which satisfy $\alpha'\alpha_\perp = 0$ and $\beta'\beta_\perp = 0$; and $\beta' X_t$ and $\alpha'_\perp X_t$ are CIVs and CTs, respectively. Using this decomposition, the estimates of $\beta' X_t$ and $\alpha'_\perp X_t$ must exhibit sharply contrasting properties in that the first are stationary while the latter are non-stationary. Because the coefficients for CIVs and CTs must be estimated, the critical values for ADF unit root tests of these estimates are different from those for ADF unit root tests of series whose values are known a priori. The requisite critical values are obtained from MacKinnon (1996).

Third, the Harris–Inder (HI) cointegration methodology (Harris & Inder, 1994) is implemented.[8] This methodology combines the Engel–Granger single-equation procedure (Engle & Granger, 1987) with the KPPS stationarity test of residuals (Kwiatkowski et al., 1992). Thus, unlike Johansen and standard Engel–Granger tests with the null hypothesis of no cointegration, the HI approach is based on the reversed null hypothesis of cointegration. Harris and Inder (1994) recommend testing both null hypotheses unless there is a priori belief that cointegration is present or absent. This suggestion is especially of relevance. Whether or not various spot exchange rates are truly cointegrated is not known and there is no underlying economics or finance theory suggesting so. A brief review of the HI procedure follows. The Engel–Granger test equation is estimated

[6] These include, for example, when variables are $I(d, d>0.5)$ or have non-stationary long memory, but they are difficult to be distinguished from the pure $I(1)$ processes using standard unit root tests; and when the VAR representation of variables has a singular or near-singular error covariance matrix.

[7] The unit root test of the CIV estimates is also performed in Crowder and Wohar (1998) and Phengpis and Apilado (2004) for the system of international stock market price indices. It is not found that prior studies have conducted unit root tests of the CIV and CT estimates in the system of spot exchange rates.

[8] The HI procedure is also used, for example, in Dutt (1998) and Choudhry (1999a) for testing the purchasing power parity (PPP), in Choudhry (1999b) and Dutt and Ghosh (1999a,b) for testing the forward market unbiasedness hypothesis in foreign exchange markets, and in Chang (2001) for testing cointegration between stock market price indices of Taiwan and its major trading partners. It is not found that previous studies have employed the HI test in investigating cointegration among spot exchange rates.

via the OLS for each exchange rate X_{it} as in (8):

$$X_{\mathrm{it}} = \mu_i + \beta X_t + \varepsilon_{\mathrm{it}} \tag{8}$$

where μ_i is a constant, β the vector of (slope) coefficients, X_t the vector of other exchange rates in the system as explanatory variables, and $\varepsilon_{\mathrm{it}}$ is a stationary or equilibrium error if cointegration exists. Then, the KPSS stationarity test is conducted on the residual $\hat{\varepsilon}_{\mathrm{it}}$ from Eq. (8) and the KPSS test statistic is compared against the critical values in Sephton (1996) under the null hypothesis of cointegration.

Finally, the Gregory–Hansen cointegration methodology (Gregory & Hansen, 1996) which considers the possibility of an endogenous structural shift in a cointegrating relation is conducted. The GH procedure is particularly useful for two reasons. First, it mitigates the bias towards finding no evidence of cointegration because standard cointegration tests may spuriously fail to reject the null hypothesis of no cointegration even though a cointegrating relation is actually present but subjected to a structural shift. Further, it alleviates the "data mining" problem by allowing the date for a structural shift to be determined endogenously by the test statistic from model estimation. Given these benefits, two models of endogenous structural shifts are examined. For each possible date on which a structural shift can occur, T_{PS}, the "level shift" model which is a variant of Eq. (8) and enables a shift in the intercept of a cointegrating relation is estimated as in (9):

$$X_{\mathrm{it}} = \mu_{i1} + \mu_{i2}\mathrm{DU}_t + \beta X_t + \varepsilon_{\mathrm{it}} \tag{9}$$

where μ_{i1} represents the intercept before the shift; μ_{i2} represents a change in the intercept at the time of the shift; and $\mathrm{DU}_t = 1$ for $t > T_{\mathrm{PS}}$, and zero otherwise. Further, for each T_{PS}, the "regime shift" model which permits a shift in both intercept and slope coefficients of a cointegrating relation is estimated as in (10):

$$X_{\mathrm{it}} = \mu_{i1} + \mu_{i2}\mathrm{DU}_t + \beta_1 X_t + \beta_2 X_t \mathrm{DT}_t + \varepsilon_{\mathrm{it}} \tag{10}$$

where β_1 represents the cointegration slope coefficients before the shift; μ_{i2} represents a change in the slope coefficients at the time of the shift; and $\mathrm{DT}_t = 1$ for $t > T_{\mathrm{PS}}$, and zero otherwise. T_{PS} is allowed to vary from the observation $0.15T$ to the observation $0.85T$, where T is the sample size. The residual $\hat{\varepsilon}_{\mathrm{it}}$ for each T_{PS} iteration is then subjected to the ADF unit root test where the number of augmented lags, k, in the test equation is determined by the general-to-simple procedure.[9] Ultimately, the smallest value of the ADF statistics calculated for all T_{PB} iterations becomes the relevant test statistic, ADF*, under the null hypothesis of no cointegration against the alternative hypothesis of cointegration with a structural shift. Should the null hypothesis be rejected, the T_{PS} associated with the ADF* becomes T_{S} or the date on which a structural shift in a cointegrating relation occurs.

3. Empirical results

Table 1 shows the results from the ADF and ZA unit root tests and KPSS stationary tests of the exchange rate series. The null hypothesis of a unit root or non-stationarity in the ADF and ZA tests cannot be rejected for any exchange rate, except for the Italian lira (IL) where the null

[9] This procedure is identical to the one for the ZA test and $k_{\max}$ is set equal to 12.

Table 1
Unit root tests of exchange rates

Exchange rate	ADF test statistic[a]	KPSS test statistic[b]	ZA test[c]	
			ZA test statistic	T_{PB}
European				
British pound (BP)	−2.05	0.39**	−4.00	10/16/1992
French franc (FF)	−1.77	0.40**	−4.27	10/16/1992
German mark (GM)	−1.88	0.34**	−4.48	10/16/1992
Italian lira (IL)	−3.50**	0.13*	−4.87**	12/28/1992
Asian				
Indonesian rupiah (IR)	−1.65	0.65**	−4.65	01/01/1998
Korean won (KW)	−1.04	1.43**	−4.31	11/20/1997
Malaysian ringit (MR)	−1.98	1.30**	−3.42	01/30/1998
Thai baht (TB)	−2.93	1.41**	−3.26	02/02/1998

Note: The exchange rate is measured as the US dollar price of one unit of foreign currency; * and ** indicate rejection of the null hypothesis at the 10% and 5% levels, respectively.

[a] Critical values from MacKinnon (1996) are −3.1266 and −3.4098 at the 10% and 5% levels, respectively.

[b] Critical values from Kwiatkowski et al. (1992) are 0.119 and 0.146 at the 10% and 5% levels, respectively.

[c] The ZA test statistic is the smallest value of the t_α statistics calculated for all possible dates for a structural break. T_{PB} is the possible structural shift date associated with the ZA statistic. The critical value from Zivot and Andrews (1992) is −4.80 at the 5% level.

Table 2
Johansen cointegration tests in the full VAR system

European crisis			Asian crisis		
Equation	$k=1$	$k=2$	Equation	$k=8$	$k=9$
Panel A: lag length selection[a]					
BP	17.56	9.31	IR	9.62	5.60
FF	6.50	6.92	KW	49.01**	11.36
GM	5.76	5.27	MR	6.55	4.28
IL	25.47**	14.75	TB	5.17	3.33

European crisis				Asian crisis			
H_0	λ_{trace}	H_0	λ_{max}	H_0	λ_{trace}	H_0	λ_{max}
Panel B: cointegration rank[b]							
$r=0$	58.01**	$r=0$	32.86**	$r=0$	48.42	$r=0$	21.80
$r<1$	25.15	$r=1$	14.87	$r\leq 1$	26.62	$r=1$	12.42
$r\leq 2$	10.28	$r=2$	6.50	$r\leq 2$	14.20	$r=2$	10.32
$r\leq 3$	3.78	$r=3$	3.78	$r\leq 3$	3.88	$r=3$	3.88

Note: **indicates rejection of the null hypothesis at the 5% level.

[a] The full VAR is estimated based on the deterministic specification that constants are present in the cointegration space. k is the number of lags included. The number shown is the L–B $Q(12)$ calculated from each equation for exchange rates under the null hypothesis of no serial correlation in residuals up to 12 lags.

[b] Based on the VAR with $k=2$ for the European group or on the VAR with $k=9$ for the Asian group, the λ_{trace} and λ_{max} statistics are compared against critical values tabulated in MacKinnon et al. (1999).

Table 3
Weak exogeneity tests

European crisis					Asian crisis				
r	BP	FF	GM	IL	r	IR	KW	MR	TB
1	0.14	11.14**	7.91**	7.39**	1	5.51**	1.31	3.36*	6.93**
2	1.56	13.71**	12.12**	10.90**	2	7.60**	1.70	3.41	8.67**
3	3.80	14.67**	12.80**	12.75**	3	13.93**	5.80	8.26**	14.30**

Note: r denotes the number of potential cointegrating vectors (CIVs) in the full VAR system. Conditional on the presence of r CIVs, the number shown is the likelihood ratio test statistic under the null hypothesis that the exchange rate i is weakly exogenous. This test statistic is distributed as χ^2 with r degrees of freedom; * and ** indicate rejection of the null hypothesis at the 10% and 5% levels, respectively.

hypothesis can be rejected at the 5% level. Further, the null hypothesis of stationarity in the KPSS test can be rejected at the 5% level for all exchange rates, except for IL where the null hypothesis can be rejected at only the 10% level. Because none of the exchange rate series can be clearly and consistently considered stationary, all of them are subjected to cointegration tests.

Table 2 presents the results from Johansen cointegration tests for the full VAR systems. In the European group, the lag length k in the VAR is set equal to 2 because it is the minimum sufficient to eliminate serial correlation in residuals based on the L–B Q-statistics. If k were reduced to 1, the VAR equation for IL would have serial correlation in residuals which is statistically significant at the 5% level (see the first three columns in Panel A). With $k=2$, the λ_{trace} and $\lambda_{\max}$ statistics consistently indicate one CIV among European exchange rates during the European currency

Table 4
Johansen cointegration tests in the partial VAR system

European crisis			Asian crisis		
Equation	$k=1$	$k=2$	Equation	$k=4$	$k=5$
Panel A: lag length selection[a]					
FF	10.88	10.34	IR	31.53**	12.13
GM	13.91	10.39	MR	10.67	9.91
IL	26.50**	10.08	TB	3.60	3.02

European crisis				Asian crisis			
H_0	λ_{trace}	H_0	$\lambda_{\max}$	H_0	λ_{trace}	H_0	$\lambda_{\max}$
Panel B: cointegration rank[b]							
$r=0$	50.44**	$r=0$	32.72**	$r=0$	41.05*	$r=0$	24.99*
$r\leq 1$	17.72	$r=1$	13.45	$r\leq 1$	11.79	$r=1$	16.05
$r\leq 2$	4.27	$r=2$	4.27	$r\leq 2$	4.26	$r=2$	4.26

Note: * and ** indicate rejection of the null hypothesis at the 10% and 5% levels, respectively.

[a] The partial VAR is estimated based on the deterministic specification that constants are present in the cointegration space. The partial VAR for the European group has BP as a weakly exogenous exchange rate, while the partial VAR for the Asian group has KW as a weakly exogenous exchange rate. k is the number of lags included. The number shown is the L–B $Q(12)$ calculated from each equation in the partial VAR for endogenous exchange rates under the null hypothesis of no serial correlation in residuals up to 12 lags.

[b] Based on the partial VAR with $k=2$ for the European group or on the partial VAR with $k=5$ for the Asian group, the λ_{trace} and $\lambda_{\max}$ statistics which have different distributions from those from the full VAR are compared against critical values tabulated separately in MacKinnon et al. (1999).

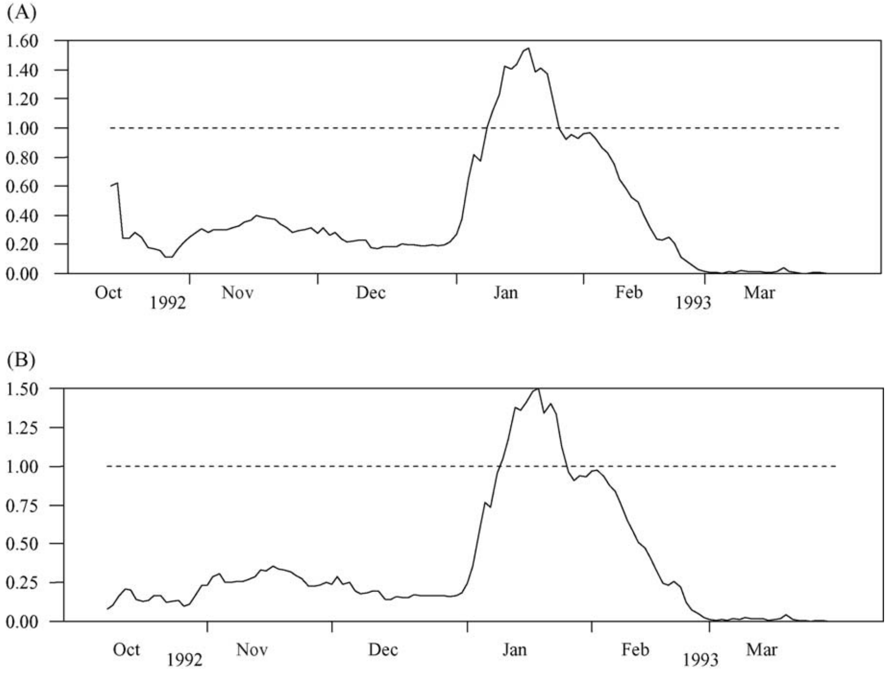

Fig. 1. Recursive tests of CIV stability in the European group. Panel A: full VAR system; Panel B: partial VAR system. *Note*: Conditional on the presence of one CIV in the full VAR or in the partial VAR, the recursive likelihood ratio test statistics (scaled by the 5% critical value) are plotted against the *end* of each estimation window. The first estimation window is recursively increased until the last estimation window covers the full sample. The plot above the critical value line of one indicates rejection of the null hypothesis that the β estimate from the respective window is not statistically different from the β estimate derived from the full sample at the 5% level.

crisis at the 5% level (see the first four columns in Panel B). In the Asian group, k in the VAR is set equal to 9. If k were reduced to 8, the VAR equation for the Korean won (KW) would have serial correlation in residuals which is statistically significant at the 5% level (see the last three columns in Panel A). With $k=9$, both λ_{trace} and $\lambda_{\max}$ statistics indicate no CIV among Asian currencies during the Asian crisis (see the last four columns in Panel B). Hence, with the appropriate selection criterion for k, the conventional Johansen test results for the group of Asian exchange rates already appear contradictory to those in previous studies.[10]

Prior to the partial VAR analysis, the weak exogeneity tests are performed. The results are shown in Table 3. Given the possibility that Johansen tests may not correctly determine the number of CIVs (r's), the exchange rates which allow the non-rejection of the weak exogeneity null hypothesis across all potential r's ($r=1$, 2 and 3) are chosen to represent weakly exogenous variables in the partial VAR. With this criterion, the British pound (BP) and the Korean won (KW) are the only exchange rates in their respective groups which consistently appear weakly exogenous. The null hypothesis that BP or KW is weakly exogenous cannot be rejected across

[10] As pointed out by one of the referees, different lengths of the European and Asian crisis periods (6.5 and 12 months, respectively) may partially result in different cointegration findings. This is because the number of observations used in the tests (as well as the lag length chosen for the VAR) can affect the size and power properties of cointegration test statistics.

all possible r's. Hence, should cointegration exist, these two exchange rates must be one of the sources of common trends and do not respond to deviations from long-run equilibria. These results are also interesting in that BP was the first currency withdrawn from the EMS during the European crisis while KW was the last of the Asian currencies falling victim to the Asian crisis.

Based on the knowledge of weak exogeneity from the full VAR system, Table 4 sets forth the results from Johansen tests for the partial VAR system. Conditional on BP being weakly exogenous in the European group, $k = 2$ is needed for each equation for endogenous exchange rates to exhibit no serial correlation in residuals (see the first three columns in Panel A). With $k = 2$, both λ_{trace} and λ_{max} statistics consistently indicate one CIV among European exchange rates at the 5% level (see the first four columns in Panel B). These results are identical to those in the full VAR framework. Further, contingent on KW being weakly exogenous in the Asian group, k can be reduced from 9 in the full VAR to only 5 in the partial VAR. $k = 4$ is inappropriate because it would result in the statistically significant serial correlation in residuals at the 5% level for the Indonesian rupiah (IR) equation (see the last three columns in Panel A). Interestingly, unlike the full VAR system with no cointegration evidence, both λ_{trace} and λ_{max} statistics indicate one CIV among Asian exchange rates at the 10% level (see the last four columns in Panel B). Among other possibilities, these findings can result from an increase in the test power due to a decrease in the number of included lags *or* the possibility that the Johansen tests based on the λ_{trace} and λ_{max} statistics cause indeterminate or spurious findings of cointegration *or* the combination of the two. Hence, additional tests must be performed to ensure that the cointegration evidence is adequately robust.

Fig. 1 plots the recursive likelihood test statistics (normalized by the 5% critical value) for the stability of cointegration parameters obtained from the Johansen tests for the European currency group. The normalized test statistics for the full VAR and for the partial VAR (see Panels A and B, respectively) appear almost identical. As indicated by the plots which abruptly emerge above the critical value line of one in both VARs, the null hypothesis that the cointegration parameter is not statistically different from that derived over the full sample can be rejected for the recursive windows ending during January 1993. This instability is considerably more apparent in the group of Asian currencies with the plots of the same test statistics shown in Fig. 2. The null hypothesis that the cointegration parameter is not statistically different from that derived over the full sample can be rejected for the recursive windows ending during approximately the first half of the sample

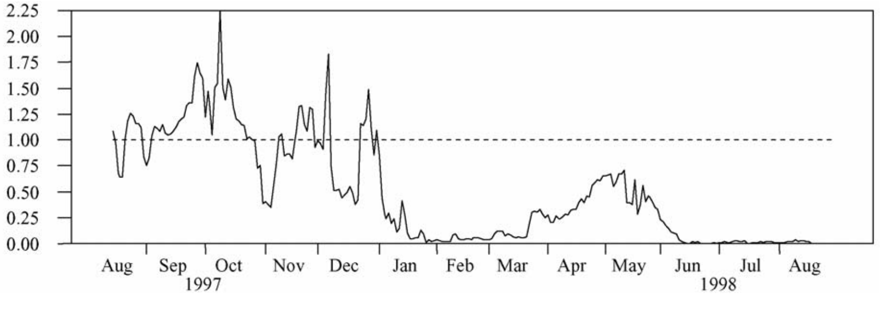

Fig. 2. Recursive test of CIV stability in the Asian group. *Note*: Conditional on the presence of one CIV in the partial VAR, the recursive likelihood ratio test statistics (scaled by the 5% critical value) are plotted against the *end* of each estimation window. The first estimation window is recursively increased until the last estimation window covers the full sample. The plot above the critical value line of one indicates rejection of the null hypothesis that the β estimate from the respective window is not statistically different from the β estimate derived from the full sample at the 5% level.

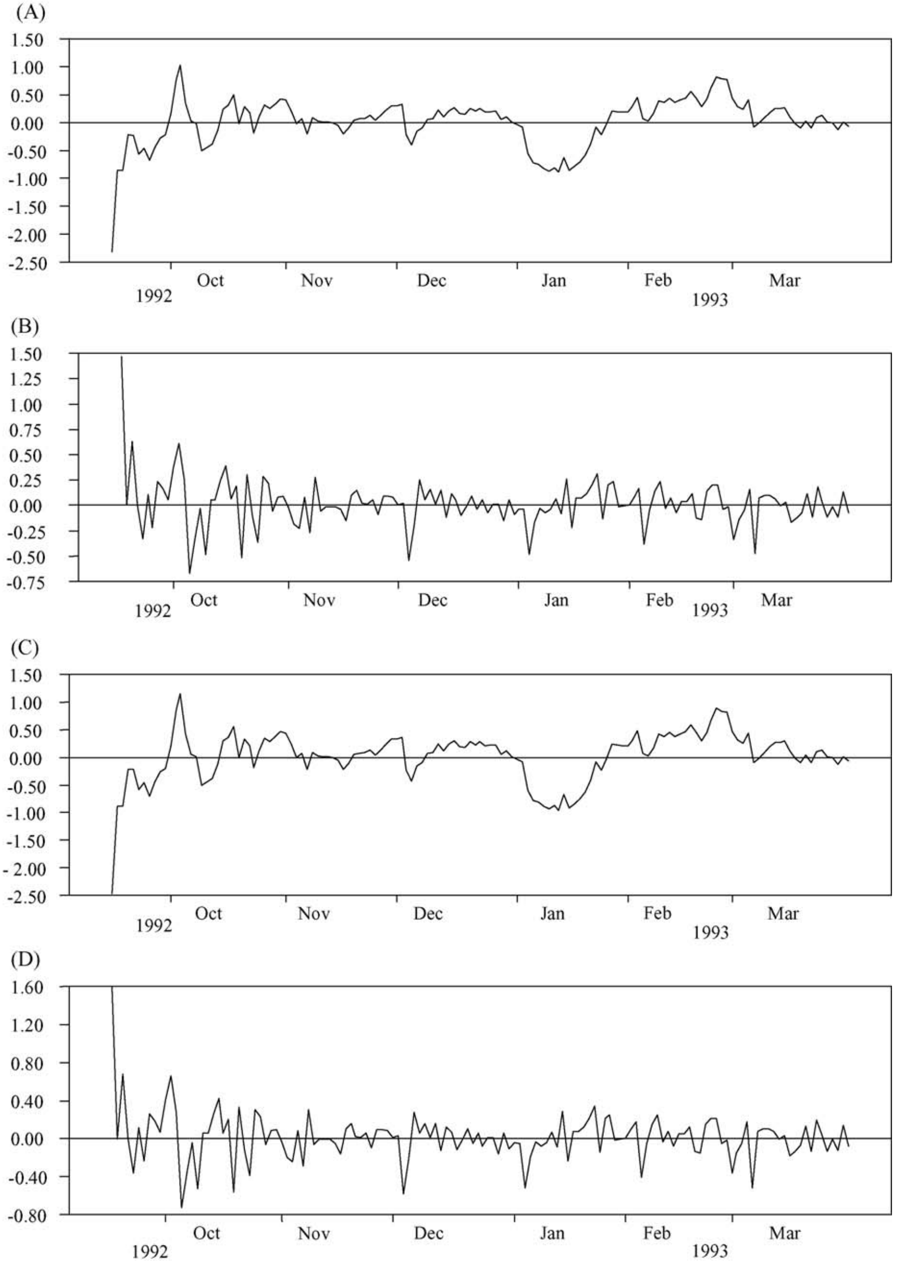

Fig. 3. Plots of CIV estimates in the European group. Panel A: CIV estimate in the full VAR system (in level); Panel B: CIV estimate in the full VAR system (in first difference); Panel C: CIV estimate in the partial VAR system (in level); Panel D: CIV estimate in the partial VAR system (in first difference). *Note*: The CIV estimate in the full VAR is 1.00BP +1.85FR −3.01GM −0.05IL +0.95 while the CIV estimate in the partial VAR is 1.00BP +1.87FR −3.06GM −0.04IL +1.00.

(i.e., August 1997–January 1998). Therefore, a cointegrating relation is not consistently stable during any of the two crisis periods.

Further, the plots of CIV estimates from the Johansen tests for the European group are shown in Fig. 3. The CIV estimates in levels from the full and partial VARs look virtually indistinguishable and more importantly non-stationary or $I(1)$ (see Panels A and C). This is in contrast to the plots of these estimates in first differences which look stationary or $I(0)$ (see Panels B and D). Additionally, the plots of the CIV estimate from the partial VAR for the Asian group where cointegration is evidenced are shown in Fig. 4. Similar to the European group, the CIV estimate in levels for the Asian group appears to be $I(1)$ while its first difference appears to be $I(0)$. This visual and informal assessment suggests that the ADF unit root tests of CIVs and common trend estimates (CTs) should be performed.

The unit root test results presented in Table 5 support the assessment prior. In the full VAR system in which one CIV and thus three CTs are found in the European group, the unit root null hypothesis cannot be rejected at any conventional level for both CIV and CT estimates. The non-rejection is consistent irrespective of the number of lags included in the ADF test equation (see Panel A). Further, the unit root null hypothesis cannot be rejected at any conventional level for both CIV and CT estimates in the partial VARs for European exchange rates and for Asian exchange rates. The non-rejection is also consistent irrespective of the number of lags included in the ADF test equation (see Panel B). Therefore, the estimates of CIVs which are stationary by definition show strong statistical evidence of non-stationarity as do the estimates of CTs. These findings substantially negate the presence of cointegrating relations during the two crises detected through the Johansen tests.

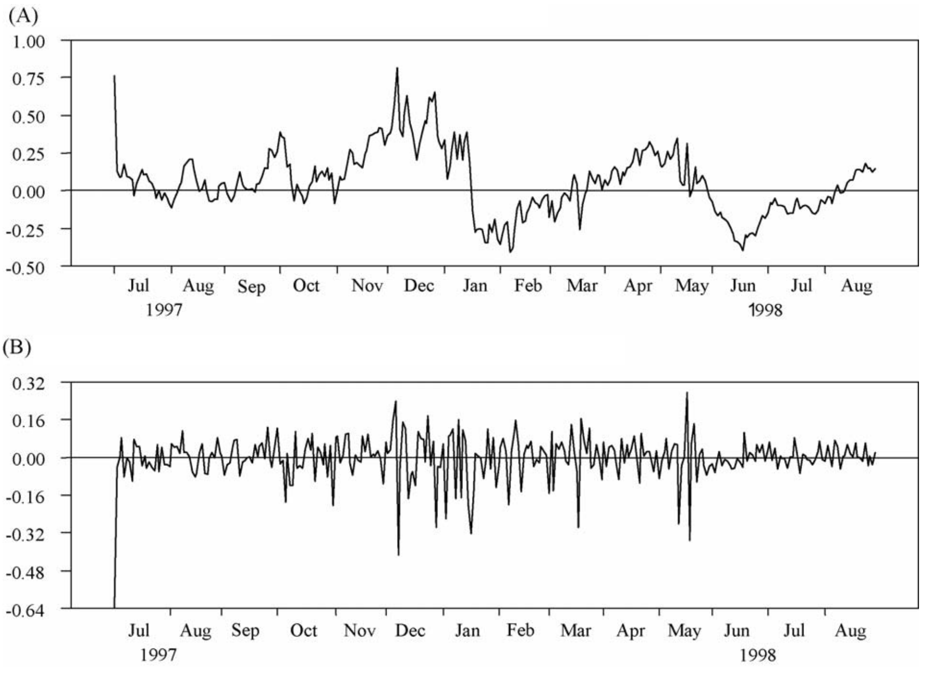

Fig. 4. Plots of CIV estimate in the Asian group. Panel A: CIV estimate in the partial VAR system (in level); Panel B: CIV estimate in the partial VAR system (in first difference). *Note*: The CIV estimate is 1.00IR −1.29KW −4.72MR +3.74TB +7.42.

Table 5
ADF unit root tests of CIV and common trend estimates

Lags	European crisis			
	CIV	CT1	CT2	CT3
Panel A: full VAR system[a]				
1	−3.73	−1.22	−2.00	−2.40
3	−3.36	−1.08	−1.33	−1.91
6	−3.04	−0.96	−1.23	−1.99
9	−2.76	−0.69	−1.40	−2.86
12	−2.62	−0.29	−0.45	−2.48

Lags	European crisis			Asian crisis		
	CIV	CT1	CT2	CIV	CT1	CT2
Panel B: partial VAR system[b]						
1	−3.73	−1.25	−1.31	−3.25	−1.01	−2.24
3	−3.39	−1.11	−1.01	−2.43	−0.75	−2.31
6	−3.02	−0.95	−0.87	−2.57	−0.75	−2.44
9	−2.70	−0.75	−0.64	−2.08	−0.81	−2.32
12	−2.68	−0.44	0.00	−2.40	−0.92	−2.32

Note: The ADF test statistic is under the null hypothesis of a unit root. Critical values obtained from MacKinnon (1996) are −3.8106 and −4.0964 at the 10% and 5% levels, respectively; * and ** *would* indicate rejection of the unit root null hypothesis at the 10% and 5% levels, respectively.

[a] "Lags" indicates the number of lags included the unit root test equation. The number shown is the ADF test statistic from the unit root test equation for the CIV estimate (CIV) or for each common trend estimate (CT1, CT2 or CT3) in the full VAR system.

[b] "Lags" indicates the number of lags included the unit root test equation. The number shown is the ADF test statistic from the unit root test equation for the CIV estimate (CIV) or for each common trend estimate (CT1 or CT2) in the partial VAR system.

Further, the HI cointegration test is performed for each exchange rate group with the results set forth in Table 6. In the European group, the null hypothesis of cointegration can be rejected at the 5% or 10% level irrespective of the choice of normalization (i.e., the selection of one exchange rate as a dependent variable). Similar findings are obtained in the Asian group as well. The null

Table 6
Harris–Inder cointegration tests

European crisis		Asian crisis	
Equation[a]	KPSS test statistic[b]	Equation[c]	KPSS test statistic[d]
BP	0.16*	IR	0.82**
FF	0.18**	KW	0.57**
GM	0.14*	MR	0.56**
IL	1.14**	TB	0.61**

Note: [a,c]"Equation" indicates the cointegration test equation in which the exchange rate listed is a dependent variable. [b,d]Obtained from the KPSS test of residuals from the equation, the KPSS statistic is under the null hypothesis of cointegration. Critical values, which are conditional upon the number of included variables and the sample size, are obtained from Sephton (1996). For the European group, critical values are 0.125 and 0.166 at the 10% and 5% levels, respectively. For the Asian group, critical values are 0.123 and 0.161 at the 10% and 5% levels, respectively; * and ** indicate rejection of the null hypothesis at the 10% and 5% levels, respectively.

Table 7
Gregory–Hansen cointegration tests

Equation[a]	ADF test statistic[b]	Level shift model[c]		Regime shift model[d]	
		ADF* test statistic	T_{PS}	ADF* test statistic	T_{PS}
Panel A: European crisis					
BP	−3.22	−4.05	12/18/1992	−3.80	12/18/1992
FF	−3.10	−3.81	11/18/1992	−3.96	11/18/1992
GM	−3.43	−3.91	11/18/1992	−3.86	11/18/1992
IL	−0.90	−3.24	10/29/1992	−3.24	10/29/1992
Equation[e]	ADF test statistic[f]	Level shift model[g]		Regime shift model[h]	
		ADF* test statistic	T_{PS}	ADF* test statistic	T_{PS}
Panel B: Asian crisis					
IR	−1.96	−3.50	01/01/1998	−3.67	02/06/1998
KW	−2.34	−4.49	11/25/1997	−4.51	12/03/1997
MR	−2.01	−3.13	03/10/1998	−4.15	01/14/1998
TB	−2.36	−4.16	02/26/1998	−2.34	09/18/1997

Note: [a,e]"Equation" indicates the cointegration test equation in which the exchange rate listed is a dependent variable. [b,f]The ADF test statistic is under the null hypothesis of no cointegration based on the conventional Engel and Granger procedure in which a possible shift in a cointegration relation is not considered. Critical values taken from MacKinnon (1996) are −3.8106 and −4.0964 at the 10% and 5% levels, respectively. [c,g]The cointegrating relation is allowed to have an endogenous shift in its intercept. [d,h]The cointegrating relation is allowed to have an endogenous shift in its intercept and slope coefficients. The ADF* test statistic is under the null hypothesis of no cointegration against the alternative hypothesis of a cointegrating relation with a structural shift. It is the smallest value of the ADF statistics calculated for all possible dates for a structural shift. T_{PS} is the possible structural shift date associated with the ADF*. Critical values obtained from Gregory and Hansen (1996) are −5.02 and −5.28 at the 10% and 5% levels, respectively, for the level shift model and −5.75 and −6.00 at the 10% and 5% levels, respectively, for the regime shift model; * and ** *would* indicate rejection of the null hypothesis of no cointegration at the 10% and 5% levels, respectively.

hypothesis of cointegration can be clearly rejected at the 5% level irrespective of the choice of normalization. Therefore, it is consistently convincing that a cointegrating relation is not present in either crisis period based on the reversed null hypothesis of cointegration.

Finally, the GH cointegration test which allows a structural shift in a cointegrating relation is conducted on each exchange rate group. The resultant ADF* statistics are shown in Table 7. For comparative purposes, the ADF test statistics from the conventional Engel–Granger procedure, which does not consider a structural shift, are also reported. The conventional ADF test statistics indicate non-rejection of the null hypothesis of no cointegration for every exchange rate equations. More importantly, the ADF* test statistics confirm that this non-rejection (as well as the instability and non-stationarity of the Johansen CIV estimates) does not result from failing to account for a possible structural shift in a cointegrating relation. In both groups, the ADF* statistics indicate that the null hypothesis of no cointegration cannot be rejected at any significance level. This finding holds true irrespective of the choice of normalization and whether the level shift or the regime shift is specified under the alternative hypothesis.[11] Hence, even when the possibility for

[11] For any one structural shift specification in either exchange rate group, the finding that T_{PS}'s associated with the ADF* statistics vary within only a few months of one another partially implies that the choice of normalization does not considerably alter the test results. For instance, for the level shift model in the European group, T_{PS}'s vary from 10/29/1992 (in the IL equation) to 12/18/1992 (in the BP equation).

a structural shift is considered, a cointegrating relation is not evident during either the European or the Asian currency crisis.

4. Conclusions

This study extends the study of foreign exchange market efficiency. It employs several verification testing procedures to re-examine if cointegration among different spot exchange rates is actually present during the 1992–1993 European currency crisis and during the 1997–1998 Asian currency crisis. This approach differs from prior studies which have relied primarily on conventional Johansen tests, have detected evidence of cointegration, and consequently, have inferred the existence of market inefficiency during these two crises.

The Johansen test results for the full and partial VAR systems collectively indicate that cointegration is present during the two crises, and specifically, that the cointegration evidence appears stronger during the European crisis than during the Asian crisis. However, additional tests cast strong doubts on the validity of these cointegration findings. The estimated cointegrating relations show evidence of instability and non-stationarity. This potentially negates the reliability and usefulness of the resultant ECM to predict the affected exchange rates during these two crises. Further, the reversed null hypothesis of cointegration in the Harris–Inder test can be clearly rejected. Finally, the null hypothesis of no cointegration in the Gregory–Hansen test cannot be rejected irrespective of the specification for a possible endogenous structural shift in a cointegrating relation.

The finding that the existence of cointegration during the European and Asian currency crises is in fact doubtful or unlikely has useful implications. First, even if foreign exchange markets are truly inefficient during these two crises, the inefficiency is not revealed through evidence of cointegration and thus a cointegration test does not appear to be useful or relevant in detecting such inefficiency. This conclusion is consistent with prior empirical suggestions that cointegration is a long-run property and long spans of data are needed to appropriately detect a cointegrating relation and that cointegration tests after all are not appropriate tests of market efficiency for any period. In fact, the absence of cointegration during the European crisis may simply be due to the departure of some EU currencies from the exchange rate mandates such as ERM and the Maastricht Treaty (1992).

Second, investors in foreign exchange markets should be discouraged from generalizing that cointegration among relevant spot exchange rates usually emerges during periods of economic uncertainty.[12] This caution should further prevent them from taking unnecessarily risky currency positions and hoping to gain superior profits based on cointegration and its implied predictability which in fact do not exist. Specifically, because a stationary and stable cointegrating relation is not present during either the European or the Asian crisis even if the possibility for a structural shift is considered, the cointegration-based model is unlikely one of the forecasting models which truly provide investors with arbitrage opportunities and superior returns. Thus whether or not other increasingly sophisticated econometric models (or otherwise basic but overlooked statistical

[12] To make further generalizations that Johansen tests lead to spurious cointegration findings for periods of economic certainty, the editor and one of the referees suggest that another recent crisis, the Mexican crisis of 1994–1995, be analyzed. The results which are not tabulated, but are available upon request, suggest that there exists a cointegrating relation between the U.S. dollar and Mexican peso during the Mexican crisis based on Johansen tests. However, additional test results clearly undermine the cointegration presence. The cointegration parameters are statistically unstable and the instability evidence is more persistent through time during this crisis than during the European and Asian crises.

measures such as correlations) can better detect the relationships among currencies and measure market inefficiency during the crisis period should be of interest to researchers.

Third, prior empirical evidence that the reliance on the Johansen test can result in spurious findings of cointegration and thus erroneous inferences is strongly supported. While the Johansen test has been found to show reasonable power and size properties, even when it is carefully implemented, additional tests are clearly needed to validate Johansen test results.

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