An Examination of the Asian Crisis Part I:
Regime Shifts in Currency and Equity Markets

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Abstract

We analyze a reduced-form structural relation between currency and equity markets in Indonesia, Malaysia, the Philippines, South Korea, Taiwan and Thailand during the recent Asian crisis using a new non-parametric technique for the identification of regime shifts. We examine both returns and return volatility time series and show how information shocks in these markets moved from country to country in the 1992-1998 time period. We find that volatility breaks occurred before return breaks: shifts in the volatility structure took place in the fall of 1997 for most of the countries in our sample, while the majority of the shifts in returns are concentrated in the first half of 1998. The sequential nature of the observed regime shifts is consistent with the notion that information spillover effects induced return and volatility shocks across Asian markets. In addition, we find that, after the estimated regime breaks occurred, all the Asian equity markets in our study became more responsive to the volatility of the corresponding domestic exchange rate. A companion paper analyzes the relation between these breaks and portfolio rebalancing activity by foreign investors.

JEL classification: C51; G15

Keywords: Regime shifts; Asian crisis; Information spillover; Exchange rates

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Abstract

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I. Introduction

To most observers the Asian currency crisis began in July 1997 with the devaluation of the Thai baht. It spread quickly through East Asia, although with widely differing intensity and duration, and eventually affected the global economy, pushing several emerging countries into a deep recession. However, prior to this date, a number of the Asian economies had already begun to stagnate. As well, several countries experienced further currency crises into 1998 and suffered a decline during the summer of that year in correspondence with the ruble crisis and huge capital flight out of most emerging markets.

Many popular explanations for the Asian crisis emphasize the relationship between equity and currency markets in each country (see Corsetti, Pesenti, and Roubini (1998) for an overview). A collapse in equity values led to an outflow of foreign investment, which exerted downward pressure on the domestic currency. From another perspective, currency devaluation, perhaps triggered by speculation, disrupted the domestic financial sector, then the entire domestic economy, finally bringing the equity market to a collapse. The magnitude of the losses in equity and currency markets during this period was in fact remarkable. Between June 1997 and September 1998, the Indonesian rupiah depreciated versus the U.S. dollar by 79%, while Indonesia’s equity index declined by 56%. The corresponding figures for Malaysia and the Philippines are 61% and 40%, and 67% and 34%, respectively (see Lin and Kuo (2000)).

One of the most significant questions raised by the events of 1997 and 1998 is the extent to which the crisis was the result of natural dependencies rather than an example of an irrational or “herding” phenomenon. Thus, the purpose of this study is to examine how equity and currency markets reacted during this crisis. Many studies of financial crises focus on the correlations among countries’ equity and currency returns before and after a crisis. If the correlations are significantly higher, a contagion effect is indicated, as in Bertero and Mayer (1990), and Lee and Kim (1993). But, as Forbes and Rigobon (2002) point out, this approach is flawed, since the standard tests of increases in covariance are biased toward acceptance.

Furthermore, Andersen, Bollerslev, Diebold and Labys (1999) show that, in foreign exchange markets, high volatilities are associated with high correlations. A common empirical approach in the literature uses ARCH or GARCH specifications for volatility and
return time series: for example, see Hamao, Masulis, and Ng (1990), and Karolyi and Stulz (1996). These applications use high frequency data to test for cointegration, and typically focus on the transmission of volatility across markets.

Our approach is different. We posit a simple, reduced-form structural relation between the equity and currency markets in each of six Asian countries - Indonesia, Malaysia, the Philippines, South Korea, Taiwan and Thailand. We select these countries for our study because they arguably experienced the greatest extent of financial, economic and political turmoil during 1997 and 1998, but also because each of these countries, prior to the crisis, was in fact adopting a pegged currency regime. More specifically, we introduce and estimate a linear relation between the equity returns in a given country and leads and lags of the corresponding domestic currency returns (relative to the U.S. dollar). We estimate a model for returns and for volatility separately in each country. We then test for a regime shift in these relations. When a shift occurs, the non-parametric technique we use, developed by Bai, Lumsdaine, and Stock (1998), allows us to determine confidence intervals around the break dates, and to do statistical inference into the nature of the break with minimal restrictions on the underlying data generation process.

Our data set comprises monthly return observations. The choice of lower frequency data mitigates the potential impact of infrequent trading effects on the statistical inference, as suggested by Harvey (1995). If there is a regime break in our selected time frame, then we can assume that a structural shift occurred between these markets. Analyzing the nature of this shift can help us understand its dynamics more precisely. For example, the shift could be the result of an adjustment in the means, a change in the lead-lag relation, etc. If instead no shift occurred between 1997 and 1998, then we assume that the crisis event is a manifestation of natural dependencies across these markets.

The patterns of potential regime breaks we observe during the Asian crisis allow us to address a second issue: Are these shocks consistent with the theoretical modeling of how various types of information events affect returns and volatility? The speculative trading model of Fleming, Kirby, and Ostdiek (1998) can be applied to portfolio allocation in multiple foreign markets. Their study provides a framework for analyzing how structural breaks in a single market can spill over to others in two possible ways. First, a common information shock generates trading activity and volatility in each market simultaneously.
Second, an information shock alters expectations in one market, resulting in investors adjusting their holdings in other markets. Portfolio rebalancing occurs because of the existence of correlations between returns, i.e., because of the ensuing changes in hedging demand. Fleming et al. call the latter case “information spillover.” Thus, our second goal includes testing whether the structural breaks occur simultaneously across all countries in our sample, or if they occur in sequence. If all markets experience structural breaks simultaneously, then common information shocks predominate. If instead the timing of regime shifts differs across markets in various countries, then cross-country information spillover effects are significant.

A further aspect of this issue is to compare the timing of shocks to volatility and to returns. It has long been argued in the financial literature that price and return volatility are related to the arrival of information. Merton (1980) develops a diffusion model showing that the variance of prices equals the rate of information flow. Shalen (1993) proves that new information generates speculative trade because it modifies the dispersion of beliefs among traders.1 Under the assumption that information can arrive at different times to different types of traders, Hirshleifer, Subrahmanyam, and Titman (1994) demonstrate that profit-taking behavior and herding should be more apparent in stocks that have more late-informed traders.

If the spillover of information is not instantaneous, we test if shocks in volatility will affect future returns. This issue is part of the extensive literature dealing with return predictability. In a study of the U.S. stock market, French, Schwert, and Stambaugh (1987) find evidence that a positive relation between expected equity volatility and risk premiums leads to a negative relation between the unpredicted component of volatility and excess returns. French et al. argue that if risk premiums are positively related to the predicted component of market volatility, then a larger-than-predicted volatility will not only result in an upward revision of future volatility, but will also increase the discount rate for future cash flows. The persistence of this shock in the predicted volatility depends on the particular return model adopted. The higher discount rate reduces the current stock price if cash flows do not adjust. Subsequent research (for example, Haugen, Talmor, and Torous, 1991, or

1 Shalen’ s work is related to the volatility clustering models of Bollerslev, Chou, and Kroner (1992), among others.
Brown, Harlow, and Tinic, 1993) shows that future stock returns are positively correlated to current shifts in volatility.

In our empirical setting, we thus wish to determine whether or not information shocks lead to shifts in the structure of the volatility relation before or after they affect the return structure. Providing an answer to these questions is of interest not only to researchers aiming to understand how financial crises propagate but also to policymakers and regulators hoping to predict and prevent future crises.

We find that regime shifts in the hypothesized return and volatility relations did occur in each of the six Asian countries in our dataset during 1997 and 1998. The estimated regime shifts between equity and currency returns differ widely, both in timing and nature, from country to country. We show that shifts in volatility lead shifts in returns. While regime shifts in returns take place almost a year after the currency crisis, volatility breaks occur around the crisis period. The shocks in the return relations affect Malaysia and Thailand first, and then the rest of Asia. Volatility breaks are instead concentrated in the fall of 1994 for Malaysia, Taiwan and Thailand, and in the last months of 1997 for the other countries in our sample. Our results also indicate that, after the estimated regime shifts occurred, most Asian equity markets became more responsive to the volatility of the corresponding domestic exchange rate. The sequential nature of breaks in the structural relationship between currency and equity market returns and volatility in various Asian countries is consistent with information spillover effects.

A final, important issue, the role of flows of funds during the Asian crisis, is addressed in a companion paper, Kallberg, Liu, and Pasquariello (2002). There we use data on the flows of funds in each of the countries in this study to see if the “herding” behavior of international investors played a significant role in the crisis. In particular, we test whether large-scale shifts in foreign investment were occurring at or around the time of the estimated structural breaks, and if changes in the flows of funds were responsible for any clustering in the estimated regime breaks.

The organization of the paper is as follows: Section II reviews the literature on the East Asian crisis that relates to our analysis. Section III illustrates the statistical methodology. Section IV describes the data set and examines our results. Section V concludes.
II. Explaining financial crises

Financial crises have generated many economic explanations and models; Calomiris (1995) presents an overview. Several plausible reasons for the Asian crisis and its propagation across world economies have emerged. The disruptive actions of hedge funds, the high volumes of foreign currency borrowing by Asian firms, the mishandling by the IMF and the Asian governments, and the excessive risk-taking by banks and local entrepreneurs are among the most frequently cited.

Many economists, e.g., Corsetti et al. (1998) and Allen and Gale (2000), have focused on the country “domino effect,” i.e., on the impact of the occurrence of a financial crisis in one country on other, not necessarily contiguous markets. Others, like Forbes and Rigobon (2002), downplay the contagion story and attribute the outcomes of the crisis to existing inter-market correlations. For example, as real estate markets plummeted, banks suffered enormous losses due to their exposures to real estate developers. These losses then spread to the rest of the financial sector (see Renaud (1997, 1999)).

Foreign exchange plays a central role in each of these possible explanations. Hence, the Asian financial crisis is usually labeled as a currency crisis, as in Park and Lee (2001). Choe, Kho, and Stulz (1999) argue that the currency crisis led to problems in the banking sector, as a response to deteriorating expectations about future economic activity. They document the extremely low returns to Asian banks between January 1997 and July 1998. High levels of dollar-denominated corporate borrowing further magnified the crippling role of foreign exchange during the crisis, as emphasized by Thurow (1998). Currency crises are often modeled on the basis of restrictive assumptions about government strategies (for example focusing on “unilateral pegs” and bilateral country structures). Buiter, Corsetti, and Pesenti (1998) analyze these flaws. One rationale underlying the currency and equity market contagion effect is that trade linkages provide a channel for the spread of financial problems, as in Glick and Rose (1999) and Dasgupta (2000).

Crisis, like bank runs, can also contain a self-fulfilling element. In this case, Obstfeld (1998) states that a sharp transition from equilibrium to a crisis state is more likely than a

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2 Thurow (1998) observes that by the end of 1997 most Korean private companies appeared to have borrowed some $160 billion in foreign currency, half in short-term loans.
smooth decline of the domestic economy. Allen and Gale (2000) model this effect. Because liquidity preference shocks are imperfectly correlated across geographical regions, banks hold interregional claims on other banks to provide insurance against these shocks. A small liquidity preference shock in one region can then spread throughout the economy. Allen and Gale find that the likelihood and intensity of contagion is highly dependent on the completeness of the market for interregional claims. Related to these models is the idea that the spread of financial crises is exaggerated because of incomplete information, as in Calvo and Mendoza (1997). Thus, a crisis in one country can signal information about the financial condition of other countries that share underlying regional factors and trade patterns. The widespread withdrawal of foreign funds, which creates pressure on the domestic currencies and interest rates, would then follow.³

Research that examines the transmission of information shocks across markets is relevant to understanding financial crises. The most significant paper for our study is Fleming et al. (1998), who analyze volatility shocks across debt, equity and money markets. Their work extends the single-asset, mean-variance model of Tauchen and Pitts (1983) to multiple markets, and investigates empirically the effects of cross-market dependencies. As previously mentioned, Fleming et al.’s main goal is to investigate volatility linkages of two distinct types. In the first case, a common information shock affects traders’ expectations in each market simultaneously. Reacting to this shock, traders adjust their speculative demand across markets.

In the second case (information spillover), an information shock perturbs expectations in one market, resulting in investors rebalancing their portfolios across other markets. With frictionless markets, the information spillover effect is complete and instantaneous. In these circumstances, we could not distinguish empirically between shocks that are attributable to common information or to an information spillover effect. However, practical considerations, i.e., market frictions, mitigate the intensity and speed of the information spillover effect, and permit its empirical identification. Transaction costs and institutional constraints, such as position and capital limits, restrict the trading activity of portfolio

³ In 1997, the Prime Minister of Malaysia, Dr. Mahathir Mohamed, called hedge fund managers the “highwaymen” of the global economy. See Chancellor (2000). In Kallberg et al. (2002) we explore in greater depth the role played by capital flight in the Asian crisis.
managers for both speculative and hedging purposes. These factors further reduce the managers’ ability to trade simultaneously in disparate markets. Differences in market depth may also reduce the speed with which an information shock spills over to other markets.

In the next section we model empirically the interaction among equity and currency markets as a (linear) structural relation between equity and currency returns, and between their corresponding volatilities. We then test whether a break occurred in any of these relations during the Asian crisis. The chronology of the estimated regime shifts will help us determine which kind of information shock affected the Far East markets in 1997 and 1998.

III. Methodology

Testing for regime breaks

In this section we describe the empirical method we use to test for the existence of shocks in the structural relation between equity and currency returns, and equity and currency volatility. By observing the timing and the intensity of the breaks in the hypothesized structural relations between returns and volatilities, we can test several implications of the crisis models discussed in the previous section. Although the problem of detecting break dates in economic and financial time series has received increasing attention in the literature, formal measures of the precision of these estimates could not be obtained.4 The estimation of a confidence interval around the break date is important to economists, because it incorporates a measure of sampling uncertainty in the analysis.

In this paper we adopt the statistical method of Bai, Lumsdaine, and Stock (1998), as it permits statistical inference about regime breaks, including interval estimation of the break date, with minimal restrictions on the underlying data generation process (DPG). In fact, Bai et al.’s non-parametric technique searches for a single break in univariate or multivariate time series models (with or without stationary regressors) assuming only that the DPG is a stationary vector autoregression (VAR) before and after the break, and specifies asymptotic confidence intervals for the estimated break point. We use this approach to test for regime shifts in the hypothesized linear relation between equity and currency returns, and equity and

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4 See, for example, Perron (1989) and Banerjee, Lumsdaine and Stock (1992).
currency return volatility, for each individual country.\(^5\) We start by applying the technique to a reduced-form model for returns. The model uses month \(t\) equity index returns denominated in local currency as the dependent variable. The independent variables are the currency returns in each of the five months \(t - 2\) to \(t + 2\) (to account for possible lead-lag relations) and the equity return in month \(t - 1\) (to account for first-order autocorrelation in returns). We perform a similar analysis on monthly volatility, estimated using a 12-month rolling window.\(^6\) We have the following equation:

\[
y_t = \mu + A y_{t-1} + \sum_{i=1}^{5} b_i x_{t+i-3} + d_t(k) \left[ \lambda + \alpha y_{t-1} + \sum_{i=1}^{5} \beta_i x_{t+i-3} \right] + \varepsilon_t, \tag{1}
\]

with \(d_t(k)\) equal to 1 if \(t\) is greater or equal to \(k\), and zero otherwise.\(^7\)

Here, \(k\) is a potential break date, \(y_t\) is the equity index return for a certain country in month \(t\), and \(x_t\) is the corresponding currency return in month \(t\). If we define \(S\) as a binary selection row vector with columns of ones corresponding to the parameters of Eq. (1) that are allowed to change, we can write Eq. (1) in a stacked form,

\[
y_t = V_t' \vartheta + d_t(k) V_t' S S \delta + \varepsilon_t, \tag{2}
\]

with \(V_t' = (1, y_{t-1}, x_{t-2}, ..., x_{t+2})\) and \(\vartheta = (\mu, A, b_1, ..., b_5)\) and \(\delta = (\lambda, \alpha, \beta_1, ..., \beta_5)\). In matrix form this is equivalent to

\[
y_t = Z_t' (k) B + \varepsilon_t, \tag{3}
\]

where \(Z_t' = (V_t', d_t(k) V_t' S')\) and \(B = (\vartheta', (S \delta)')\).

The model allows a wide range of assumptions about parameter shifts. The model is one of a full structural change if all coefficients are allowed to change. If we instead assume that only a subset of the coefficients undergoes a regime shift, then a partial structural model is more appropriate. For example, if we suspect a break in only the intercept, then we assume

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5 This technique was first applied by Bekaert, Harvey, and Lumsdaine (1998) to date the integration of world financial markets.
6 Officer (1973) and Merton (1980) were the first to use a rolling 12-month standard deviation. Our results are not sensitive to the particular averaging length.
7 It is important to observe that the above lead-lag specification does not necessarily imply a causal relation from the market for the domestic currency to the local equity market, or vice versa, in a country. Indeed, Eq. (1) also allows for the possibility that one market reacts to a common information shock faster than the other.
that \( S = [1, 0, 0, 0, 0, 0] \). As reported in Bai et al. (1998), tests for partial structural changes tend to have more power than those for full structural changes. However, our analysis indicates that the breaks are statistically significant when we use a full structural model rather than simply allowing the intercept to break.

A variety of tests for a break based on Wald statistics has been proposed in the literature. For a null hypothesis that \( S \delta = 0 \), for \( k = k_1^* + 1, \ldots, T - k_2^* \), where \( k_1^* \) and \( k_2^* \) are some trimming values,\(^8\) Bai et al.’s test, which is similar to that of Quandt (1958, 1960), considers the maximum of the following \( F \) process:

\[
\hat{F}(k) = T(R\hat{B}(k))'\left[ R T^{-1} \sum_{t=1}^{T} Z_t(k) \hat{\Sigma}_k^{-1} Z_t(k)' R' \right]^{-1} R(k)',
\]

where \( R = [0, I_r] \) is such that \( RB = S\delta \) and \( \hat{B}(k) \) and \( \hat{\Sigma}_k \) are the estimators for \( B \) and \( \sigma^2 \varepsilon \), respectively, evaluated under the alternative hypothesis, i.e., given \( k \). The estimated break date is then \( k^* \), the value that maximizes \( \hat{F}(k) \), and is statistically significant if \( \hat{F}(k^*) \) is greater than the corresponding critical value (for the selected significance level).\(^9\)

When we analyze a time series, if there is in fact a break, then the problem of constructing confidence intervals for the true break dates is not trivial. Various authors have examined this problem under the assumption of i.i.d. Gaussian errors.\(^10\) Bai et al. (1998) instead assume that the disturbances form a sequence of martingale differences with some moment conditions, and use limit theorems to construct asymptotic confidence intervals for the true break date, with coverage of at least \( 100(1 - \pi)\% \), of the following form:

\[
k^* \pm c \frac{1}{2} \left[ (S\delta)' S (\hat{\Sigma}_k^{-1} \hat{Q}) S (S\delta) \right]^{-1}.
\]

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\(^8\) Trimming is necessary for the model of Eq. (1) to be full rank before and after any possible break date \( k \). Because of the lead-and-lag structure of Eq. (1) and the definition of \( d_t(k) \), the trimming values used in our analysis are \( k_1^* = 10 \) and \( k_2^* = 9 \).

\(^9\) In order to compute critical values for \( \hat{F}(k) \), Bai et al. (1998) suggest to approximate the limiting distribution of the \( F \) process with partial sums of normal random variables for each possible dimension of the test statistic \( \hat{F}(k) \), i.e., for the rank of the selected vector \( S \). Bekaert et al. (1998) report one such table with critical values for dimensions up to 68, which we use for this research.

\(^10\) For a more complete review of the available literature on the topic, see Bai et al. (1998).
In Eq. (5), \( \hat{k} \) and \( \hat{\Sigma}_k \) are the estimated values for the break date and \( \sigma^2 \) respectively; \( c_{1/2} \pi \) is the \( 100(1 - \pi/2) \) quantile of the Picard (1985) distribution for \( \pi = 0.05 \). Next, we describe statistically the two regimes identified by the structural break in the return time series. More specifically, we estimate the hypothesized model assuming that \( \hat{k} \) is in fact the true break date, i.e., we estimate the parameters of the following regression model:

\[
y_t = Z_t (\hat{k}) B + \epsilon_t .
\]

Eq. (6) allows us to examine the effects of the identified structural break, if any, over the parameters of the structural relations between equity and currency returns.

We then focus on the second moments of returns, by extending the basic model of Eq. (1) to the case of equity and currency volatility. More specifically, we search for breaks in a structural relation between monthly rolled volatility for the available series of equity returns and the corresponding monthly rolled volatility series for currency returns. We calculate the rolled volatility as a moving standard deviation of 12 monthly returns for the equity \( (\sigma_y) \) and currency \( (\sigma_x) \) index returns. The proposed model is now

\[
\sigma_y = \mu + A(\sigma_{y,t-1} + b\sigma_{x,t} + d_t(k[B + \beta\sigma_{x,t}]) + \epsilon_t .
\]

Finally, as for returns in Eq. (6), we estimate Eq. (7) for each country at the measured break date to examine the effects of the regime shift over such reduced-form relation for volatility.

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11 Bai et al. (1998) investigate some of the finite sample properties of these test statistics. Using Monte Carlo simulations on I(0) and I(1) models, they show that the tests perform adequately, in terms of size and power, under the null hypothesis of no break and under the alternative hypothesis of a break in the mean of the DPG. Bai et al. also observe that, for a fixed break magnitude, the confidence interval does not depend on the sample size and, in particular, does not shrink for bigger sample sizes. Nonetheless, their results suggest that the asymptotic confidence intervals tend to be too tight, and that the performance of all the univariate intervals improves substantially when \( T \) is increased. Further exploration of such issues for more general structural break specifications, available in Bekaert et al. (1998), confirms that the size properties of the Wald test statistics are satisfactory in univariate models regardless of whether exogenous regressors are included in the analysis.

12 For this case, the trimming values we adopted are \( k_1 = 5 \) and \( k_2 = 4 \).

13 Note that, by construction, the rolled volatility series has positive serial correlation. We take this into account by including a lagged term in the regression model. However, since the focus of our analysis is on breaks in the relation between markets, and because most of the series’ persistence is induced by the way it is computed, we do not allow the serial correlation coefficient to vary in Eq. (7). Allowing the persistence term to vary after date \( \hat{k} \) results in mostly insignificant post-break coefficient’s changes, and does not affect the inference based on the reported Wald statistics.
As previously mentioned, many studies in the growing literature on financial crises use ARCH and GARCH models to analyze the behavior of volatility through time. GARCH models are especially attractive because of their ability to capture the time-varying nature of volatility with a limited number of coefficients. However, this approach is much less successful in capturing regime shifts in economic time series due to low-probability events, such as financial crises (see Susmel (2000) for a review). To overcome this shortcoming, Hamilton and Susmel (1994) extended ARCH specifications to account explicitly for potential structural changes in the underlying DGP.

Edwards and Susmel (2000, 2001) use a variant of this methodology, known as SWARCH, to test whether higher volatility regimes and increased co-dependence of those volatility regimes across countries characterize weekly interest rate and equity time series for a sample of Asian and Latin-American markets. Their results indicate that volatility co-movements are more likely among stock markets than among money markets. The SWARCH setting is, however, highly parametric, for it requires conditional normality of residuals and the formulation of a specific probability law (e.g., a Markov chain) causing the economy to switch regimes. Inference therefore becomes more problematic, since tests of the null hypothesis of no breaks are simultaneously tests of the validity of these assumptions. The technique of Bai et al. (1998) employed in this paper allows us to test whether a reduced-form relation between equity and currency return volatility underwent a statistically significant break over the selected interval while at the same time imposing the simplest structure possible on the DGP.

One of the main goals of this paper is to test whether the interaction between equity and currency markets was affected by the events of 1997 and 1998. The specifications selected to that purpose in Eqs. (1) and (7) result from a balance between the statistical significance of the estimated coefficients, economic rationale, and data availability constraints. Nevertheless, as emphasized by Rigobon (2001), stock market returns and exchange rate data, especially at high frequencies, appear to be “plagued with simultaneous equations, omitted variables, conditional and unconditional heteroskedasticity, serial correlation, non-linearity and non-normality problems.” However, autocorrelations in our dataset of equity and currency return series (not reported here) are either small or not statistically significant at the monthly
frequency we use in the analysis. Moreover, the non-parametric technique devised by Bai et al. has the attractive feature of not requiring normality in the residuals from Eqs. (1) and (7).

Albeit with these caveats, testing for the existence and timing of volatility and return shocks, i.e., for breaks in the posited linear relations can still help us shed some light on whether events observed in Asia in the past two years are attributable to either common information shocks or information spillover from one country to another. If common information shocks are associated with the Asian crisis, then we should see concurrent regime shifts in volatility and in returns across countries. On the other hand, if the timing of these regime shifts differs across countries, then we can attribute this to information spillover effects.

IV. Empirical results

Data

We use monthly time series of currency and equity returns in our analysis of Indonesia, Malaysia, the Philippines, South Korea, Taiwan and Thailand.

We assume that the exchange rate of each local currency relative to the dollar is the key exchange rate variable. Therefore, the monthly currency returns are the spot rates that correspond to the noon buying rate for cable transfers payable in foreign currencies, as recorded by the Federal Reserve Bank of New York every business day. For each of the six markets, we use the major equity index in local currency to calculate a monthly time series of local returns: JCI Jakarta Composite Index, EMAS Equity Index, PSE Index, Kospi 200, the Thailand Stock Exchange Index and the TWSE Stock Index. Monthly equity indexes for each country are from Bloomberg.

A variety of reasons justify the composition of our dataset. Firstly, Indonesia, Malaysia, the Philippines, South Korea, Taiwan and Thailand arguably suffered the greatest degree of economic, financial, and political turmoil during 1997 and 1998. A brief chronology of the crisis is reported in Table 1. These economies, except Taiwan, were explicitly entitled by the Asian Development Bank (ADB) to access emergency funding facilities made available by Japan in the fall of 1998 (the "New Miyazawa Initiative") in recognition of the severe impact
of the currency crisis on their economic and financial systems. Secondly, those countries were all unilaterally pegged to the U.S. dollar by relatively tight fluctuation bands around the time of the Asian crisis. Thailand abandoned its fixed exchange rate regime on July 2, 1997. In the Philippines, the Central Bank was forced to relax its previously successful band of fluctuations for the peso by the end of July 1997. Indonesia gave up a similar enlarged currency band on August 14, 1997. The Indonesian rupiah crashed soon afterwards. South Korea renounced battling the increasing selling pressure on the won in November of 1997, as did Malaysia.

As mentioned in Section I, the choice of monthly data is also not casual, and is consistent with previous applications of the technique used in this paper to detect the occurrence of structural breaks in economic and financial time series (e.g., Bai et al. (1998), Bekaert et al. (1998) and, more recently, Bekaert, Harvey, and Lumsdaine (2002)). Employing weekly or even daily equity and currency data for the identification of structural breaks raises in fact concerns related to the possibility of biases induced by infrequent and/or nonsynchronous trading. Although the trading activity of many of the emerging markets is surprisingly intense, Harvey (1995) suggests that it might be appropriate to use monthly rather than weekly time series to mitigate the possible influences of these biases and short-term noise on the resulting statistical inference. Furthermore, as reported by Bai et al. (1998) (and summarized in footnote 10 of Section III), Monte Carlo analysis of the finite sample performance of the proposed estimator for \( \hat{k} \) reveals little or no change in the estimated confidence intervals for increased frequency and constant sample length \( T \); substantial improvements in the precision for the estimated break date appear to be obtained only for constant frequency and increased \( T \).

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14 The international political isolation of Taiwan in the rest of Asia may explain its exclusion from the “New Miyazawa Initiative”.
15 Grier and Grier (2001) find that currency depreciation and equity price declines were more pronounced than what could have been justified by existing macroeconomic fundamentals in developing countries that started 1997 with an exchange rate peg.
16 Taiwan was also hit by the crisis, although less severely. Taiwan was able at first to fend off expectations of increased depreciation, through its Central Bank’s policy of active foreign exchange intervention. Therefore, the country initially appeared to be unaffected by the economic instability undermining the rest of the region. However, following increased pressures on the domestic currency, a steep rise of short-term interest rates and significant losses of foreign reserves, on October 17, 1997 Taiwan decided to adopt a floating exchange rate regime. The New Taiwan Dollar subsequently fell by 30% versus the U.S. dollar. A severe decline of the stock market ensued. After rising by 13.15% in 1997, while many other Asian equity markets were plummeting, the TWSE Stock Index in fact fell by more than 21% in 1998.
Basic statistics

Table 2 presents summary statistics for each country. Sample periods end in March 1999. Mean returns and sample volatilities reveal a wide disparity across the region. Mean monthly equity returns are negative in the Philippines (−0.29%) and Thailand (−0.67%). Table 1 also reports (not surprisingly) negative returns on currencies over this period. Correlation matrices for currency and equity returns, and their corresponding volatility time series (not reported here) characterize the interaction among the six countries we consider. The correlations for equity returns are mostly positive, ranging between .306 and .691. The currency return correlations are similar, although somewhat lower than those obtained for equity index returns. These facts seem to indicate that the markets and the asset classes in our study are only moderately correlated.17 Thus, there appear to be ex-ante hedging and portfolio rebalancing opportunities for fund managers investing in the region.

Volatility correlations across the six countries tend to be much more significant (and higher) than the return correlations, ranging from .507 to .969. We also observe a very wide dispersion of values, which suggests that the lead-lag relations among these countries are complex. Correlations between equity and currency returns are all quite low, from a minimum of −.206 to a maximum of .408. This confirms our previous observation that diversification effects should be strong, and that the existence of common factors in these returns is not apparent. As before, correlations between equity and currency volatility show higher values than those between the corresponding return time series. Most of the countries have volatility correlations higher than .636.

Analyzing regime shifts

We start by testing for breaks in the assumed structural relationships between local equity returns and local currency returns (Table 3), and between the volatility of local equity returns and the volatility of local currency returns (Table 4). As an example of the analysis underlying Tables 3 and 4, we report in Figures 1A to 1D the time series of equity and currency indexes (A), their rolled volatility series (B), and the corresponding Wald statistics.

17 This conclusion is even stronger if we take into account the fact that those matrices are affected by the increase in correlation that followed the events of 1997 and 1998. For an analysis of the behavior of correlations in Asia before and after 1997, see Baig and Goldfajn (1998).
(C and D), computed according to Eq. (4) in Section III, for Thailand. Finally, Figures 2 and 3 display the confidence intervals at the 5% level around those estimated break dates \( \hat{k} \), ranked in order of increasing statistical significance, over two time intervals: March to December 1997 and January to October 1998. These two intervals correspond to most of the observed break dates for equity versus currency volatility (Figures 2b and 3b) and equity versus currency returns (Figures 2a and 3a).

Tables 3 and 4 (and Figures 2 and 3) contain the main features of our analysis of regime shifts in returns and volatility for equity indexes versus their corresponding exchange rates. The tables report the median break point date and its associated significance. For the relation between equity and currency returns in Table 3, we find that all the countries in our sample experience a statistically significant regime shift (at the 5% level). More importantly, all the shifts (except Taiwan) occur from February to June 1998. The shocks appear to affect Malaysia and South Korea first and then quickly move to the other countries. For example, the shift in the parameters of Eq. (1) for Thailand is statistically significant in March of 1998, in the aftermath of the devaluation of the baht but before the Russian default (see Table 1). Our results also offer some support to the notion (supported, among others, by Wu (1998)) that financial instability in Taiwan during the Asian crisis was less acute than in the rest of the region.

The picture offered by the analysis of the relation between equity and currency return volatility in Table 4 is very different. The estimated regime shifts are more dispersed in time than in the case of returns. Structural breaks in Eq. (7) affect Malaysia, Taiwan and Thailand in October 1994. Around the time of the currency crisis, regime shifts involve Indonesia and South Korea. Finally, in August 1998 the Philippines undergo a switch.

An analysis of the pre- and post-break regressions, given in Tables 5 and 6, illuminates the nature of these regime shifts and of the lead-lag relationships in these countries. To correct for heteroskedasticity and autocorrelation in the residuals of the regressions (1) and (7), statistical significance of the coefficients is measured using Newey-West standard errors. As discussed in Section II, the post-break coefficients represent incremental changes to the pre-break coefficients. The data in Table 5 formalize the regime shift analysis for volatility in Table 4. The posited structural relations fit the data relatively well; adjusted \( R^2 \)'s range from 83% (Taiwan) to 97% (Malaysia). The most complete regime shift seems to have occurred in
Malaysia, Taiwan and Thailand, where both the intercept term ($\mu$) and the coefficient for the contemporaneous relationship between equity and currency return volatility (A) in Eq. (7) experience a statistically significant change. The most startling result from Table 5 is the uniformity across countries in the sign of $\beta$, the change in the contemporaneous linear relationship between equity and currency return volatility after a break occurred. All the coefficient changes are positive, and most of them are statistically significant. This indicates that, following the estimated regime shifts, equity markets became more responsive to the volatility in the corresponding domestic exchange rate. However, because the breaks occur quite late in our sample period, only few of the coefficients are statistically significant. In Table 6, which reports parameter estimates for shifts in the return relations, we observe less pronounced structural shifts than the ones recorded for the volatility relations. The adjusted $R^2$s now range from 37% (for Taiwan) to 63% (Philippines), with a much lower average. Again, since many of the regime breaks in returns occur quite late in our sample period (but also because of the presence of multicollinearity), some of the individual coefficients’ changes had weak statistical significance, and did not suggest any consistent interpretation.18

In short, Tables 3 to 6 reveal that for three of the countries in our sample, Indonesia, the Philippines and South Korea, where the greatest extent of political turmoil ensued from the Asian crisis, where the domestic central banks strictly controlled the corresponding currencies, and where market regulations were poor or nonexistent, breaks in the reduced-form relations described in Eqs. (1) and (7) do occur in 1997 and 1998 for both returns and return volatility. Regime shifts in the lead-lag relation between equity and currency returns were registered for Malaysia, Taiwan and Thailand as well. Nonetheless, for those three countries breaks in return volatility cluster by the end of 1994, before the Mexican Peso crisis. This evidence suggests that the increase in the volatility of stock and currency returns observed in these economies during the Asian turmoil was a manifestation of natural

\[18\] This issue arises because, when the Wald statistic identifies a break late in the sample period, there are fewer observations with which to estimate the seven post-shift coefficient changes in Eq. (1). Indeed, in our analysis, the breaks in the hypothesized structural relation between equity and currency returns occur in the first few months of 1998 for 5 of the 6 countries in the sample. The problem is somewhat less serious in the case of Eq. (7), where there are just two post-break parameters to estimate. In addition, multicollinearity affecting the many explanatory variables in Eq. (1) might also obscure their individual contribution to the fit. In short, the resulting post-shift estimates for the parameters in Eq. (1) may be statistically and economically insignificant, even if the joint effect of the break dummies is significant. Therefore, these coefficients are not reported in Table 6, but are available from the authors on request.
dependencies among the domestic equity and exchange rate markets rather than the result of a sequence of regime shifts in their interaction.

Furthermore, breaks in the hypothesized structural relations between equity and currency return volatility appear to consistently anticipate the corresponding break events in the returns’ relations for most of the countries in our sample, even after accounting for the width of the confidence intervals around the estimated dates \( \hat{k} \) in Tables 3 and 4. For instance, in Indonesia the reduced-form linear relation between the volatility of the domestic stock market and the volatility of the corresponding exchange rate breaks in a statistically significant fashion in June of 1997, ahead of the devaluation of the Thai baht and a full month before the local Central Bank eventually abandoned the rupiah trading band (as reported in Table 1). Nonetheless, the relation between the corresponding equity and currency returns does not shift until June of 1998, in the aftermath of the Asian crisis and before the Russian default and the subsequent collapse of Long Term Financial Management (LTCM). This evidence is even stronger when we consider that the time series for the volatility of equity and currency returns used in the analysis are rolled 12-month-window moving averages, hence by construction slow in absorbing shocks.

These results are consistent with the empirical findings of Haugen, Talmor and Torous (1991). In the context of U.S. markets, they show that regime shifts in volatility, whether sudden or protracted over several months, induce adjustments in the level of stock prices and realized returns. An intuitive explanation of this phenomenon is that any information shock that affects the volatility of a single market’s equity index returns, the volatility of the exchange rate, or the relation between the two, will eventually induce a shock in returns through the adjustment of the predicted component of the index volatility itself. Significant cross-country correlations then channel the information shock that originates in a specific country into other markets. Market frictions, differences in economic fundamentals, and constraints to hedging make the spillover phenomenon more or less pronounced.19

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19 The available empirical evidence on the state of the economies involved in the Asian crisis (see Eichengreen, Rose, and Wyplosz (1996), and Corsetti et al. (1998)) leads us to exclude the possibility that idiosyncratic shocks occurred in each of the countries for which the Wald statistic of Eq. (4) is statistically significant, hence that such shocks would explain the chronology of regime shifts in the posited reduced-form relations between equity and currency markets (in Tables 3 and 4).
The sequential nature of the estimated structural breaks appears to suggest that information spillover effects created return and volatility linkages between Asian markets. Information alters expectations in one market and affects returns and volatility in other markets through changes in hedging demand. The effect of these changes is protracted and delayed by the existence of market frictions. In a companion paper, we test more explicitly for cross-country information spillover and explore in greater detail the potential role of portfolio rebalancing and herding in the transmission of such volatility shocks across East Asia.

Testing for volatility spillover

A substantial issue still remains unresolved. Does the observed sequence of structural breaks imply a causal relation from volatility to returns? In their study of the U.S. equity market, French et al. (1987) argue that larger-than-expected shocks to volatility may induce upward revisions of future volatility predictions, eventually increasing risk premiums. We extend the argument of French et al. (1987) to emerging markets, where the foreign investors’ component of the total investment is especially significant and the cash flows of export-oriented firms (which typically represent a relevant percentage of the local equity indexes) are affected by fluctuations in local currencies.

Unexpected shocks to the volatility of the domestic market index might result from an unexpected shock to the volatility of the local currency equity index, to the volatility of the exchange rate, or to the correlation between the two. For example, an increase in currency volatility makes the cash flows of export-oriented firms more uncertain. Higher equity volatility could instead raise the discount rates for those cash flows. Their combined effect could induce an upward revision of both predicted currency and equity volatility, therefore causing stock prices in local currency to decline and affecting returns in both domestic and foreign currencies.

We study this dynamic interaction between returns and volatility using the timing information on the break events contained in Tables 3 and 4. The chronology of the structural breaks in volatility allows us to test for the existence of such information spillover by estimating the impact of the break event in country i at time t on the structural relation itself. We use the estimated break dates to test whether the chronologically sequential relation between volatility breaks and return breaks we found in Tables 3 and 4 reflects a
more fundamental interaction between the two underlying hypothesized structures of Section III. Thus, we estimate the following regression:

\[ y_t = \mu + A y_{t-1} + \sum_{i=1}^{5} b_i x_{t+i-3} + d_t(\hat{k}) \left( \lambda + \alpha y_{t-1} + \sum_{i=1}^{5} \beta_i x_{t+i-3} \right) + \varepsilon_t + \gamma d_t(\hat{k}_{VOL}) \]  \hspace{1cm} (8)

Eq. (8) tests whether the statistical power of Eq. (1) for each country can be improved by including a dummy for break dates in the corresponding reduced-form relation between equity and currency return volatility. In the above expression, \( d_t(\hat{k}_{VOL}) \) is equal to zero when \( t \) is below the lower limit of the confidence interval for the corresponding volatility break, and one otherwise. If volatility shocks were at least partially responsible for the break events in returns, we would expect the estimated \( \gamma \)s to be statistically significant. Moreover, if such shocks depress future return, we would expect the estimated \( \gamma \)s to be negative. Table 7 reports the results of the estimation of Eq. (8). We include all the countries in the experiment, although for some of them the registered volatility break occurred after the return event. This allows us to test for the robustness of our analysis. Indeed, when the lag between return and volatility breaks is large and positive, we expect the resulting estimated \( \gamma \)s to be insignificant. For example, in the case of Taiwan, where the structural relation between equity and currency volatility breaks almost one year after the corresponding relation for returns, the resulting \( \gamma \) is not significantly different from zero.

However, for most the countries in our sample for which the break in volatility precedes the one in returns the estimated \( \gamma \)s are negative and, in many cases, significant, with the exception of South Korea, where \( \gamma \) is instead positive but statistically insignificant. That the evidence of a causal relation from volatility to return breaks is weaker for South Korea has two, not necessarily conflicting explanations. First, the confidence interval around the estimated volatility break date for South Korea (in Table 4) encompasses the events that eventually led the country to abandon its defense of the won, as shown in Table 1. Nonetheless, albeit significant at the 1% level, the Wald statistic corresponding to such regime shift in the reduced-form volatility relation of Eq. (7) is the lowest (and, as is clear from Figure 2b, the confidence interval around \( \hat{k} \) is the widest) among its East Asian peers.

\[ \text{\footnotesize 20 We are grateful to Robert Dittmar for suggesting this analysis of the break chronology.} \]
many of which were at the same time experiencing more extreme forms of economic and political turmoil. Additionally, the Kospi 200 stock market index managed to recover during the first quarter of 1998, which is when a break in the linear relation between equity and currency returns is detected (March 1998, in Table 3). This should not be surprising, given the observation that the Korean economy was, and still is among the most developed, resilient and dynamic of the region, and given the massive bailout (U.S. $53 billion) the country received from the IMF in the fall of 1997 (see Bank for International Settlements (1999) and Table 1).

Hence, our evidence is consistent with the results of French et al. (1987) for the U.S. markets. Indeed, in our sample, past regime shifts in the linear relation between equity and currency return volatility appear to have induced a statistically and economically significant decline in equity returns, even after we controlled for the occurrence of a break in the relation between equity and currency returns of Eq. (1). In particular, volatility shocks affected future returns for most of the Asian nations that experienced structural shifts in the posited relations between equity and currency markets during 1997 and 1998.

V. Conclusions

Foreign exchange markets played a central role in the events that occurred in Asia in 1997 and 1998. Our study investigates the structural relation between currency and equity markets in six Asian countries, Indonesia, Malaysia, the Philippines, South Korea, Taiwan and Thailand, around the time of the crisis. For each country, we specify reduced-form linear relations between equity index returns and lead-and-lag currency returns. We choose the lead-and-lag specification to allow the data to determine whether and how one market anticipates or follows another. We adopt similar specifications to describe the structural relation between equity and currency return volatility. We apply the non-parametric statistical technique of Bai, Lumsdaine, and Stock (1998) to test for the existence of a single structural break in the posited relations between the time series of equity and currency returns and return volatility, respectively. We also compute confidence intervals for the estimated break dates for each country.

We find that the estimated regime shifts between equity and currency returns differ widely, both in timing and nature from country to country. The shocks in the return
relations appear to affect Malaysia and South Korea first, in early 1998, and then move to the rest of Asia. Volatility breaks instead cluster during the Mexican Peso crisis, in the fall of 1994, for Malaysia, Taiwan and Thailand, and at the time of the first Asian currency crisis, in the last months of 1997, for the other countries in our sample. Our analysis shows that for all the countries in the sample the corresponding domestic equity markets became more responsive to the volatility in currency markets after the occurrence of the regime breaks.

We use the chronology of breaks implied by our analysis to distinguish whether market linkages between countries are due to common information shocks or to information spillover effects. Most of the estimated breaks in the hypothesized return and return volatility structural relations are not simultaneous, even when the uncertainty surrounding our estimates is taken into account. The sequential nature of these regime shifts indicates that information spillover effects created linkages across the Asian markets during 1997 and 1998. Information alters expectations in one market and affects returns and volatility in other markets through changes in hedging demand. The effect of these changes is protracted over time and delayed by the existence of market frictions.

We examine the related issue of whether the negative relation between volatility and returns found by French et al. (1987) for the U.S. market holds for Asian markets. We find that not only is there a negative relationship between volatility and returns in the data but also the breaks in volatility appear to have anticipated the breaks in returns. Moreover, the observed sequence of breaks appears to imply a causal relationship from volatility to returns. We in fact show that past volatility shocks negatively affected future returns for most of the nations that experienced regime shifts in the posited relations between equity and currency markets during the crisis.

Many studies (e.g., Fleming et al. (1998) and, more recently, Kodres and Pritsker (2002)) suggest that portfolio rebalancing was a major channel for transmitting information shocks across markets during 1997 and 1998. In a companion paper, we address the closely related issues of whether the sequence of breaks estimated here implies a causal relationship across countries, and of whether cross-country spillover at or around the time of the Asian turmoil can be attributed to herding behavior by foreign investors.
References


Table 1
A chronology of the Asian Crisis, 1997-1999

This table, from the authors, Bank for International Settlements (1999) and Kaminsky and Reinhart (2000), reports selected significant economic, political and financial events that occurred between 1997 and the first three months of 1999.

<table>
<thead>
<tr>
<th>Date</th>
<th>Event</th>
</tr>
</thead>
<tbody>
<tr>
<td>May 1997</td>
<td>The Thai baht comes under strong depreciation pressure</td>
</tr>
<tr>
<td>July 2, 1997</td>
<td>Devaluation of the Thai baht</td>
</tr>
<tr>
<td>End of July</td>
<td>The Philippines Central Bank relaxes its fluctuation band for the peso</td>
</tr>
<tr>
<td>August 14, 1997</td>
<td>Indonesia abandons the rupiah trading band</td>
</tr>
<tr>
<td>Fall of 1997 (1)</td>
<td>The IMF offers U.S. $1 billion to the Philippines to rescue the peso</td>
</tr>
<tr>
<td>Fall of 1997 (2)</td>
<td>An aid fund of U.S. $16 billion is granted by the IMF to Thailand</td>
</tr>
<tr>
<td>October 17, 1997</td>
<td>Taiwan’s Central Bank adopts a clean floating foreign exchange regime</td>
</tr>
<tr>
<td>October 20, 1997</td>
<td>The Hong Kong dollar (HKD) falls victim to speculation</td>
</tr>
<tr>
<td>October 20, 1997</td>
<td>The Taiwanese dollar (NTD) depreciates by 5% against the U.S. dollar</td>
</tr>
<tr>
<td>October 24, 1997</td>
<td>The overnight interest rate in Hong Kong soars from 5% to 300%</td>
</tr>
<tr>
<td>October 27, 1997</td>
<td>The Hang Seng index plunges; panic selling in New York and Europe</td>
</tr>
<tr>
<td>November 1997</td>
<td>Indonesia is granted U.S. $23 billion by the IMF</td>
</tr>
<tr>
<td>November 17, 1997</td>
<td>South Korea abandons its defense of the won</td>
</tr>
<tr>
<td>November 18, 1997</td>
<td>U.S. and Japan meet with South East Asian countries in Manila</td>
</tr>
<tr>
<td>November 22, 1997</td>
<td>The Korean government formally asks the IMF for bail out</td>
</tr>
<tr>
<td>December 1997</td>
<td>IMF launches the biggest international aid plan in history for Korea</td>
</tr>
<tr>
<td>December 4, 1997</td>
<td>Markets question the plan; the won falls to nearly 2000 for a U.S. dollar</td>
</tr>
<tr>
<td>July 6, 1998</td>
<td>Salomon Brothers dismantles its bond arbitrage desk</td>
</tr>
<tr>
<td>July 20, 1998</td>
<td>First Wall Street Journal headlines on LTCM losses</td>
</tr>
<tr>
<td>August 17, 1998</td>
<td>Russian effective default and ruble devaluation</td>
</tr>
<tr>
<td>September 1, 1998</td>
<td>Malaysia introduces capital controls</td>
</tr>
<tr>
<td>September 2, 1998</td>
<td>LTCM shareholder letter, announcing the fund’s collapse, is issued</td>
</tr>
<tr>
<td>September 23, 1998</td>
<td>LTCM recapitalization</td>
</tr>
<tr>
<td>October 15, 1998</td>
<td>Inter-meeting Federal Reserve rate cut</td>
</tr>
<tr>
<td>Fall 1998</td>
<td>Japan launches the “Miyazawa Initiative” to rescue East Asian countries</td>
</tr>
<tr>
<td>January 10, 1999</td>
<td>China refuses to help foreign creditors of GITIC; markets disrupted</td>
</tr>
<tr>
<td>January 13, 1999</td>
<td>Fears of debt crisis in China sweep through Hong Kong; Brazil devalues</td>
</tr>
</tbody>
</table>
Table 2
Descriptive statistics
This table displays descriptive statistics (mean and standard deviation) for time series of monthly equity indexes and local currency returns for each of the countries included in the study over different intervals of our sample. Equity returns are computed from local equity indexes' monthly time series obtained from Bloomberg. Currency returns are calculated from the exchange rate versus the U.S. dollar. The minus sign represents devaluation of the local currency. Currency data are from the Federal Reserve Bank of New York.

<table>
<thead>
<tr>
<th>Country</th>
<th>Equity returns</th>
<th>Currency returns</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean</td>
<td>Stdev</td>
</tr>
<tr>
<td>Indonesia</td>
<td>0.06%</td>
<td>11.88%</td>
</tr>
<tr>
<td>Malaysia</td>
<td>0.19%</td>
<td>10.19%</td>
</tr>
<tr>
<td>Philippines</td>
<td>-0.29%</td>
<td>10.06%</td>
</tr>
<tr>
<td>South Korea</td>
<td>0.21%</td>
<td>10.52%</td>
</tr>
<tr>
<td>Taiwan</td>
<td>1.02%</td>
<td>9.38%</td>
</tr>
<tr>
<td>Thailand</td>
<td>-0.67%</td>
<td>11.84%</td>
</tr>
</tbody>
</table>
Table 3
Analysis of equity indexes monthly returns: structural break in all parameters

This table reports estimated break dates $\hat{k}$ for the structural relation between equity and currency returns

$$y_t = \mu + A y_{t-1} + \sum_{i=1}^{s} b_i x_{t-i-3} + d_t(k) \left[ \lambda + c y_{t-1} + \sum_{i=1}^{s} \beta_i x_{t+i-3} \right] + \epsilon_t \quad . \tag{1}$$

The Median column in the Table shows the estimated break date $\hat{k}$ for the time series of monthly equity returns in local currency for each of the countries in the sample. Break dates are estimated using the Wald statistic $F$ described in Eq. (4). We test the null hypothesis that the post-break coefficient changes are not significantly different from zero, i.e., that no break occurred in the sample period, by comparing the maximum value in the estimated time series $\hat{F}(k)$ to the 5% quantile of its limiting distribution. The null hypothesis is rejected when the maximum value for $\hat{F}(k)$, reported in the Max-Wald column of the table, is lower than the critical value for the selected significance level. Bekaert et al. (1998) compute a table with critical values for max $\hat{F}(k)$ (for dimensions up to 68), approximating the limiting distribution of the $F$ process with partial sums of normal random variables for each possible dimension of the test statistic $\hat{F}(k)$, i.e., for the rank of the selected vector $S$. From such table we use the asymptotic 5% and 1% critical values of 27.02 and 22.21, respectively, corresponding to a rank of 7 for the vector $S$ in Eq. (2). The 2.5th and 97.5th Percentile columns display estimated lower and upper bands, respectively, for the confidence intervals for the “true” break dates, computed according to Eq. (5) with quantiles of the Picard (1985) distribution.

<table>
<thead>
<tr>
<th>Country</th>
<th>2.5th Percentile</th>
<th>Median</th>
<th>97.5th Percentile</th>
<th>Max-Wald</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>May-98</td>
<td>Jun-98$^a$</td>
<td>Jun-98</td>
<td>24.19</td>
<td>&lt; 0.05</td>
</tr>
<tr>
<td>Malaysia</td>
<td>Oct-97</td>
<td>Feb-98$^b$</td>
<td>May-98</td>
<td>35.64</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>Philippines</td>
<td>Feb-98</td>
<td>Apr-98$^b$</td>
<td>May-98</td>
<td>47.33</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>South Korea</td>
<td>Dec-97</td>
<td>Mar-98$^a$</td>
<td>May-98</td>
<td>26.52</td>
<td>&lt; 0.05</td>
</tr>
<tr>
<td>Taiwan</td>
<td>Jan-94</td>
<td>Feb-94$^a$</td>
<td>Feb-94</td>
<td>23.93</td>
<td>&lt; 0.05</td>
</tr>
<tr>
<td>Thailand</td>
<td>Jan-98</td>
<td>Mar-98$^b$</td>
<td>Apr-98</td>
<td>27.45</td>
<td>&lt; 0.01</td>
</tr>
</tbody>
</table>

$^a$ Significant at the 5% level (for a critical value of 22.21).
$^b$ Significant at the 1% level (for a critical value of 27.02) or less.
Table 4
Analysis of equity indexes monthly return volatility: structural break in selected parameters

This table reports estimated break dates \( \hat{k} \) for the structural relation between equity and currency return volatility

\[
\sigma_{yt} = \mu + A\sigma_{y,t-1} + b\sigma_{xt} + d_t[k(\lambda + \beta\sigma_{xt})] + \epsilon_t.
\]  

(7)

The Median column in the Table shows the estimated break date \( \hat{k} \) for the time series of monthly equity returns in local currency for each of the countries in the sample. Break dates are estimated using the Wald statistic \( F \) described in Eq. (4). We test the null hypothesis that the post-break coefficient changes are not significantly different from zero, i.e., that no break occurred in the sample period, by comparing the maximum value in the estimated time series \( \hat{F}(k) \) to the 5% quantile of its limiting distribution. The null hypothesis is rejected when the maximum value for \( \hat{F}(k) \), reported in the Max-Wald column of the table, is lower than the critical value for the selected significance level. Bekaert et al. (1998) compute a table with critical values for \( \max \hat{F}(k) \) (for dimensions up to 68), approximating the limiting distribution of the \( F \) process with partial sums of normal random variables for each possible dimension of the test statistic \( \hat{F}(k) \), i.e., for the rank of the selected vector \( S \). From such table we use the asymptotic 1% critical value of 16.37, corresponding to a rank of 2 for the vector \( S \) in Eq. (2). The 2.5th and 97.5th Percentile columns display estimated lower and upper bands, respectively, for the confidence intervals for the “true” break dates, computed according to Eq. (5) with quantiles of the Picard (1985) distribution.

<table>
<thead>
<tr>
<th>Country</th>
<th>2.5th Percentile</th>
<th>Median</th>
<th>97.5th Percentile</th>
<th>Max-Wald</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>May-97</td>
<td>Jun-97\textsuperscript{a}</td>
<td>Jun-97</td>
<td>70.70</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>Malaysia</td>
<td>Jun-94</td>
<td>Oct-94\textsuperscript{a}</td>
<td>Jan-95</td>
<td>23.56</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>Philippines</td>
<td>Jul-98</td>
<td>Aug-98\textsuperscript{a}</td>
<td>Aug-98</td>
<td>63.61</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>South Korea</td>
<td>Sep-97</td>
<td>Nov-97\textsuperscript{a}</td>
<td>Dec-97</td>
<td>17.67</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>Taiwan</td>
<td>Sep-94</td>
<td>Oct-94\textsuperscript{a}</td>
<td>Oct-94</td>
<td>32.55</td>
<td>&lt; 0.01</td>
</tr>
<tr>
<td>Thailand</td>
<td>Sep-94</td>
<td>Oct-94\textsuperscript{a}</td>
<td>Oct-94</td>
<td>21.22</td>
<td>&lt; 0.01</td>
</tr>
</tbody>
</table>

\textsuperscript{a}Significant at the 1% level (for a critical value of 16.37) or less.
Table 5
Break in the structural relation between equity and currency return volatility

This table displays regression coefficients estimated at the break date $\hat{k}$ for the structural relation between equity and currency return volatility

$$\sigma_{yt} = \mu + A \sigma_{yt-1} + b \sigma_{xt} + d_i (k) [\lambda + \beta \sigma_{yt}] + \varepsilon_t ,$$

(7)

The first column of the table reports the corresponding adjusted regression $R^2$, $R^2_a$, for each of the countries in the sample. The three pre-break coefficient columns display the estimated coefficients in the hypothesized structural relationship before the break occurred. We identify the break dates $\hat{k}$ through the Wald Statistic described in Eq. (4). Break dates for each of the countries in our sample are collected in Table 2. The two post-break coefficient columns show the change ($\Delta$) in the structural coefficients after the break occurred. To correct for heteroskedasticity and autocorrelation in the residuals of the above regression, statistical significance of the coefficients is measured using Newey-West standard errors.

<table>
<thead>
<tr>
<th>Country</th>
<th>$R^2_a$</th>
<th>Intercept</th>
<th>$X$</th>
<th>$Y-1$</th>
<th>$\Delta$Intercept</th>
<th>$\Delta$X</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>92.39%</td>
<td>0.076\textsuperscript{b}</td>
<td>$-8.371\textsuperscript{b}$</td>
<td>0.627\textsuperscript{b}</td>
<td>$-0.031\textsuperscript{b}$</td>
<td>8.402\textsuperscript{b}</td>
</tr>
<tr>
<td>Malaysia</td>
<td>96.73%</td>
<td>0.016</td>
<td>0.817</td>
<td>0.735\textsuperscript{b}</td>
<td>$-0.009$</td>
<td>$-0.220\textsuperscript{b}$</td>
</tr>
<tr>
<td>Philippines</td>
<td>93.60%</td>
<td>0.017\textsuperscript{b}</td>
<td>0.204\textsuperscript{b}</td>
<td>0.721\textsuperscript{b}</td>
<td>$-0.003$</td>
<td>0.403</td>
</tr>
<tr>
<td>South Korea</td>
<td>94.34%</td>
<td>0.022\textsuperscript{b}</td>
<td>$-0.006$</td>
<td>0.685\textsuperscript{b}</td>
<td>0.005</td>
<td>0.309</td>
</tr>
<tr>
<td>Taiwan</td>
<td>82.86%</td>
<td>0.055\textsuperscript{b}</td>
<td>$-1.503\textsuperscript{b}$</td>
<td>0.724\textsuperscript{b}</td>
<td>$-0.037\textsuperscript{b}$</td>
<td>1.680\textsuperscript{b}</td>
</tr>
<tr>
<td>Thailand</td>
<td>95.27%</td>
<td>$-0.056\textsuperscript{b}$</td>
<td>$-2.012\textsuperscript{a}$</td>
<td>0.650\textsuperscript{b}</td>
<td>$-0.034\textsuperscript{b}$</td>
<td>2.393\textsuperscript{b}</td>
</tr>
</tbody>
</table>

\textsuperscript{a} Significant at the 10% level.
\textsuperscript{b} Significant at the 5% level or less.
Table 6
Break in the structural relation between equity and currency returns

This table displays regression coefficients estimated at the break date $\hat{k}$ for the structural relation between equity and currency returns

$$y_t = \mu + A y_{t-1} + \sum_{i=1}^{5} b_i x_{t-i-3} + d_t(\hat{k}) \left[ \lambda + \alpha y_{t-1} + \sum_{i=1}^{5} \beta_i x_{t+i-3} \right] + \varepsilon_t \quad (1)$$

The first column of the table reports the corresponding regression adjusted $R^2$, $R^2_{a}$, for each of the countries in the sample. The seven pre-break coefficient columns display the estimated coefficients in the hypothesized structural relationship before the break occurred. We identify the break dates $\hat{k}$ through the Wald Statistic described in Eq. (4). Break dates for each of the countries in our sample are collected in Table 3. Because the breaks in the hypothesized structural relation between equity and currency returns occur late in the sample for 5 of the 6 countries in the sample and because the interaction among the explanatory variables may obscure their individual contribution to the fit even if their joint effect is significant, several of the resulting post-shift estimates for the parameters in Eq. (1) are either statistically or economically insignificant. Hence, those estimates are not reported here, but are available from the authors on request. To correct for heteroskedasticity and autocorrelation in the residuals of the above regression, statistical significance of the coefficients is measured using Newey-West standard errors.

<table>
<thead>
<tr>
<th>Country</th>
<th>$R^2$</th>
<th>$R^2_a$</th>
<th>Pre-break coefficients</th>
<th>Post-break coefficients</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>41.88%</td>
<td>0.002</td>
<td>-0.094b</td>
<td>-0.136</td>
</tr>
<tr>
<td>Malaysia</td>
<td>42.57%</td>
<td>0.013</td>
<td>-0.701</td>
<td>-0.055</td>
</tr>
<tr>
<td>Philippines</td>
<td>62.66%</td>
<td>0.009</td>
<td>-0.166</td>
<td>0.037</td>
</tr>
<tr>
<td>South Korea</td>
<td>44.79%</td>
<td>0.001</td>
<td>0.632b</td>
<td>-0.067</td>
</tr>
<tr>
<td>Taiwan</td>
<td>26.72%</td>
<td>0.207b</td>
<td>7.580b</td>
<td>-0.296b</td>
</tr>
<tr>
<td>Thailand</td>
<td>29.58%</td>
<td>0.003</td>
<td>-0.517</td>
<td>-0.084</td>
</tr>
</tbody>
</table>

a Significant at the 10% level.
b Significant at the 5% level or less.
Table 7
Volatility versus return breaks

This table displays the results of tests for the chronologically sequential relation between equity volatility breaks and equity return breaks in each country. We estimate the following regression:

\[ y_t = \mu + A y_{t-1} + \sum_{i=1}^{5} b_i x_{t-i-3} + d_t(\hat{k}) \left[ \lambda + \alpha y_{t-1} + \sum_{i=1}^{5} \beta_i x_{t+i-3} \right] + \epsilon_t + \gamma d_t(\hat{k}_{VOL}), \]  \hspace{1cm} (8)

where \(d_t(\hat{k}_{VOL})\) is equal to one when \(t\) is higher than or equal to the lower bound of the confidence interval for the corresponding volatility structural break found in Table 3, and zero otherwise. The first column of the table shows the regression adjusted \(R^2\), \(R^2_a\), for the structural relation (8) for each of the countries in the sample. The break dates \(\hat{k}\) and \(\hat{k}_{VOL}\) have been identified with the Wald statistic described in Eq. (4), and are reported in Tables 3 and 4 respectively. \(\gamma\) is the estimated coefficient for the volatility dummy. In the timing column, the table displays S if the volatility and return breaks are simultaneous, B if volatility breaks before returns and A if volatility breaks after returns. Double letters represent a lead-lag of more than one year. To correct for heteroskedasticity and autocorrelation in the residuals of the above regression, statistical significance of the coefficients is measured using Newey-West standard errors.

<table>
<thead>
<tr>
<th>Country</th>
<th>(R^2_a)</th>
<th>Timing</th>
<th>(\gamma)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>40.50%</td>
<td>B</td>
<td>-0.0370</td>
</tr>
<tr>
<td>Malaysia</td>
<td>45.08%</td>
<td>BB</td>
<td>-0.0447(^a)</td>
</tr>
<tr>
<td>Philippines</td>
<td>61.57%</td>
<td>A</td>
<td>-0.0003</td>
</tr>
<tr>
<td>South Korea</td>
<td>44.76%</td>
<td>B</td>
<td>0.1031</td>
</tr>
<tr>
<td>Taiwan</td>
<td>25.66%</td>
<td>A</td>
<td>-0.0163</td>
</tr>
<tr>
<td>Thailand</td>
<td>34.13%</td>
<td>BB</td>
<td>-0.0676(^a)</td>
</tr>
</tbody>
</table>

\(^a\) Significant at the 10% level or less.
Figure 1
Search for a break in the structural relations between equity and currency returns and volatility: Thailand

Figure 1A displays the monthly time series for the Thailand Stock Exchange Index (on the left axis) and a $/baht currency index (on the right axis) constructed with January 1971 as base date (100). Lower values of the currency index represent a weaker baht relative to the U.S. dollar. Figure 1B shows the time series of rolled volatility for both equity and currency returns. Rolled volatility is computed from a monthly return time series using a 12-month rolling window. Figures 1C and 1D report (on the left axis) the monthly time series of the Wald statistic $F(k)$ to test for a break in the structural relation between equity and currency returns (Eq. 1) and equity and currency volatility (Eq. 7) respectively. 5% confidence intervals (plotted at the 5% significance level for the Wald statistic) around the date $\hat{k}$ that maximizes $\hat{F}(k)$ over the sample interval are computed according to Eq. (5). Figures 1C and 1D also show (on the right axis) the monthly time series of flows of funds for the Thailand Stock Exchange, in millions of U.S. dollars.
Figure 2
Return and volatility breaks in East Asia: March to December 1997

Figures 2A and 2B display confidence intervals at the 5% significance level around the estimated break date $\hat{k}$, i.e., the one that maximizes the Wald statistic $\hat{F}(k)$ over the sample interval, for the structural relations between equity and currency return and volatility respectively. The confidence intervals are computed according to Eq. (5). The confidence interval measure of 6 (left axis) is associated with the country for which we measure the most significant $\hat{F}(k)$ in the sample. The confidence interval measure of 1 (left axis) is associated to the country for which we measure the least significant $\hat{F}(k)$ in the sample.
Figure 3
Return and volatility breaks in East Asia: January to October 1998

Figures 3A and 3B display confidence intervals at the 5% significance level around the estimated break date \( \hat{k} \), i.e., the one that maximizes the Wald statistic \( \hat{F}(k) \) over the sample interval, for the structural relations between equity and currency return and volatility respectively. The confidence intervals are computed according to Eq. (5). The confidence interval measure of 6 (left axis) is associated with the country for which we measure the most significant \( \hat{F}(\hat{k}) \) in the sample. The confidence interval measure of 1 (left axis) is associated to the country for which we measure the least significant \( \hat{F}(\hat{k}) \) in the sample.