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Changing Risk, Return, and Leverage: The 1997 Asian Financial Crisis

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Abstract

This paper explores risk and return relations in six Asian equity markets affected by the 1997 Asian financial crisis. After the start of the crisis, national equity betas increased and average returns fell substantially. Beta increases due to leverage linked to exchange rates. The increase in expected return needed to accompany this rise in beta is made possible through the creation of capital losses that lower average returns. We propose a new probability-based asset pricing model that captures leverage effects using valuation ratios. Results show the role of leverage in explaining the likelihood of the financial crises. Cross-sectional evidence supports time-series findings.

I. Introduction

The Asian crisis materialized with the devaluation of the Thai baht July 2, 1997 and spread quickly to other Asian equity markets. The severity of the crisis provides a unique opportunity to understand financial crises and their manifestations. It has caused international investors and academics to reassess the benefits of investing in emerging markets. Based on a sample of six Asian countries (Indonesia, South Korea, Malaysia, the Philippines, Taiwan, and Thailand) from October 1990 to June 1998, currencies collectively lost almost half their value and their U.S. dollar denominated equity market returns dropped sevenfold after the crisis began.[1] This paper investigates the role of leverage in explaining risk and return changes in affected Asian markets before and after the financial crisis.

Leverage is a key feature of current financial crises models. Krugman (1999) states that high leverage financed through extensive foreign currency borrowing starting after 1990 made Asian economies particularly susceptible to financial crisis. Bris and Koskinen (2002) show that depreciation is a solution to a debt

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[1]These countries are commonly used in examining the 1997 Asian crisis. See Kaminsky and Schmukler (1999), Harvey and Roper (1999), and Baig and Goldfajn (1998).

overhang problem. Gande, John, and Senbet (2002) view financial crises as usually involving a corporate debt problem. Harvey and Roper (1999) assert that firms accumulated "excess" leverage in what they call the Asian "bet." Leading up to the crisis, Asian firms bet on future growth by committing increasing amounts of short-term U.S. dollar denominated debt to their already heavily levered capital structures.[2] Once devaluation ensued, leverage, linked to exchange rates, substantially increased and interest payments became harder to meet. This effect was universally experienced in these countries, which lacked significant industrial diversification. Consistent with the deteriorating balance sheets of Asian firms, Baig and Goldfajn (1998) report that sovereign debt yields skyrocketed.

Researchers have also addressed the moral hazard problems associated with financial liberalization and maintaining fixed exchange rates that effectively reduce the relative cost of debt financing. Mishkin (1999) argues that asymmetric information problems started with financial liberalization. Lenders and supervisory agencies are ill equipped to handle rapidly growing credit and lack both the resources and expertise to effectively monitor contracts. Corsetti, Pesenti, and Roubini (1999) discuss moral hazard issues related to maintaining fixed exchange rates. Central Banks implicitly give guarantees to creditors by providing hedging services and "over borrowing" results.

We observe that the average national equity market beta increases fivefold after the start of the crisis. Devaluations are infrequent events and by themselves are not expected to change risk premia, but in the presence of unhedged debt, devaluation directly affects firms' capital structures and therefore discount rates. The rise in discount rates and the necessary creation of capital losses after the start of the crisis resolve an apparent asset pricing irregularity where post-crisis betas rise and average returns fall. Average returns are negatively correlated with changes in beta because they contain capital gains and losses.

The date that defines the start of the crisis must be carefully considered to model the changes in risk and return. The Thai bhat devaluation of July 2, 1997 is the "official" date of the crisis;[3] however, investors' risk assessments do not change overnight. Evidence gleaned from news services indicates investors started revising their expectations earlier than that date and continued to update them in the post-crisis period. Tests show significant discrete shifts in risk and return after the official start of the crisis, and structural change tests confirm significant changes in risk and return without imposing economic priors. Consistent with economic news, statistical break point estimates identify a crisis date five or more months after the official date; at the depth of the crisis.

Our empirical asset pricing models seek to capture the evolution of discount rate changes with portfolios *managed* by valuation ratios.[4] Fama and French (1993) suggest the price-to-book (P/B) ratio could proxy for an omitted *distressed*

[2]Harvey and Roper (1999) report annual book leverage ratios for emerging markets from 1992 to 1996. The average debt-to-equity ratios for Asian countries increased dramatically. For 1992 and 1996, the average total debt-to-equity ratios are: Indonesia (137% and 202%), South Korea (123% and 257%), Malaysia (188% and 336%), the Philippines (55% and 105%), Taiwan (27% and 201%), and Thailand (261% and 417%). In the early stages of financial liberalization, long-term financing arrangements are not available and international lenders are unwilling to lend in the local currency.

[3]See Kaminsky and Schmukler (1999).

[4]We use managed portfolios to characterize time variation in beta. See Cochrane (2001).

firm risk factor.[5] Ferguson and Shockley (2003) illustrate that firm leverage is directly related to valuation ratios. In a discounted cash flow model, Berk (1995) finds lower P/B ratios reflect higher discount rates. We find valuation ratios are half their pre-crisis levels.

The first asset pricing model examined in this paper, commonly seen in the literature, characterizes the time variation in beta as a linear function of valuation ratios and world market conditions.[6] This model is compared to a benchmark asset pricing model with a structural break at the start of the crisis and has success in describing shifts in beta but not average returns. Second, we propose a new probability-based asset pricing model to describe changes in average returns. This model uses, in addition to the world market portfolio, a portfolio that is the world market return managed by an index of investor sentiment. The index of investor sentiment is based on the likelihood of the crisis modeled as a logistic function of leverage indicators and world market conditions. This approach allows the explicit incorporation of investors' assessment of the likelihood of a financial crisis into evolution of beta and discount rates. We estimate the resulting logit and normal regression moment conditions simultaneously and bootstrap critical values of tests' statistics.

Our probability-based asset pricing model can account for the substantial increase in beta and lower average returns after the crisis began. Moreover, this model predicts a rise in risk premia well before the crisis; the other model does not. This suggests that investors started to revise their expectations before the crisis officially started. This prediction is supported by news of currency speculation and the troubles with the Thai economy prior to the crisis. Results show the role of leverage in explaining the likelihood of financial crises. Further, firm-level evidence from the six countries also shows that changes in leverage are associated with cross-sectional differences in average returns.

In Section II, we outline the asset pricing models and testing methodology. Section III shows descriptive statistics of the six sample countries. Section IV presents results from the benchmark model, provides statistical break point tests, reports estimates of break dates, and reconciles break point tests with economic news. Section V describes the leverage indicators, the moment conditions examined, estimation results, and presents cross-sectional evidence. Section VI concludes the paper.

II. Models of Expected Returns

We implement an asset pricing model that uses portfolios managed by valuation ratios to incorporate the impact of changes in leverage and business risk on average returns. Expected returns are mimicked by portfolios that are a combination of the world market portfolio managed by information about valuation ratios. The introduction of managed portfolios effectively allows the beta of each country to time vary according to a chosen information set. Consider an international asset pricing model (IAPM) where time variation in beta and conditional

[5]Fama and French (1998) also find that the book-to-market effect is present in an international setting.

[6]For example, Lewellen (1999) adopts this approach to examine U.S. data.

expected returns are characterized by a set of *managed* portfolios that mimic expected returns,

$$(1) \qquad E[r_t \,|Z_{t-1}] \;=\; \gamma E[r_{Wt} Z_{t-1}],$$

where r_t is a vector of time t excess dollar returns for n countries and is a linear combination of returns on k managed portfolios that are the product of the excess world market return r_{Wt} and k information variables, Z_{t-1} (i.e., valuation ratios) with γ as the vector of coefficients.[7] In this approach, portfolios are managed in the sense that information is used to assign weights to the world market portfolio creating new portfolios that have their own distinct payoffs. The traditional international capital asset pricing model (ICAPM) is obtained if a constant is the only instrument used to weight the world market portfolio. Equation (1) can be expressed in terms of the portfolio weights given to each country,

$$(2) \qquad E[r_t|Z_{t-1}] \;=\; \beta_t^{\mathrm{TV}} \,|\, Z_{t-1} E[r_{Wt}],$$

where $\beta_t^{\mathrm{TV}}|Z_{t-1} = \gamma Z_{t-1}$, is a vector of time-varying *betas* that are linear in the instruments. We posit expected returns in this manner to allow for the instrument set to directly affect beta through the covariance between the world market return and each country's return. We call this model TV-IAPM.

In this paper, we propose an alternative model of expected returns that depends on a portfolio managed by investors' assessment of the likelihood of a financial crisis. Instead of betas depending directly on the instruments set, they depend on the conditional probability of the financial crisis given leverage indicators and world market conditions. The conditional probability reveals a latent index of investor sentiment. We write expected returns for n countries as a combination of the world market and the world market portfolio managed by an index of investor sentiment as

$$(3) \qquad E[r_t|Z_{t-1}] \;=\; \beta_t^{p} \,|\, Z_{t-1} E[r_{Wt}],$$

where $\beta_t^p|Z_{t-1} = \lambda_0 + \lambda_1 p(I_t|Z_{t-1})$, is a vector of time-varying betas depending on n-dimensional vectors of fixed coefficients λ_0 and λ_1. Investors' assessment of the likelihood of the financial crisis given the available information is $p(I_t|Z_{t-1})$. The index of investors' sentiment is based on the logistic relation,

$$(4) \qquad p\,(I_t\,|Z_{t-1}\,) \;=\; \frac{\exp\,(Z_{t-1}\theta)}{1+\exp\,(Z_{t-1}\theta)},$$

where $p(I_t|Z_{t-1})$ is the conditional probability of the financial crisis I_t given known information Z_{t-1} with θ as coefficients. I_t is a binary variable that is zero before the crisis period and one after the crisis begins. We call this formulation of expected returns the P-IAPM.

A benchmark model with a structural break at the start of the crisis period is used to assess the magnitude of the shifts in beta and average returns. The expected returns pre- and post-crisis are

$$(5) \qquad E\,(r_t\,|I_t = 0\,) \;=\; \beta E\,(r_{wt}) \quad \text{and} \quad E\,(r_t\,|I_t = 1\,) \;=\; (\beta + \phi)\,E\,(r_{wt})\,,$$

[7]This model is a variant of the standard conditional international asset pricing models utilized by Dumas and Solnik (1995), Bekaert and Harvey (1995), Ferson and Harvey (1998), and Ghysels (1998), among others.

where the vector of coefficients ϕ depicts the shift in betas after the crisis begins.

After establishing the significance of the benchmark model, we test whether our managed portfolio formulations can account for the shifts in beta. We cast each model in terms of the benchmark,

$$(6) \qquad \begin{aligned} E(r_t \,|I_t = 0) &= \beta_t^q \,|\, Z_{t-1} E(r_{wt}) \text{ and} \\ E(r_t \,|I_t = 1) &= (\beta_t^q \,|\, Z_{t-1} + \phi^q) E(r_{wt}) \end{aligned}$$

where $q =$ (TV,P) for TV-IAPM and P-IAPM specifications. If ϕ^q is statistically zero, then the particular formulation of expected returns q can offer an explanation for the shifts in risk and return when the crisis started.

III. Data

The six Asian countries that abandon their exchange rate management policy at the height of the financial crisis in the summer of 1997 are Indonesia (ID), South Korea (KR), Malaysia (MY), the Philippines (PH), Taiwan (TW), and Thailand (TH). We obtain weekly data on national equity market total returns, exchange rates, P/B and price-to-earnings (P/E) ratios of each country, together with 48 emerging market country averages of these variables from Standard & Poor's Emerging Markets Database (EMDB). Our sample spans 401 weekly (end of Friday) observations from October 5, 1990 to June 5, 1998.[8] The world return (WD) is represented by the Morgan Stanley Capital Market International (MSCI) world index. U.S. dollar returns for the national markets are computed by translating the local index return into U.S. dollars using the end of week exchange rate. Excess returns are obtained by subtracting the three-month U.S. Treasury bill yield.

Table 1 presents descriptive statistics of the sample countries. The crisis culminates with the abandonment of exchange rate supports and a drastic depreciation of Asian currencies. Panel A reports devaluation dates and total depreciation of exchange rates before and after the start of the crisis. After Thailand, a wave of devaluations follows and currencies collectively lose almost half their value. A rank correlation between the timing of devaluation and its extent indicates that there is weak evidence (i.e., $\rho = 0.22$) that countries which wait to devalue suffer the worst devaluation.

The returns to international investors not only suffer from currency devaluation, but also from disruptions in local markets. Panels B and C of Table 1 provide a decomposition of annualized dollar returns before and after devaluations. The equity returns of all six Asian countries drop after the crisis and standard deviations increase substantially while these effects are more pronounced in dollar terms. The six-country average U.S. dollar return decreases by more than sevenfold and standard deviations increase more than threefold when comparing before and after devaluation samples. The effect of the crisis is slowly realized. Even in the last subperiod after devaluations (Sub2) returns are still negative.

[8]Harvey (1995) claims that EMDB backfilling problems and the associated survivorship bias is confined to the pre-1981 sample period. The dividend price ratio is excluded from this study due to missing observations for Indonesia.

Before devaluation, the local market return almost exclusively determines the U.S. dollar return because of currency management policies. After devaluation, exchange rates are responsible for more than half of the drop in dollar returns (six-country average in Table 1, panel B). The increase in the volatility of dollar returns is due to increases in both local market volatility and exchange rates. The standard deviation of dollar returns before devaluations is determined by local returns; after devaluations, half of the volatility of dollar returns is related to exchange rates (six-country average in panel C). Taiwan is an exception with its standard deviations changing very little. The local return and the exchange rate movements reinforce one another, further amplifying their effect on dollar return volatility. Referring to panel C, the correlation between local returns and exchange rates $\rho_{R,S}$ change from virtually zero in the pre-devaluation period to more than 40% in all countries and as much as 58% in South Korea in the post-devaluation period.

TABLE 1

Descriptive Statistics

Panel A. Total Currency Devaluation (−) and Appreciation (+) in % Ordered by Event Date

		Country							
		TH	PH	MY	ID	TW	KR		
		Devaluation Date							
Period	Weeks	07/02/97	07/11/97	07/14/97	08/14/97	10/18/97	11/07/97	Avg.	Rank ρ
10/05/90 to 06/27/97	352	−2	−6	7	−25	−4	−19	−8	0.49
07/04/97 to 06/05/98	49	−33	−33	−37	−79	−19	−36	−39	0.22
10/05/90 to 06/05/98	401	−42	−37	−32	−84	−22	−49	−44	0.09

Panel B. U.S. Dollar Returns, Local Returns, and Exchange Rate Changes

		Before Devaluation			After Devaluation			
		Sub1	Sub2	Total	Sub1	Sub2	Total	Whole
ID	Rus	0.032	0.065	0.049	−2.743	−0.345	−1.516	−0.119
	R	0.075	0.120	0.097	−0.972	0.149	−0.399	0.044
	S	−0.043	−0.056	−0.050	−1.894	−0.737	−1.302	−0.184
KR	*Rus*	0.103	−0.206	−0.051	−0.302	−0.544	−0.427	−0.080
	R	0.138	−0.160	−0.011	0.809	−1.111	−0.182	−0.024
	S	−0.035	−0.048	−0.042	−1.483	0.523	−0.447	−0.073
MY	*Rus*	0.299	0.024	0.161	−2.134	0.037	−1.025	0.022
	R	0.308	−0.007	0.150	−1.273	−0.112	−0.680	0.053
	S	−0.008	0.031	0.012	−0.918	−0.009	−0.454	−0.043
PH	*Rus*	0.462	0.004	0.232	−1.581	0.323	−0.629	0.129
	R	0.486	−0.006	0.240	−0.799	0.214	−0.292	0.176
	S	−0.022	0.010	−0.006	−0.816	0.039	−0.388	−0.052
TW	*Rus*	0.274	0.092	0.183	0.011	−0.659	−0.334	0.140
	R	0.266	0.119	0.192	0.367	−0.536	−0.098	0.169
	S	0.006	−0.028	−0.011	−0.370	−0.129	−0.246	−0.031
TH	*Rus*	0.343	−0.223	0.060	−1.929	0.078	−0.905	−0.058
	R	0.345	−0.220	0.063	−0.800	−0.141	−0.464	−0.001
	S	−0.002	−0.003	−0.003	−1.147	0.141	−0.490	−0.062
Six-country avg.	*Rus*	0.252	−0.041	0.106	−1.446	−0.185	−0.806	0.006
	R	0.270	−0.026	0.122	−0.445	−0.256	−0.353	0.070
	S	−0.017	−0.016	−0.017	−1.105	−0.029	−0.555	−0.074

(continued on next page)

TABLE 1 (continued)

Descriptive Statistics

Panel C. Return Variance Decomposition

		Before Devaluation			After Devaluation			
		Sub1	Sub2	Total	Sub1	Sub2	Total	Whole
ID	$\sigma(Rus)$	1.708	1.651	1.678	6.511	12.075	9.724	3.560
	$\sigma(R)$	1.689	1.576	1.631	4.795	5.747	5.271	2.306
	$\sigma(S)$	0.195	0.265	0.233	3.691	9.025	6.896	2.279
	$\rho_{R,S}$	0.050	0.229	0.147	0.271	0.258	0.256	0.207
KR	$\sigma(Rus)$	1.976	1.770	1.880	10.110	4.414	7.580	2.753
	$\sigma(R)$	1.971	1.691	1.840	5.656	3.210	4.586	2.169
	$\sigma(S)$	0.131	0.309	0.237	5.430	1.853	4.064	1.141
	$\rho_{R,S}$	0.005	0.207	0.127	0.716	0.514	0.579	0.350
MY	$\sigma(Rus)$	1.327	1.414	1.376	4.267	6.881	5.795	2.383
	$\sigma(R)$	1.327	1.398	1.370	3.422	4.498	4.008	1.891
	$\sigma(S)$	0.342	0.237	0.294	1.841	2.923	2.471	0.895
	$\rho_{R,S}$	−0.123	−0.023	−0.086	0.293	0.650	0.548	0.374
PH	$\sigma(Rus)$	1.984	1.565	1.798	3.423	4.514	4.078	2.209
	$\sigma(R)$	1.903	1.525	1.739	2.412	3.125	2.809	1.902
	$\sigma(S)$	0.732	0.350	0.573	1.997	1.915	1.983	0.875
	$\rho_{R,S}$	−0.081	−0.001	−0.061	0.236	0.630	0.471	0.166
TW	$\sigma(Rus)$	2.682	1.705	2.246	2.694	1.822	2.276	2.250
	$\sigma(R)$	2.617	1.638	2.182	2.077	1.598	1.873	2.157
	$\sigma(S)$	0.316	0.245	0.283	1.078	0.430	0.808	0.360
	$\rho_{R,S}$	0.136	0.219	0.164	0.402	0.435	0.341	0.179
TH	$\sigma(Rus)$	2.010	2.001	2.023	5.107	5.869	5.545	2.717
	$\sigma(R)$	2.007	1.978	2.010	4.846	4.375	4.576	2.467
	$\sigma(S)$	0.174	0.413	0.316	2.053	2.546	2.384	0.892
	$\rho_{R,S}$	−0.020	−0.047	−0.035	0.000	0.382	0.215	0.131
Six-country avg.	$\sigma(Rus)$	1.948	1.684	1.834	5.352	5.929	5.833	2.645
	$\sigma(R)$	1.919	1.634	1.795	3.868	3.759	3.854	2.149
	$\sigma(S)$	0.315	0.303	0.323	2.682	3.115	3.101	1.074
	$\rho_{R,S}$	−0.006	0.097	0.043	0.320	0.478	0.402	0.235
Panel D. Country Average and World Excess Returns								
Country avg.	Mean	0.211	−0.091	0.060	−1.497	−0.235	−0.857	−0.041
	Std. dev.	1.949	1.685	1.834	5.352	5.929	5.833	2.646
WD	Mean	0.088	0.067	0.078	−0.132	0.255	0.066	0.076
	Std. dev.	0.840	0.643	0.747	1.067	0.896	0.992	0.779
Between country & WD	$\rho(R_i, R_j)$			0.202			0.504	0.351
	$\rho(R_i, R_W)$			0.199			0.492	0.269

Panel A reports devaluation dates of Indonesia (ID), South Korea (KR), Malaysia (MY), Philippines (PH), Taiwan (TW), Thailand (TH), total currency changes for pre-crisis, post-crisis, and the whole sample period that spans October 5, 1990 to June 5, 1998 (401 weeks). Panel B reports annualized means and standard deviations of U.S. dollar returns (*Rus*), local market returns (*R*), and exchange rates (*S*). Local EMDB index returns (end of week) are converted to U.S. dollars using the end of week exchange rate. The before and after devaluation periods are based on the closest Friday after each event date reported in panel A and subperiods (Sub1 and Sub2) split in half before and after devaluation periods. The variance decomposition in panel C uses the approximation that ignores the cross-term between the local return and the exchange rate and reports annualized standard deviations of U.S. dollar returns $\sigma(Rus)$, local market returns $\sigma(R)$, exchange rates $\sigma(S)$, and the correlation of exchange rates with local market returns $\rho_{R,S}$. Panel D reports annualized means and standard deviations of the MSCI world index in excess of the three-month T-bill rate, the average correlation between countries $\rho(R_i, R_j)$, and the average of correlations with the world index $\rho(R_i, R_W)$. The subperiod for the world index is based on July 2,1997.

This positive correlation between exchange rates and local returns is consistent with leverage linked to exchange rates. Devaluation in export-oriented countries normally enhances investment opportunities, leading to an increase in firm value resulting in a negative correlation of exchange rates and local return. As Krugman (1999) points out, Asian firms that were characterized by heavy leverage financed through extensive foreign currency borrowing could not take advantage of these growth opportunities because their balance sheets eroded after devalua-

tion destroying entrepreneur wealth used as collateral for lending. Depreciation works against firm value if firms are heavily levered in foreign currency, resulting in the positive correlation seen in these data. Accounting for risk can generate the same effect that Krugman predicts. Leverage increases with exchange rate depreciation cause equity betas to rise. Investors suffer capital losses because the equities they hold become more risky. Local returns have positive correlation with exchange rate changes because they are associated with capital gains and losses in the local market.

Panel D of Table 1 reports annualized means and standard deviations of U.S. dollar returns in excess of the three-month T-bill for the average of the six countries and the MSCI world index. The average world return and its standard deviation change very little over the sample period. As found in the literature, these countries display very high volatility relative to the world market return.[9] Panel D also depicts the average correlation between countries and the correlations between local returns and the world returns pre- and post-devaluations. The correlations increase after the start of the crisis because countries devalue around the same time.[10] More striking is that average correlation with the world return more than doubles after devaluations. In sum, investing in the markets became more risky while yielding lower average returns and fewer diversification opportunities. The next section documents the magnitude and significance of shifts in risk and return around the start of the crisis.

IV. Tests for Changing Exposure to Systematic Risk

A. Shifts in Beta at Devaluation and Crisis Dates

The Thai bhat devaluation on July 2, 1997 is the earliest economic date used in the literature as the start of the Asian crisis. An ICAPM with a discrete structural break at the start of the crisis is a potential candidate benchmark model for evaluating our specifications of beta. From equations (5), the candidate multi-country benchmark model is

$$r_{jt} = \alpha_j + \phi_{0j} I_{\mathrm{TH}t} + [\beta_j + \phi_{1j} I_{\mathrm{TH}t}]\ r_{Wt} + \varepsilon_{jt}, \tag{7a}$$

with moment conditions,

$$E[(1, I_{\mathrm{TH}}, r_{Wt}, I_{\mathrm{TH}t} r_{Wt})\ \varepsilon_{jt}] = 0, \quad j = (1,6), \tag{7b}$$

where r_{jt} and r_{Wt} are the excess returns of each country and the world indices, I_{TH} is the indicator variable for the start of the crisis in Thailand, and ε_t is the residual vector for the six countries. The coefficients ϕ_{0j} and ϕ_{1j} assess changes in intercept and slope.

Table 2 contains our regression findings. For comparison purposes, we present structural break results for each country at its devaluation date and at the date the crisis started. Referring to the coefficient averages in panel A, beta jumped

[9]See De Santis and Imrohoroglu (1997), Bekaert and Harvey (1997), and Aggarwal, Inclan, and Leal (1999).

[10]Bekaert, Harvey, and Ng (2004) find evidence of contagion effects during the Asian crisis

fivefold from before to after the crisis began. The intercept terms are negative and significant after devaluation. Consistent with any equilibrium asset pricing model, as beta increases, investors suffer capital losses lowering average returns. The results are virtually the same whether the devaluation date or the crisis date is used as the indicator. Five of six countries have significant breaks at the 10% level except Taiwan.

TABLE 2

Benchmark ICAPM

Panel A. Structural Shifts at Devaluation Date and Crisis Date

	At Devaluation Date					At Crisis Date (benchmark)				
Country	α_D	β_D	ϕ_{D0}	ϕ_{D1}	R^2	α	β	ϕ_0	ϕ_1	R^2
ID	0.000	0.208	−0.034	4.240	0.199	0.000	0.197	−0.035	4.174	0.200
	(0.937)	(0.361)	(0.001)	(0.000)		(0.928)	(0.075)	(0.052)	(0.005)	
KR	−0.003	0.625	−0.018	2.210	0.108	−0.001	0.547	−0.017	1.783	0.112
	(0.312)	(0.001)	(0.062)	(0.000)		(0.460)	(0.000)	(0.212)	(0.068)	
MY	0.001	0.607	−0.024	2.183	0.216	0.001	0.621	−0.026	2.130	0.216
	(0.554)	(0.000)	(0.000)	(0.000)		(0.296)	(0.000)	(0.029)	(0.076)	
PH	0.003	0.407	−0.018	1.940	0.163	0.003	0.480	−0.019	1.799	0.166
	(0.157)	(0.005)	(0.003)	(0.000)		(0.086)	(0.002)	(0.023)	(0.006)	
TW	0.002	0.638	−0.012	0.624	0.074	0.003	0.676	−0.014	0.385	0.078
	(0.428)	(0.000)	(0.124)	(0.116)		(0.275)	(0.001)	(0.013)	(0.249)	
TH	0.000	0.450	−0.022	2.675	0.183	0.000	0.451	−0.022	2.675	0.183
	(0.875)	(0.011)	(0.003)	(0.000)		(0.848)	(0.016)	(0.065)	(0.001)	
Coeff. avg.	0.001	0.489	−0.021	2.312		0.001	0.495	−0.022	2.158	

Panel B. Joint Tests

	Description	Tests	df	Wald	J-Stat.	LR
H_1	Before crisis intercepts zero	$\alpha = 0$	6	0.464	0.485	0.655
H_2	No change in intercept after crisis	$\phi_0 = 0$	6	0.080	0.239	0.000
H_3	No discrete shift in beta after crisis	$\phi_1 = 0$	6	0.003	0.211	0.000
H_4	No structural change in the ICAPM relationship at crisis date	$\phi_0 = \phi_1 = 0$	12	0.000	0.282	0.000
H_5	Intercepts zero before and after crisis	$\alpha = \phi_0 = 0$	12	0.212	0.415	0.002
H_6	No structural change with intercepts zero	$\alpha = \phi_0 = \phi_1 = 0$	18	0.000	0.360	0.000

Table 2 presents benchmark ICAPM regression coefficients and joint tests for six Asian countries. The set of coefficients in panel A presents structural break results in the ICAPM relation first using the devaluation date of each country and then the benchmark crisis date, July 2, 1997, for all countries in the rightmost columns. The coefficients α and β is the intercept and beta of the ICAPM and ϕ_0 and ϕ_1 represent the shifts in intercept and slope. All returns are in U.S. dollars in excess of the U.S. three-month T-bill rate. Local EMDB index returns (end of week) are converted to U.S. dollars using the EMDB end of week exchange rate. The market return is the MSCI world index return. These data span October 5, 1990 to June 5, 1998 ($T = 401$). p-values are in parentheses. Panel B presents the p-values from Wald, J-stat, and LR tests joint hypotheses described in the panel that include all countries and df denotes the degrees of freedom. For panel B, the benchmark system (equations (7)) is estimated with GMM and a first-order Newey and West correction applied.

Our investigation into whether expected return formulations capture the structural breaks and whether the resulting model conforms to asset pricing restrictions leads us to examine additional related hypotheses defined in panel B of Table 2. Joint tests of the benchmark model use Wald, J-statistics, and Likelihood Ratio (LR) test statistics on a system estimated by Hansen's (1982) GMM with a Newey and West (1987) correction for heteroskedasticity and first-order autocorrelation.[11] Panel B shows evidence of a *discrete* shift in beta across countries. The Wald and LR tests show similar results. They strongly *reject* the joint hypothesis involving *no* discrete shifts in beta across countries. They also confirm that the intercept changes after the crisis. The J-statistic tests fail to reject all of the hy-

[11] We selected lag lengths of one to 10 to check robustness. Our conclusions are the same.

potheses. Ghysels (1998) and Garcia and Ghysels (1998) discuss the notoriously low power of J-statistics in detecting structural change.

B. Evaluating Statistical Breaks and Break Date Estimates

Evidence for large significant shifts in risk and return in Asian countries using either their devaluation dates or the official crisis date is convincing. However, investors may have anticipated changes in risk long before the first devaluation. We first look at changes in the ICAPM relation without imposing a particular break date for each country to evaluate whether the Thai devaluation is a contributing rather than an initiating factor in reassessing asset prices. We identify statistical break dates, form confidence intervals for the break dates, and assess if the earliest date given by the confidence interval is before the official start of the Asian crisis.

We use a battery of tests on the traditional ICAPM (equation (7a) without the break) for each country separately to identify structural breaks without imposing economic priors. We employ the standard CUSUM, the CUSUMSQ, the Chow test, and the Andrews (1993) LM-test. The Andrews test has the most attractive power properties of the four.[12]

Under the Andrews (1993) sup-LM test for a structural break with an unknown break point, a particular interval is inspected for a possible break. The $\sup_{\pi \in \Pi} \mathrm{LM}_T(\pi)$ test statistic is the maximum value of successive tests that occurs at π (proportion of sample) in the sampling interval defined by Π. Supremum (sup) tests are based on a distribution of the maximum value from successive test statistics where critical values depend on the window size and the numbers of break coefficients. The Andrews sup-test is analogous to recording the maximum value from successive F-tests, which places a dummy variable depicting a pre- and post-test date for each date inside a designated test window. We employ the most general of Andrews' tests, the sup-LM test, which allows for variances to change in successive pre- and post-dates and Newey West corrections of covariance matrices (Eq. 4.3 in Andrews (1993)). This methodology permits the examination of breaks in intercept and/or slope as well as inspecting a system for full or partial breaks.

Collectively, these tests confirm the presence of a structural break in the ICAPM relation. Panel A in Table 3 presents the results of our structural break tests. The CUSUMSQ and Chow tests find strong evidence for structural change with the exception of Taiwan confirming regression results.[13] The CUSUM tests, which have the weakest power of the three, do not reject.[14] Sup-LM tests confirm the presence of a full or partial structural change for each country without imposing a change date. We computed successive LM statistics over the entire sample and observed that global maximum values occur between four and one-half to seven months after the Thai devaluation date. As long as the testing interval includes these globally maximum dates, the results are the same. We report test

[12]Ghysels (1998) and Garcia and Ghysels (1998) use the Andrews LM-test to detect structural change in returns data.

[13]We examined CUSUMSQ plots to gain an understanding of when the ICAPM relation becomes unstable, but they do not reveal a clear pattern. The Chow tests split the sample in half.

[14]Garbade (1977) reports the weak power properties of CUSUM tests.

results using the last half of our sample for the testing interval, which encompasses all the devaluation dates.

TABLE 3

Structural Change Tests and Break Date Estimates

Panel A. Structural Change Tests

				Andrews $\sup_{\pi \in \Pi} LM_T(\pi)$		
Country	CUSUM	CUSUMSQ	CHOW	α, β Change	β Change Only	α Change Only
ID	0.651	0.000	0.000	43.935***	25.445***	4.011
KR	0.658	0.000	0.039	10.701*	12.873***	7.786*
MY	0.321	0.000	0.000	27.642***	30.354***	12.807***
PH	0.112	0.000	0.000	23.950***	18.970***	10.045**
TW	0.104	0.000	0.126	29.156***	5.758	15.229***
TH	0.266	0.000	0.000	68.048***	78.088***	14.621***
System				281.928***	212.826***	99.583***

Panel B. Break Point Confidence Intervals

		Stat. Break Point		95% Confidence Interval	
Country	Devaluation	$\hat{k}$	Date ($\hat{k}$)	Symmetric	Skewed
ID	360	380	01/02/98	[375, 385]	[376, 392]
KR	372	385	02/16/98	[358, 412]	[357, 415]
MY	356	377	12/12/97	[368, 386]	[373, 404]
PH	355	380	01/02/98	[368, 392]	[370, 396]
TW	370	380	01/02/98	[342, 418]	[337, 386]
TH	354	374	11/21/97	[360, 388]	[364, 406]

Panel C. News of Bond Downgrades

Date	09/96	10/96–03/97	04/97	05/97–07/97	08/97	09/97	10/97	11/97	12/97	01/98	02/98
ID	–	–	–	–	–	–	1	–	5	4	–
KR	–	–	–	–	1	–	1	1	4	–	–
MY	–	–	–	–	–	1	–	–	2	–	–
PH	–	–	–	–	–	–	–	–	–	–	3
TW	–	–	–	–	–	–	–	–	–	–	–
TH	1	–	1	–	1	1	–	2	4	3	–

Panel A shows p-values from CUSUM, CUSUMSQ, and CHOW tests and reports sup LM(π) statistics using Andrews' (1993) equation 4.3 using critical values from Table 1 (p. 840). The observation window is 201 ($\pi_1 = 0.5$) to 389 ($\pi_2 = 0.97$), $\lambda = 32.33$. The number of break points $p = 2$ for full structural change $\{\alpha, \beta\}$ and $p = 1$ under partial change, α only or β only. For the system tests, $p = 12$ under full structural change and $p = 6$ for α only or β only. Panel B provides confidence intervals on break dates based on Bai (1997) equations (17) and (18) and distributions presented in his Appendix B. The break date estimate $\hat{k}$ is the observation that corresponds to sup LM(π) for each country. Skewed confidence intervals allow for different covariance matrices pre- and post-break estimate. Panel C presents the number of unique news reports of bond downgrades from September 1996–February 1998.

***,**,* denote 1%, 5%, and 10% significance levels, respectively.

Bai (1997) shows the dates that correspond to the Andrews test statistic are the least squares estimates of the break dates. The estimate of the break date is $\hat{k} = \arg\ \min_{1 \leq k \leq T} S_T(k)$, where $S_T(k)$ is the sum of squared residuals, $k = [\pi T]$, $\pi \in (0, 1)$, and $[\cdot]$ denotes the greatest integer function. The break date estimate maximizes the difference between sets of coefficients from the two subsamples defined by the break date thereby minimizing the sum of squares. This objective is equivalent to choosing the date that maximizes the Andrews sup-LM statistic.

Panel B in Table 3 presents break date estimates. The break date estimate $\hat{k}$ is the observation that corresponds to $\sup_{\pi \in \Pi} LM_T(\pi)$ for each country. The skewed confidence interval allows for different covariance matrices for pre- and post-$\hat{k}$. The statistical break dates fall well after the devaluation event dates. We cannot reject that the statistical break points and devaluation dates are the same for Korea and Taiwan, the last to devalue. In other countries, devaluation dates

fall before the start of each country's confidence interval. Only the interval of Taiwan encompasses the start of the crisis. Taiwan is the country with the weakest evidence of a structural shift and therefore has the widest of all the confidence intervals. This evidence indicates that the crisis came as a surprise to investors and that the Thai devaluation is the start of significant risk and return changes in Asian countries.

C. Reconciling Economic and Statistical Event Dates

Statistical tests may lead one to conclude the first devaluation is an initiating factor to the reassessment of risk in the region. The problem is that statistical event dates do not pin down the start of the crisis; rather they indicate when the crisis is at its worst. Break point estimates look for the maximum difference in the coefficient vector from two subsamples. As such, the maximum difference is at the height of the crisis not the start. We reconcile economic and statistical break dates and gain a sense of how investor expectations change over this tremulous period by presenting a summary of news reports of currency speculation in the Thai baht, and announcements by the World Bank and IMF. We also examine the timing of Standard & Poor's and Moody's downgrading of debt instruments of particular countries.[15]

Wire services reported speculation in the Thai bhat as early as February 1997, suggesting investors started to reassess risk and return before the official start of the crisis July 2, 1997. A general slowing of economic growth since 1996 and mounting bad loans to property developers made the baht continue to weaken. When devaluation was announced, it came as a surprise to some analysts as it was thought that the Bank of Thailand had been successful in keeping speculators at bay. In spite of the wave of devaluations that followed Thailand's, it took time for investors to realize the extent and duration of the crisis. Consistent with the break point estimates we find, IMF analysts did not revise their economic outlook until December of that year.

Debt rating agencies downgraded Thai debt instruments one month after the start of the crisis. Panel C in Table 3 presents the number of reported debt instrument downgrades by Moody's and Standard & Poor's. Before the Thai devaluation, there were two reports, both for Thailand. A flood of downgrades followed in December when the debts of Indonesia, South Korea, and Thailand sunk to junk bond status. Many of these downgrades came two days after the issuance of the December 20, 1997 revised IMF outlook. We did not find any reports for Taiwan, which is consistent with our earlier findings.

Taiwan was arguably the country least affected by the Asian crisis. The Central bank announced a change in exchange rate policy, but certainly had sufficient foreign currency reserves to thwart speculators. In 1997, Taiwan had the third largest stock of foreign currency reserves in the world after Japan and China.[16] Taiwan also had at least two and one-half times the foreign currency reserves

[15]Based on a LexisNexis search of wire reports. Search terms available from authors.

[16]Agence France Presse News Wire, October 1, 1997.

of any of the other countries affected by the Asian crisis.[17] The Central Bank of China continued to support the Taiwan dollar even after it announced that it would stop defending the currency.[18]

V. Modeling Exposure to Systematic Risk with Leverage Managed Portfolios

A. Leverage Indicators and Other Financial Instruments

Leverage is difficult to measure and the data is likely to consist of annual book values. However, observable valuation ratios containing price, which are part of market leverage, respond to changes in leverage as well as business risk because price changes convey information about discount factor change. This link can be understood through a standard discounted cash flow model.[19] Today's earnings or book values predict next period's cash flows and price is the discounted value.[20] The increase in equity beta caused by leverage or business risk raises expected return; price reflects the change in discount rate and lowers the valuation ratio, as capital losses must be suffered to make higher expected returns possible.[21]

We use indicators of leverage and other controls for world market conditions to embed information into the evolution of beta. The indicators of leverage are the P/B and P/E ratios and exchange rates (FX), lagged one week. Exchange rates are included as a leverage proxy because interest payments are denominated in U.S. dollars.

Table 4 presents summary statistics on the instruments. P/B and P/E ratios are closely related, being scaled versions of price. Panel A shows both P/B and P/E ratios fell to nearly half their pre-crisis averages in the post-crisis sample indicating higher discount rates and leverage after the start of the crisis. Taiwan is again the exception with its ratios remaining constant. On average, the annualized return on holding foreign exchange is −51.7% for the post-crisis period. It was virtually zero before the crisis. U.S. dollar excess returns are negatively correlated with their indicators on average and this correlation strengthens after the crisis. A negative relation would be predicted with lower P/B and P/E ratios: the lower the P/B, P/E, or returns to foreign exchange in a particular week, the higher the expected returns in the following week.

Panel B reports summary statistics for broader emerging market conditions in 48 emerging markets as well as an average correlation between local returns and these instruments. The lagged world return, R-WD, is also included as part

[17] International Financial Statistics and database from Central Bank of China foreign exchange reserves.

[18] Agence France Presse, October 19, 1997.

[19] The argument is detailed in a simple model available from the authors.

[20] Pontiff and Schall (1998) provide evidence that P/B ratios and book values on DJIA and S&P industrials predict future market returns and earnings, respectively.

[21] A fall in growth rates is unlikely to explain a decline in valuation ratios. If expected returns increase, then the capital losses incurred make possible the higher expected returns. Some combination of higher growth rates and future capital gains must occur in order to finance higher expected returns. Also, in export-oriented economies devaluation is thought to enhance growth opportunities.

TABLE 4

Local Leverage Indicators and Global Market Conditions

Panel A. Local Leverage (z^l_{jt-1})

Country		P/B Before	P/B After	P/E Before	P/E After	FX Before	FX After
ID	Mean	2.483	1.861	19.983	14.246	−0.041	−1.214
	Std. dev.	0.515	0.455	4.512	3.867	0.200	6.461
	Corr(r, z_{t-1})	−0.098	−0.037	−0.101	−0.023	−0.050	−0.277
KR	Mean	1.173	0.692	22.280	22.957	−0.032	−0.370
	Std. dev.	0.244	0.131	6.873	18.997	0.215	3.227
	Corr(r, z_{t-1})	−0.036	−0.066	0.020	−0.005	−0.003	−0.356
MY	Mean	3.334	1.817	26.186	13.263	0.011	−0.427
	Std. dev.	0.772	0.540	5.417	4.108	0.295	2.423
	Corr(r, z_{t-1})	−0.118	−0.168	−0.115	−0.169	0.000	−0.092
PH	Mean	3.297	1.657	19.997	12.759	−0.006	−0.381
	Std. dev.	0.741	0.329	7.381	1.909	0.574	1.963
	Corr(r, z_{t-1})	−0.113	−0.261	−0.079	−0.182	−0.048	−0.215
TW	Mean	3.284	3.323	25.722	31.445	−0.005	−0.217
	Std. dev.	0.622	0.420	5.603	4.146	0.270	0.715
	Corr(r, z_{t-1})	−0.097	−0.032	−0.074	−0.059	−0.060	0.027
TH	Mean	2.862	1.091	17.323	−6.651	−0.003	−0.490
	Std. dev.	0.837	0.197	5.170	23.372	0.316	2.384
	Corr(r, z_{t-1})	0.009	−0.064	−0.044	−0.351	−0.040	−0.102
Six-country avg.	Mean	2.739	1.740	21.915	14.670	−0.013	−0.517
	Std. dev.	0.622	0.345	5.826	9.400	0.312	2.862
	Corr(r, z_{t-1})	−0.076	−0.105	−0.066	−0.132	−0.034	−0.169

Panel B. Global Market Conditions (z^g_{t-1})

	P/B-EM Before	P/B-EM After	P/E-EM Before	P/E-EM After	R-WD Before	R-WD After	R-EM Before	R-EM After
Mean	1.787	1.567	18.615	17.123	0.123	0.116	0.144	−0.380
Std. dev.	0.219	0.134	3.597	1.927	0.747	0.992	0.908	1.648
Six-country avg. Corr(r, z^g_{t-1})	−0.058	−0.118	−0.020	−0.143	0.095	−0.143	0.111	−0.185

Table 4 presents the annualized means, standard deviations, and own country correlation of each indicator with the U.S. dollar excess return. Sample for each country consists of 401 observations and covers the period from 10/05/90 to 06/05/98. The local leverage indicators (z^l_{jt-1}) in panel A are the local P/B and P/E ratios and foreign exchange returns (FX). The global market conditions (z^g_{t-1}) in panel B are the emerging market price-to-book ratio (P/B-EM), the emerging market price-to-earnings ratio (P/E-EM), the MSCI world index return (R-WD), and the emerging country index returns (R-EM). The before- and after-crisis periods are based on the closest Friday after the Thai event date July 2, 1997.

of the broader market conditions. The valuation ratios for the broader markets change very little over the sample period. As with local P/B and P/E ratios, the correlation of local market dollar returns and broader emerging market ratios are negative and strengthen after the start of the crisis. It is interesting to note that the P/B and P/E ratios of the Asian countries in our sample are much higher before the crisis and come back into line with broader emerging market values after the crisis. Taiwan has ratios similar to the broader markets.

B. Portfolios Managed by Leverage: TV-IAPM

The first attempt to describe the changes in risk and return after the start of the crisis uses portfolios managed by leverage and world market conditions

to yield time-varying betas that are linear in a set of instruments. The empirical setup from equation (6) for the TV-IAPM is

$$r_{jt} = \alpha_j + \phi_{0j} I_{\mathrm{TH}t} + \left[\beta_{jt}^{v} \middle| Z_{t-1}^{l,g} + \phi_{1j} I_{\mathrm{TH}t}\right] r_{Wt} + \varepsilon_{jt}, \tag{8a}$$

$$\text{where} \quad \beta_{jt}^{v} \middle| Z_{t-1}^{l,g} = \gamma_{0j} + \gamma_{lj} Z_{jt-1}^{l} + \gamma_{gj} Z_{t-1}^{g}. \tag{8b}$$

The moment conditions are

$$E\left[\left(1, I_{\mathrm{TH}t}, r_{Wt}, I_{\mathrm{TH}t} r_{Wt}, Z_{jt-1}^{l} r_{Wt}, Z_{jt-1}^{g} r_{Wt}\right) \varepsilon_{jt}\right] = 0, \quad j = (1,6), \tag{8c}$$

where the jth country's excess return r_{jt} is written in terms of the benchmark model and regressed on a constant α_j, an indicator variable for Thailand I_{TH} with ϕ_{0j} as its shift parameter, and the excess returns on the world portfolio r_{wt}. The time-varying coefficients of the world market portfolio are comprised of a break parameter ϕ_{1j} for the indicator variable and a time-varying beta $\beta_{jt}^{v}|Z_{t-1}^{l,g}$ that depends on *local* leverage indicators Z_{jt-1}^{l} and *global* market conditions Z_{t-1}^{g}. Beta is linear in these instruments where γ_{0j} is a constant, γ_{lj} is a vector of coefficients for local leverage indicators consisting of P/B$_j$, P/E$_j$, and FX$_j$, and γ_{gj} is a vector of coefficients on global market conditions consisting of P/B-EM, P/E-EM, R-WD, and R-EM.[22]

Panel A of Table 5 presents country-by-country results of the TV-IAPM with structural breaks. Columns 1–3 show shifts in beta become less significant for many countries. Four countries, rather than one country in the benchmark case, do not exhibit a discrete shift in beta after the crisis at the 10% level. Columns 4–6 show leverage indicators and global market conditions are insignificant in predicting returns four countries. Columns 7–8 test for a structural break in each country, whether TV-IAPM holds and there is no structural break in each country. The introduction of information into beta did not weaken the evidence for structural shifts that is virtually the same as found with benchmark model tests.

Panel B reports tests of joint hypotheses with betas dependent on local indicators, global conditions, and both information sets. The column labeled None corresponds with benchmark results reported in panel A of Table 2. Significant shifts in betas remain across countries after introducing information into the dynamics of these betas. Benchmark results remain unchanged, regardless of the information set employed. Panel B also reports joint tests of information sets independently. Entries in this section of the table correspond to the fact that information sets can only be tested when they are also used in estimation. Tests indicate both information sets are potentially important to predicting return variation among these Asian countries.[23]

Overall, results indicate that significant shifts in betas remain across countries after introducing portfolios managed by leverage and world market conditions; however, the shifts in beta become less significant in many countries. Leverage indicators and global market conditions are important in predicting returns for

[22]There are three local instruments and four global instruments in each country's equation. This raises the number of coefficients in the system from 24 in the base case to 24 + 18 = 42 where only local instruments are used, 24 + 24 = 48 where only global instruments are used, and 24 + 18 + 24 = 66 where local and global instruments are used.

[23]The LR-tests lead to the same conclusion and are not reported.

TABLE 5

TV-IAPM with a Structural Break

Panel A. Individual Country Results

Country	1 α	2 ϕ_0	3 ϕ_1	4 $\gamma_l = 0$	5 $\gamma_g = 0$	6 $\gamma_l = \gamma_g = 0$	7 $\phi_0 = \phi_1 = 0$	8 $\alpha = \phi_0 = \phi_1 = 0$
ID	0.001 (0.779)	−0.031 (0.077)	4.334 (0.001)	0.533	0.280	0.280	0.000	0.000
KR	−0.001 (0.445)	−0.013 (0.298)	1.726 (0.118)	0.538	0.391	0.195	0.201	0.186
MY	0.001 (0.353)	−0.025 (0.015)	1.455 (0.281)	0.207	0.189	0.329	0.007	0.019
PH	0.003 (0.172)	−0.017 (0.032)	1.313 (0.154)	0.524	0.327	0.369	0.042	0.074
TW	0.002 (0.284)	−0.013 (0.011)	1.562 (0.002)	0.004	0.698	0.016	0.000	0.001
TH	−0.001 (0.638)	−0.032 (0.007)	1.437 (0.106)	0.000	0.009	0.000	0.000	0.000

Panel B. Joint Tests

			Wald Tests			
		df	Local	Global	$L + G$	None
H_1	$\alpha = 0$	6	0.593	0.467	0.604	0.464
H_2	$\phi_0 = 0$	6	0.027	0.050	0.044	0.080
H_3	$\phi_1 = 0$	6	0.007	0.008	0.000	0.003
H_4	$\phi_0 = \phi_1 = 0$	12	0.000	0.000	0.000	0.000
H_5	$\alpha = \phi_0 = 0$	12	0.155	0.123	0.184	0.212
H_6	$\alpha = \phi_0 = \phi_1 = 0$	18	0.000	0.000	0.000	0.000
	$\gamma_l = 0$	18	0.000		0.001	
	$\gamma_g = 0$	24		0.000	0.037	
	$\gamma_l = \gamma_g = 0$	42			0.000	

Table 5 presents results from estimating an asset pricing model (equations (8)) where beta is made time-varying with portfolios managed by local leverage indicators and global market conditions (TV-IAPM). This model is constructed to allow for a structural break at the time of the crisis in Thailand. Refer to Table 1 and Table 3 for descriptions of variables. Sample covers 10/05/90 to 06/05/98. Columns 1–3 in panel A show the constant term, α_j, and the coefficients ϕ_{0j}, and ϕ_{1j} that allow for a structural break in the TV-IAPM at the time of crisis. Columns 4–6 provide p-values from Wald tests that examine the importance of local leverage indicators $\gamma_l = 0$, global market conditions $\gamma_g = 0$, and both $\gamma_l = \gamma_g = 0$ in predicting each country's returns. Columns 7 and 8 present p-values from Wald tests of no structural break and test of whether TV-IAPM holds where the intercept equals zero and there is no structural break. Panel B reports the p-values with Wald tests of hypotheses H_1–H_6 (described in Table 2, panel B) involving all countries with betas depending on local indicators L, global conditions G, and both information sets $L + G$. The column labeled None corresponds with benchmark results (Table 2, panel B) and *df* denotes degrees of freedom. The last set of tests refers to the importance of local leverage indicators, global market conditions, and both to predicting returns in the system of equations.

the system as a whole. The introduction of information into beta does not weaken the evidence for structural shifts that is largely due to its failure to capture the change in average returns.

C. Portfolios Managed by Probability Assessments: P-IAPM

The second formulation for including information about leverage into the evolution of beta is based on a portfolio managed by an index of investor sentiment. It is a latent variable that gives an assessment of the likelihood of the financial crisis conditioned on leverage indicators and world market conditions.

In this formulation, beta is given as a linear projection on the investors' sentiment index. The empirical analog for equations (6) for the P-IAPM is

$$\text{(9a)} \qquad r_{jt} = \alpha_{jt}^{p}\left|Z_{\mathrm{TH}t-1}^{l,g}\right. + \phi_{0j} I_{\mathrm{TH}t} + \left[\beta_{jt}^{p}\left|Z_{\mathrm{TH}t-1}^{l,g}\right. + \phi_{1j} I_{\mathrm{TH}t}\right] r_{Wt} + \varepsilon_{jt},$$

$$\text{(9b)} \quad \text{where} \qquad \alpha_{jt}^{p}\left|Z_{\mathrm{TH}t-1}^{l,g}\right. = \alpha_{0j} + \alpha_{1j} p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right).$$

$$\text{(9c)} \qquad \beta_{jt}^{p}\left|Z_{t-1}^{l,g}\right. = \lambda_{0j} + \lambda_{1j} p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right), \qquad \text{and}$$

$$\text{(9d)} \qquad p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right) = \frac{\exp\left(\theta_{0\mathrm{TH}} + \theta_{l\mathrm{TH}} Z_{\mathrm{TH}t-1}^{l} + \theta_{g\mathrm{TH}} Z_{t-1}^{g}\right)}{1 + \exp\left(\theta_{0\mathrm{TH}} + \theta_{l\mathrm{TH}} Z_{\mathrm{TH}t-1}^{l} + \theta_{g\mathrm{TH}} Z_{t-1}^{g}\right)}.$$

The associated moment conditions are

$$\text{(9e)} \qquad E\left\{\begin{array}{c}\left(1, I_{\mathrm{TH}t}, p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right), r_{Wt}, p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right) r_{Wt}, I_{\mathrm{TH}t} r_{Wt}\right) \varepsilon_{jt} \\ \left(1, Z_{\mathrm{TH}t-1}^{l,g}\right)\left(I_{\mathrm{TH}t} - p\left(I_{\mathrm{TH}t}\left|Z_{\mathrm{TH}t-1}^{l,g}\right.\right)\right)\end{array}\right\} = 0, \qquad j = (1,6).$$

As before, the jth country's excess return r_{jt} is written in terms of the benchmark model and regressed on a constant α_j, an indicator variable I_{TH} for Thailand with ϕ_{0j} as its shift parameter, and the excess returns on the world portfolio r_{wt}. The time-varying intercept $\alpha_{jt}^{p}|Z_{\mathrm{TH}t-1}^{l,g}$ and beta in each country $\beta_{jt}^{p}|Z_{\mathrm{TH}t-1}^{l,g}$ depends on a constant and an investor's sentiment index of the Thai baht devaluation $p(I_{\mathrm{TH}t}|Z_{\mathrm{TH}t-1}^{l,g})$. The sentiment index is linked through a logistic relation to *local* leverage indicators for Thailand $Z_{\mathrm{TH}t-1}^{l}$ and *global* market conditions Z_{t-1}^{g}. The first set of moment conditions are normal regression equations, but the second set of orthogonality conditions are the first-order conditions from maximizing the likelihood of a logit model. It is straightforward to estimate this mixed equation system simultaneously.[24]

The extent to which the investor index model captures the shift in beta when the crisis started depends on how well the predicted crisis probability fits the actual crisis indicator. In Table 6, we examine logit models for each country separately using their own devaluation dates. The dependent variable is the indicator variable for each country devaluation date that is equal to zero before devaluation and one after devaluation. The independent variables are the local leverage ratios and the world market conditions all lagged one week.[25]

Logit results show the investor sentiment index fits the devaluation indicators quite well and that local leverage information, specifically the P/E and P/B ratios, contributes the bulk of the explanatory power. Local leverage indicators

[24]Simultaneous estimation of multiple Probits by GMM can be found in Bertschek and Lechner (1998). Maddala (1983) discusses Logit models and mixed estimation using Likelihood methods. The comparison of GMM and Logit routines confirms that estimates are identical and standard errors are nearly identical.

[25]The variables P/B-EM and P/B-EM are excluded from the estimation due to convergence problems.

TABLE 6

Logit Estimation of Investor Sentiment Index with Local Leverage and Global Market Conditions

		Local			Global		
Country	Constant	P/B	P/E	FX	R-WD	R-EM	R^2
ID	5.031	−0.221	−0.415	−1.908			0.258
	(0.000)	(0.812)	(0.000)	(0.668)			
KR	29.190	−42.267	0.049	26.961			0.776
	(0.000)	(0.000)	(0.157)	(0.000)			
MY	37.262	35.595	−6.641	−34.951			0.916
	(0.000)	(0.000)	(0.000)	(0.012)			
PH	178.779	−145.607	11.126	3.136			0.977
	(0.033)	(0.092)	(0.185)	(0.892)			
TW	6.881	−24.349	2.355	−15.644			0.473
	(0.000)	(0.000)	(0.000)	(0.567)			
TH	10.854	−7.606	−0.092	−18.722			0.814
	(0.000)	(0.000)	(0.007)	(0.154)			
ID	4.986	0.359	−0.496	0.523	27.165	−32.416	0.298
	(0.000)	(0.646)	(0.000)	(0.900)	(0.056)	(0.002)	
KR	62.022	−86.875	0.017	24.328	122.242	64.730	0.873
	(0.022)	(0.020)	(0.724)	(0.084)	(0.067)	(0.180)	
MY	44.049	39.707	−7.497	−57.825	−63.972	77.162	0.920
	(0.000)	(0.000)	(0.000)	(0.074)	(0.035)	(0.265)	
PH	199.279	−167.585	13.213	−1.384	−137.257	123.268	0.978
	(0.027)	(0.043)	(0.070)	(0.897)	(0.263)	(0.342)	
TW	7.099	−23.711	2.276	−3.236	46.652	−28.634	0.500
	(0.000)	(0.000)	(0.000)	(0.911)	(0.046)	(0.059)	
TH	13.082	−8.836	−0.130	−26.710	−55.582	−20.047	0.832
	(0.000)	(0.000)	(0.001)	(0.060)	(0.049)	(0.450)	

Table 6 shows coefficients, p-values (in parentheses), and pseudo R^2s from logistic models using local leverage and global market conditions. The dependent variable is each country's devaluation date where the indictor is zero before devaluation and one after devaluation. Table 1 reports devaluation dates and Table 4 describes independent variables, lagged one week.

have strong predictive power in discriminating pre- and post-devaluation periods with the majority having high R^2s. Surprisingly, exchange rates contributed the least to distinguishing the two periods. Global market conditions have a minor contribution.

With the P-IAPM, we seek to capture the changes in risk and return at the start of the crisis in Thailand where indicator values I_{TH} are very closely related ($R^2 = 83\%$) to the predicted index of investor sentiment. One expects the index for investor sentiment can substitute for the crisis indicator and therefore explain the changes in risk and return at the start of the crisis. We formally test whether the crisis indicator and the investor sentiment index are substitutes in an asset pricing framework by estimating the P-IAPM equations (9) and bootstrapping their critical values.

There are several reasons for bootstrapping critical values. In the mixed model that we estimate, the sampling distributions of test statistics have well-behaved large sample properties. However, it may take a very large sample to discriminate between the P-IAPM and the benchmark model nested in equations (9) because they are such close substitutes due to the high R^2 found earlier. Also, the small sample distribution of test statistics where the indicator variable is endogenous and appears in other equations is not known. The distributions of test statistics are derived by performing fixed in repeated samples bootstrap (i.e., drawing errors only) with 3000 replications under each hypothesis. The indicator variable $I_{\text{TH}t}$ is the only dependent variable that appears on the right-hand side. This vari-

able is constructed by adding the newly drawn error to the original predicted probability index, assigned a one if the result is greater than or equal to 0.5 and zero otherwise. Drawing their errors and adding back their predicted values as modified by the new indicator variable and the restrictions imposed by each hypothesis form the returns. After each set of replications, test values for each hypothesis are ranked to determine critical values.[26]

Table 7 reports values of Wald, J-statistics, and LR[27] tests and corresponding large sample and bootstrap critical values of the six hypotheses reported before and an additional three hypotheses specific to the P-IAPM. Although the results are mixed if conventional p-values are used, bootstrap confidence intervals are much wider and do not reject that the benchmark and P-IAPM models serve as substitutes. The values from bootstrapped distributions are large. There are very few rejections. This may reflect that when two models are close substitutes their critical values become very sensitive to sample size. The Wald statistic appears the most sensitive to sample size. Stars indicate the rejections we found in our sample using bootstrapped values for each confidence level. None of the Wald and J-statistic tests reject at the 95% level and very few of the LR tests reject at the 95% level.

The index for investor sentiment explains the changes in risk and return we see at the start of the crisis. We do not reject that the shift parameter coefficients are zero (i.e., H_4: $\phi_0 = \phi_1 = 0$).[28] As expected, the close relationship between the crisis indicator and the predicted index of investor sentiment makes these two models statistical substitutes.

Figure 1 presents a comparison of betas from Thailand based on the P-IAPM, TV-IAPM, the constant beta model, and the benchmark model with structural shifts around the time of the crisis. The P-IAPM roughly follows the shift in beta depicted in the benchmark model. Not only does the P-IAPM track the shift in beta after the crisis, but it also predicts an increase in beta in the months leading up to the crisis date; the TV-IAPM does not. Based on predictions of the P-IAPM alone it would be difficult to conclude that the investor's sentiment "predicts" the increase in beta before the crisis began. Crises, by their nature, are infrequent events and all estimation takes place in sample. However, news of speculation as early as February, the Thai bhat intervention in May 1997, and mounting bad loans in the Thailand real estate sector support these predictions. This evidence indicates that investors' reassessment of risk and return preceded the official start of the crisis.

D. Cross-Sectional Evidence

Table 8 presents evidence of leverage effects in a cross-section of 589 firms from our six sample countries obtained from the EMDB. Pre- and post-crisis be-

[26]Refer to Efron (1979) and Nelson and Kim (1993) for a discussion of resampling methods. Fair (2002) adapts a bootstrapping method similar to ours for large non-linear macro models.

[27]Even though the errors may not be normally distributed, the maximum likelihood estimate of the covariance matrix is consistent and asymptotically normal. See Hayashi (2000) for computation and discussion.

[28]Symmetrically, the benchmark model can substitute equally well for the P-IAPM (i.e., H_9: $\lambda_1 = \alpha_1 = 0$, in the presence of non-zero benchmark model parameters ϕ_0 and ϕ_1).

TABLE 7

P-IAPM with Structural Break

Hypothesis	df	Stat	Test Value	Conventional p-Values	Bootstrap Critical Values 90%	95%	99%
H_1 $\alpha = 0$	6	Wald	4.5	0.611	11.2	13.6	17.4
Before crisis		J-stat	4.4	0.620	10.2	11.9	15.6
intercepts zero		LR	0.8	0.992	6.6	15.2	55.7
H_2 $\phi_0 = 0$	6	Wald	62.7	0.000	57.6[a]	75.9	120.2
No change in intercept		J-stat	21.5	0.001	26.0	29.5	35.8
after crisis		LR	19.5	0.003	17.3[a]	23.6	42.6
H_3 $\phi_1 = 0$	6	Wald	23.0	0.001	42.8	59.2	99.8
No discrete shift in		J-stat	10.7	0.097	20.5	23.2	29.1
beta after crisis		LR	27.0	0.000	20.2[a]	24.6[a]	36.3
H_4 $\phi_0 = \phi_1 = 0$	12	Wald	182.7	0.000	185.0	244.3	488.0
No structural change in		J-stat	25.8	0.011	43.2	46.3	53.5
ICAPM relationship		LR	62.6	0.000	41.7[a]	68.0	138.0
H_5 $\alpha_0 = \phi_0 = 0$	12	Wald	68.3	0.000	64.6[a]	83.7	129.3
Intercepts zero before		J-stat	27.0	0.008	32.8	36.7	44.9
and after crisis		LR	34.1	0.001	36.0	79.5	443.2
H_6 $\alpha_0 = \phi_0 = \phi_1 = 0$	18	Wald	193.7	0.000	199.4	259.4	495.0
No structural change with		J-stat	31.2	0.028	50.5	54.3	61.7
intercepts zero		LR	58.3	0.000	77.4	117.3	237.9
$H_7 \alpha_{0j} = \alpha_0$ and $\alpha_{1j} = \alpha_1$	10	Wald	12.8	0.025	32.5	40.7	61.1
Zero beta P-IAPM holds		J-stat	11.6	0.040	24.3	27.4	34.0
		LR	58.2	0.000	22.9[a]	35.2[a]	109.8
H_8 H_4 and H_7	22	Wald	227.1	0.000	189.5[a]	260.7	686.1
Zero beta P-IAPM holds		J-stat	36.0	0.005	51.3	55.3	66.7
w/ no structural change		LR	321.4	0.000	118.8[a]	189.2[a]	389.1
H_9 $\lambda_1 = \alpha_1 = 0$	12	Wald	25.6	0.012	67.0	83.9	160.7
Test for logit		J-stat	14.8	0.253	34.2	37.3	44.8
significance		LR	51.7	0.000	54.9	72.3	113.3

Table 7 presents p-values from Wald, J-stat, and LR tests of the P-IAPM (equations (9)). Refer to Tables 1 and 3 for descriptions of variables. Sample covers 10/05/90 to 06/05/98. The nine hypotheses are tested on the constant α_0, the coefficients ϕ_0, and ϕ_1 that allow for a structural break in the P-IAPM at the time of crisis, and the investor sentiment parameters λ_1 and α_1. Degrees of freedom is denoted by df. The left-hand columns of the table report the actual values of Wald, J-stat, and LR tests of each hypothesis and their large sample p-values. The right columns contain bootstrapped critical values of test statistics for conventional confidence levels.

[a]indicates rejection.

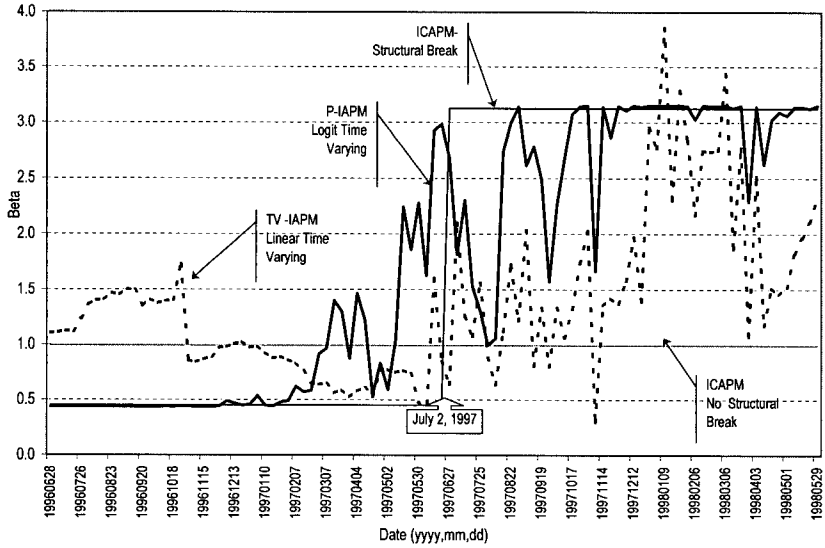
tas, pre- and post-crisis averages of returns, and the valuation ratios of each firm are used as data and grouped by market value quintile. Panel A shows the sample average for each quintile in the pre-crisis sample and the average difference between pre- and post-crisis samples (denoted by Δ).[29] Consistent with time-series evidence, average returns drop, beta increases, and P/B and P/E ratios fall across quintiles after the start of the crisis.

Panel B of Table 8 reports coefficients and their significance levels by quintile from regression models involving the firm-level change in average excess U.S. dollar return from pre- to post-crisis samples against the change in beta and average P/B and P/E ratios from pre- to post-crisis samples. We also scale the extreme P/B and P/E ratios found in firm-level emerging market data by taking natural

[29]Firms with non-positive average P/B ratios pre- or post-crisis are filtered out to remove extreme values, reducing the sample to 568 firms. Quintiles are formed on pre-crisis average market value. Results in panels A and B are insensitive to whether pre-crisis or overall sample average market values are used to create quintiles.

FIGURE 1

Thailand: Exposure to Systematic Risk around Crisis Date



logs.[30] Model I attempts to explain the cross-sectional differences in pre- and post-crisis average return with changes in pre- to post-crisis beta. Models II and III assess the effects of adding the change in P/B and P/E ratios. Overall, evidence indicates the change in P/B is consistently significant and has the right sign. This confirms what is found in the time series where a fall in the P/B ratio is associated with lower average returns. As Model II shows, the change in P/B is significant in all but the smallest quintile. The change in beta is only significant in the smallest quintile with the right sign; as beta rises average returns fall. Adding the P/E ratio in Model III restricts the sample and shows only a marginal effect in the largest quintile with the right sign.

VI. Conclusion

This paper investigates changes in risk and return in six Asian markets before and after the 1997 Asian financial crisis began. Using a six-country sample of weekly equity market returns from October 1990 to June 1998, we find that while betas increased more than fivefold after the crisis began, average returns fell more than sevenfold. We address an interesting puzzle where risk has increased and average returns fall. When beta increases, capital losses must be suffered to make higher expected returns possible. These capital losses are part of observed returns.

Leverage is a key feature of financial crisis models. Prior to a crisis, firms are highly levered with short-term, mostly unhedged, dollar denominated debt. Once

[30]It is common practice to consider only positive values of valuation ratios. As such, Models I–III contain 589, 568, and 474 firms overall.

TABLE 8

Cross-Sectional Evidence

Panel A. Average before Crisis Date and Change in Sample Average Post-Crisis Date

		R	ΔR	β	$\Delta\beta$	P/B	ΔP/B	P/E	ΔP/E
Small	1	−0.0025	−0.0215	0.077	0.746	7.800	−1.481	43.50	−29.61
	2	−0.0010	−0.0148	0.143	0.266	4.288	−1.863	28.94	−14.51
	3	0.0011	−0.0163	0.158	0.190	3.426	−1.454	34.40	−34.36
	4	0.0009	−0.0202	0.139	0.540	2.447	−0.920	35.52	−21.45
Large	5	0.0012	−0.0229	0.057	1.065	2.309	−0.904	36.90	−16.14
Overall		−0.0001	−0.0191	0.115	0.559	4.035	−1.328	35.77	−23.18

Panel B. OLS Regressions—Dependent Variable Δr is Change in Average Excess Return

		Small				Large	
Model		1	2	3	4	5	Overall
I	Const.	−0.0201	−0.0134	−0.0165	−0.0197	−0.0225	−0.0184
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
	$\Delta\beta$	−0.0019	−0.0027	0.0013	−0.0006	−0.0001	−0.0009
		(0.069)	(0.139)	(0.199)	(0.242)	(0.849)	(0.106)
	$\overline{R}^2$	0.053	0.080	0.039	0.006	−0.008	0.014
II	Const.	−0.0185	−0.0101	−0.0138	−0.0148	−0.0158	−0.0151
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
	$\Delta\beta$	−0.0020	−0.0005	0.0012	−0.0006	0.0000	−0.0005
		(0.059)	(0.501)	(0.254)	(0.321)	(0.970)	(0.214)
	Δ ln(P/B)	0.0014	0.0050	0.0034	0.0063	0.0106	0.0043
		(0.579)	(0.015)	(0.090)	(0.000)	(0.000)	(0.000)
	$\overline{R}^2$	0.057	0.064	0.074	0.124	0.087	0.050
III	Const.	−0.0146	−0.0124	−0.0138	−0.0148	−0.0153	−0.0152
		(0.003)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
	$\Delta\beta$	−0.0025	−0.0002	0.0005	−0.0008	0.0000	−0.0007
		(0.019)	(0.792)	(0.651)	(0.234)	(0.968)	(0.088)
	Δ ln(P/B)	0.0018	0.0032	0.0023	0.0078	0.0103	0.0036
		(0.633)	(0.359)	(0.322)	(0.000)	(0.000)	(0.007)
	Δ ln(P/E)	0.0018	0.0014	0.0010	−0.0021	0.0028	0.0010
		(0.694)	(0.515)	(0.291)	(0.271)	(0.095)	(0.330)
	$\overline{R}^2$	0.126	0.013	0.015	0.126	0.101	0.047

Panel A presents sample averages for U.S. dollar returns, CAPM betas (β), and P/B and P/E ratios before the start of the crisis and changes in sample averages from pre- to post-crisis (denoted by Δ) by size quintile. Quintiles are determined by average market value prior to crisis. There are 568 of 589 firms in this sample that have positive average P/B ratios. Data of individual firms are weekly from sample period defined in Table 1. Panel B contains regression coefficients and p-values (in parentheses) using White corrected standard errors by quintile of the average change in excess return of each firm from pre- to post-crisis (Δr) on changes in average beta ($\Delta\beta$ in model I, and then adding the change in log of the average P/B ratios, Δ ln(P/B), and P/E ratios, Δ ln(P/E), in models II and III, respectively. Models I–III contain 589, 568, and 474 firms overall.

monetary authorities devalue, leverage and interest payments increase. Familiar valuation ratios can be used to reveal information about the effect of leverage on discount rates, as leverage is difficult to measure directly. The increased leverage contributes to the rise in equity betas and raises expected returns. Valuation ratios decline as prices reflect the rise in discount rates. Capital losses are necessarily created to make higher expected returns possible.

Information from news reports suggests that investors do anticipate changes in risk and return before the crisis officially starts and update these expectations in the post-crisis period. Statistical break point tests find significant breaks five or more months after the crisis officially started—at the peak of the crisis.

Our model predicts an increase in beta months before the crisis officially started. Tests indicate this model, which uses portfolios managed by an index of investor sentiment, can explain the large shifts in beta and average return after the crisis started. Familiar valuation ratios tell the story that when leverage increases, beta will increase and average return falls because capital losses are necessary for

higher expected returns. We provide evidence consistent with the predictions of financial crisis models where leverage plays a key role. In addition, we show that changes in leverage indicators can explain cross-sectional differences in average returns.

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