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Speculative Attacks in the Asian Crisis

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Abstract

This paper takes the Asian crisis as an example to show that the Autoregressive Conditional Hazard (ACH) model is a powerful tool for studying the time series features of speculative attacks. The ACH model proposes a duration variable to capture the changes in the frequency of attacks, which might be an important factor influencing investors' expectations. The empirical results show that the ACH model explains the crisis far better than the Probit model. The duration variable is highly significant while most fundamentals are not. The contagion effect is tested and accepted under the ACH specification.

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I. Introduction

Since the start of the Asian financial crisis, many economists have tried to explain the magnitude and speed with which it occurred. Radelet and Sachs (1998a, b) emphasize both the self fulfilling nature of the crisis and the role of "fundamentals" such as external shocks, growing short term debt, and expanding bank credit. Corsetti, Pesenti, and Roubini (1998a, b) focus on the moral hazard problem pervasive in the East Asian countries. Chang and Velasco (1998) develop a bank run model similar to Diamond and Dybvig (1983) and argue that the root of the crisis is the inability of the governments to act as lender of last resort for debt denominated in foreign currency. Burnside, Eichenbaum, and Rebelo (1999) show that prospective deficits, arising from governments' commitment to bail out troubled banks, may destroy investors' confidence in the current exchange rate. These papers focus on modeling a crisis within one economy but do not address an important feature of the Asian crisis -- countries with quite different fundamentals were hit by the crisis around the same time. As Krugman (1998) points out, a class of "third generation" models is needed to explain what happened in Asia in terms of an apparent phenomenon of contagion.

The development of contagion theory has been impressively fast in the 1990s. Theoretical research shows that contagion could happen through three channels: global shocks, weak fundamentals, or "pure" contagion. In the global shock models, common external shocks, such as the changes of oil price in 1973 and 1979, could trigger crises in different countries. Masson (1998) refers to this phenomenon as "monsoons". The "fundamental based contagion" (Kaminsky and Reinhart, 1999) is also called "spillovers" (Masson, 1998). Eichengreen, Rose and Wyplosz (1996b) show how a currency crisis in one country can have a real effect on the economy of its trade partners. The third mechanism covers all instances not included in the above two cases. In these models, the contagion could be caused by herding behavior of investors (Calvo and Mendoza, 2000), by a shift in investor's expectation (Masson 1998, Rodrik and Velasco 1999), or through a "liquidity squeeze" effect (Valdes, 1998). In section 3, I review the empirical literature of contagion.

Despite the fast growth of contagion theory and the obvious phenomenon of contagion in the Asian crisis, there has been little empirical work done to test for a contagion effect in the Asian crisis. Radelet and Sachs (1998b) and Tornell (2000) show that the spread of crisis is determined by the cross-country variation of fundamentals. However, they do not test contagion explicitly in their models. Testing for a contagion effect requires modeling the temporal variation of fundamentals and the crisis. It can not be captured by a cross-country regression.

Little effort has been made to explain the Asian crisis using a time series model. As it is shown later, traditional econometric techniques can not give a satisfactory answer. Hamilton and Jorda (2000) proposed an Autoregressive Conditional Hazard (ACH) specification to model changes in the federal funds rate. This paper utilizes the ACH model to illustrate that the duration dynamics played a much more important role than the "fundamentals" in the Asian crisis. The duration approach emphasizes the role of a duration variable, which measures how frequently speculative attacks happened in the past. The frequency of past speculative attacks has been neglected in previous literature, despite its obvious importance in influencing investors' behavior in herding models and models that emphasize a "shift in expectation". The performances of the ACH model and a probit model are compared and the ACH model gives a much better explanation of the crisis in terms of the log likelihood.

The ACH specification offers a straightforward way to test for a contagion effect in the Asian crisis. The contagion effect is strongly supported by this model. Compared to the fundamentals, the regional duration dynamics played a dominant role in the crisis that happened in Indonesia, Korea, and Thailand.

The paper is organized as follows. Section 2 discusses how to measure speculative attacks. Section 3 reviews the empirical literature on contagion. Section 4 reviews the specification of the ACH model. Section 5 applies the ACH model to data from Asian countries and shows the results. Section 6 concludes and provides direction for future research.

II. Identifying Speculative Attacks in the Asian Crisis

Testing for contagion requires testing if the probability of a currency being attacked in one period is influenced by knowing the history of speculative attacks on all currencies in the sample, even after controlling for fundamentals. In order to carry out the test, one needs a measure of speculative attacks. There are three approaches to identify speculative attacks. The first approach is identify attacks qualitatively, by simply plotting the exchange rate and picking up the sharp jumps, or by citing journalistic record and IMF reports. One example is Blanco and Garber (1986). This approach works when the researcher is interested in analyzing one or several of crises that are well documented.

The second approach is proposed by Frankel and Rose (1996) They define a "currency crash" as "a nominal depreciation of the currency of at least 25% that is also at least a 10% increase in the rate of depreciation". The requirement on the increase in the rate of depreciation is a device to screen out high inflation countries that depreciate their currencies year after year. The third approach to solve the problem is to construct an index that summarizes changes of exchange rates, reserves, and interest rate. This index was first proposed by Eichengreen, Rose and Wyplosz (1996 a) (ERW index hereafter). It is constructed in the following way:

$$K_t = w_1 \Delta e_t - w_2 \Delta r_t + w_3 \Delta i_t \tag{2.1}$$

where K_t is the ERW index, Δe_t, Δr_t and Δi_t are percentage changes in exchange rate, reserves and interest rate, respectively. The parameters w_1, w_2, w_3 are the inverse of the standard deviations of Δe_t, Δr_t and Δi_t. Eichengreen, Rose and Wyplosz claim that this index captures the "speculative pressure" in both successful attacks (changes in exchange rates) and unsuccessful attacks (changes in reserves and interest rate). Using this index, they construct a dummy variable to identify the speculative attacks in EMU. The crisis dummy for period t takes the value of 1 if

$$K_t > 2 \times \sigma_k + \overline{K} \tag{2.2}$$

where σ_k is the standard deviation of K_t, $\overline{K}$ is the mean of K_t.

This index approach has been widely used in the currency crisis literature. Some people model the pressure index directly (Sachs, Tornell, and Velasco 1996). Others use it to identify the crisis and check the behavior of fundamental variables during tranquil periods and pre-crisis periods (Eichengreen, Rose and Wyplosz 1996 a, Kaminsky and Reinhart 1999, 2000). There are two problems for this index approach. First, as Flood and Marion (1998) point out, if Krugman (1979) is right, then "two out of three ERW indicators point in the wrong direction at the devaluation time". "Immediately following the devaluation, interest rates will fall back…Reserves will flow back…". That means if the devaluation is anticipated, the changes in reserves and interest rates may cancel out at least some of the changes in exchange rate (in equation 1, a positive Δr_t and a negative Δi_t would dampen the effect of a positive Δe_t on K). As a consequence, the ERW approach may fail to identify the attacks in the devaluation period. This is exactly what happened in the Philippines in September 1997. The Philippine currency depreciated by more than 12%, but the ERW index fails to count this month as a period of speculative attack because the reserves increased by 6%.

The second problem comes from the threshold value in the index. Eichengreen, Rose and Wyplosz (1996 a) use two standard deviations above the mean as the threshold. If the sample is large and there are different regimes in the sample, then a high volatility regime will dominate the whole sample. The threshold will be too high to identify a crisis that happened in the low volatility regime. This is why the ERW index fails to identify May 1997 as a crisis period in Thailand. The exchange rate and reserves were highly volatile in the 1980s, but relatively tranquil in the 1990s. A fixed threshold can not address the regime shifts in the exchange rate market.

It is not the goal of the ERW approach to track the evolution of crises. In most empirical work using the ERW index, when a crisis is identified in one month, the half-year following that month is regarded as a crisis period. One can use it to identify crises for cross-country regression as in Sachs, Tornell, and Velasco (1996). But for time series analysis, one needs a method to capture the evolution of crisis. In other words, it is critical to know whether a currency is under high speculative pressure for every month in the sample.

I propose an alternative way to identify speculative attacks in East Asia using monthly data. The interest rate data are not available for the whole sample, so I only consider exchange rates and reserves. My method is to identify extreme values in reserves and the exchange rate separately. Specifically, the dummy variable takes the value of 1 if

$$\Delta e_t > 3 \times \sigma_{\Delta e,t} + M_{\Delta e,t} \tag{2.3}$$

or

$$\Delta r_t < -3 \times \sigma_{\Delta r,t} + M_{\Delta r,t} \tag{2.4}$$

where $\sigma_{\Delta e,t}$ is the standard deviation of Δe in the sample of (t-36, t-1). $M_{\Delta e,t}$ is the mean of Δe in the same sample. $\sigma_{\Delta r,t}$ and $M_{\Delta r,t}$ are the corresponding parts for reserves. In other words, I identify speculative attacks as periods when changes in exchange rate or reserves take extreme values. In order to define an extreme value, the change in the exchange rate (or reserves) in one period is compared with changes in the previous 3 years. The time varying feature of the threshold is designed to avoid the regime changes.

I claim that this method solves the problems encountered when using the ERW index. First, the Flood-Marion argument does not apply to the moving window method. When a crisis is anticipated, the speculative attacks in the devaluation period will be identified because it satisfies equation 2.3. A capital inflow will not affect equation 2.3. Utilizing Δe_t and Δr_t separately, this technique successfully identified the attack against the Philippines' currency in September 1997. Second, the 3-year moving window addresses the problem of regime shifts. For example, in the case of Thailand, the threshold in May 1997 is not affected by the high volatility regime in 1980s. Therefore, the moving window method captured the speculative attack against Thailand in May 1997. The results from the moving window method and from the ERW index are compared in Table 1 and Table 2. The moving window method gives results that are more consistent with the IMF report (1997) on the Asian financial crisis.

Table 1. Speculative Attacks in 1997 Identified by the ERW Index

	Indonesia	Korea	Malaysia	Philippines	Thailand
January	0	0	0	0	0
February	0	0	0	0	0
March	0	0	0	0	0
April	0	0	0	0	0
May	0	0	0	0	0
June	0	0	0	0	0
July	0	0	1	1	1
August	1	0	1	0	1
September	0	0	1	0	0
October	1	0	1	0	1
November	0	1	0	0	1
December	1	1	NA	NA	1

III. Empirical Literature on Contagion

This section reviews several papers relevant to testing theories about contagion. Emphasis is given to papers that either directly study contagion in the Asian crisis or give useful clues for studying that issue.

Table 2. Speculative Attacks in 1997 Identified by A Three Year Moving Window Approach

	Indonesia	Korea	Malaysia	Philippines	Thailand
January	0	0	0	0	0
February	0	0	0	0	0
March	0	1	0	0	0
April	0	0	0	0	0
May	0	0	0	0	1
June	0	0	0	0	0
July	1	0	1	1	1
August	1	0	1	0	1
September	0	0	1	1	0
October	1	1	0	0	0
November	0	1	0	0	1
December	1	1	1	1	1

The global shock models have not been rigorously tested in the empirical literature. In the case of Asian crisis, the global shock models can not provide a satisfactory explanation of the contagion phenomenon. Although some global variables (the shift of US interest rates, the continued recession of Japan's economy) had some effects on East Asia, those effects are not large enough to explain a crisis of the magnitude observed in 1997.

The "fundamental based" contagion models have been tested and supported by Eichengreen, Rose and Wyplosz (1996 b), Glick and Rose (1999), Sachs, Tornell and Velasco (1996), and Tornell (2000). Eichengreen, Rose and Wyplosz (1996 b) use panel data to test the existence of contagion effect by estimating the following model:

$$Crisis_{i,t} = \omega D(Crisis_{-i,t}) + \lambda Macro_{i,t} + \varepsilon_{i,t} \qquad (3.1)$$

where $Crisis_{i,t}$ is a dummy variable for country i at time t, constructed by using the ERW index; $D(Crisis_{-i,t})$, the contagion variable, takes the value of 1 if the crisis dummy for any country in the sample other than i is 1, 0 otherwise; $Macro_{i,t}$ contains current and lagged macro variables for country i. Eichengreen, Rose, and Wyplosz use quarterly data for 20 industrial countries from 1959 through 1993. They find that ω is significant in that regression. The second step of their analysis is to test which channel the contagion spreads through. They estimate the model:

$$Crisis_{i,t} = \omega \sum_{j \neq i} W_{ij,t}(Crisis_{j,t}) + \lambda Macro_{i,t} + \varepsilon_{i,t} \tag{3.2}$$

where $W_{ij,t} = w_{ij,t}$ if $Crisis_{j,t} = 1$ for any $j \neq i$. $w_{ij,t}$ is a weight which represents the "relevance" at time t of country j for country i. They find evidence to support the weighting scheme that $w_{ij,t}$ reflects the trade link between two countries. Glick and Rose use data from several crises (including the Asian crisis in 1997) and draw the same conclusion.
Sachs, Tornell, and Velasco argue that the Tequila effect in 1995 hit countries with weak fundamentals harder. Their benchmark regression is:

$$\begin{aligned} IND_i = {} & \beta_1 + \beta_2(\Delta RER_i) + \beta_3(LB_i) + \beta_4(D^{LR}{}_i \times \Delta RER_i) \\ & + \beta_5(D^{LR}{}_i \times LB_i) + \beta_6(D^{LR}{}_i \times D^{WF}{}_i \times \Delta RER_i) \\ & + \beta_7(D^{LR}{}_i \times D^{WF}{}_i \times LB_i) + \varepsilon_t \end{aligned} \tag{3.3}$$

where IND is the ERW index, ΔRER is real exchange rate misalignment, measured by the percentage deviation of the average real exchange rate index over 1986-89 from the average over 1990-94, LB is a lending boom variable measured by the ratio of banks loans to GDP, D^{LR} is a dummy variable for countries with low reserves, D^{WF} is a dummy variable for countries with weak fundamentals. The fundamentals they use to construct D^{WF} include ΔRER, LB, and $M2/R$. $M2/R$ is the ratio of M2 to reserves. The effects of fundamentals on the crisis index in countries with low reserves ($D^{LR} = 1$) but strong fundamentals ($D^{WF} = 0$) are given by $\beta_2 + \beta_4$ and $\beta_3 + \beta_5$. These effects turn out to be close to zero in their model. Tornell (2000) extends their method to the Asian crisis and confirms their finding.
Edwards (2000) gives a definition of "pure" contagion in the following equation:

$$X_{jt} = \alpha + \lambda Y_{jt} + \beta G_t + \gamma \sum_{k \neq j} Z_{kt} + \varepsilon_t \tag{3.4}$$

where X_{jt} is some variable measuring crisis for country j, Y_{jt} is a vector of domestic variables for country j, G_t is a vector of global variables, Z_{kt} is a vector of variables from a related county k. He argues that the contagion effect is captured by three terms in the above equation: βG_t is the global factor, $\gamma \sum_{k \neq j} Z_{kt}$ is the spillover effect, and ε_t is the "pure" contagion effect.

IV. The Autoregressive Conditional Hazard Model

This section briefly reviews the autoregressive conditional hazard (ACH) model developed in Hamilton and Jorda (2000). The ACH model estimates the probability of an event (a speculative attack in this application) that would happen in a given period of time. The logic behind the ACH specification is intuitive. Suppose one had a prediction of the expected length of time that would pass until the next speculative attack. For example, perhaps the expected duration is 8 months. If an attack at any point in the future were deemed equally likely, then the possibility of an attack next month would be 1/8. With Poisson arrival times, the expected duration is the reciprocal of the hazard rate.

Engle and Russell (1998) suggested that a natural way to forecast the expected duration until the next event is to use a distributed lag on recent past observed durations. Hamilton and Jordon (2000) suggested that the reciprocal of this magnitude is a logical starting point for a prediction of the probability of an event within the next month. The mathematical representation of the above logic is as follow.

Define x_t to be a random variable that takes on the value of one if the pressure index goes above the threshold during month t and zero otherwise. Let $\{w_{1t}\}, t = 1,2,\ldots,T,$ be a sequence that, for any month t, records the most recent month in which the index breaks the threshold:

$$w_{1t} = \begin{cases} t & \text{if } x_t = 1 \\ \mathrm{w}_{1,\mathrm{t-1}} & \text{if } x_t = 0 \end{cases} \qquad \text{for } t = 1,2,\ldots,T \tag{4.1}$$

Let w_{2t} denote the month the index breaks the threshold before that:

$$w_{2t} = \begin{cases} w_{1,t-1} & \text{if } x_t = 1 \\ w_{2,t-1} & \text{if } x_t = 0 \end{cases} \qquad \text{for } t = 1,2,\ldots,T \tag{4.2}$$

In this notation, $w_{1,t-1} - w_{2,t-1}$ measures the length of the most recent tranquil period. Let ψ_t denote the expectation of $w_{1,t} - w_{2,t}$. In the absence of any other explanatory variables, ψ_t is the expectation of how long it will be until the next attack. The specification of ψ_t under the ACD (1,1) model is:

$$\psi_t = \omega + \alpha d_t + \beta \psi_{t-1} \tag{4.3}$$

where

$$d_{t-1} = w_{1,t-1} - w_{2,t-1}. \tag{4.4}$$

Note that ψ_t is a step function that only changes when a new event was observed the preceding month, i.e., only when $x_{t-1} = 1$.

Now define the hazard rate, h_t, as the conditional probability that the index breaks the threshold in the month t given Y_{t-1}, which represents information observed as of month $t-1$:

$$h_t = P(x_t = 1 \mid Y_{t-1}) \tag{4.5}$$

The generalized expression for h_t proposed by the ACH (1,1) model is

$$h_t = \frac{1}{\psi_t + \delta' z_{t-1}} = \frac{1}{\omega + \alpha d_t + \beta \psi_{t-1} + \delta' z_{t-1}} \tag{4.6}$$

where z_{t-1} denotes a vector of additional variables that is known at time t-1. Hamilton and Jorda (2000) use a smoothing transform function[2] to make sure that the hazard h_t lies between 0 and 1.

V. Test for the Contagion Effect

In this section I apply the ACH (1,0) model to the data from countries that were hit by the Asian crisis most seriously[3]. I use monthly data from 1993:12 to 1997:12. The four countries I include in my benchmark model are Indonesia, Korea, the Philippines, and Thailand. First, I check if the fundamentals found useful in cross-country regressions can also explain what happened in Asia in time series models. I stack the data from the four countries and estimate a probit model. A crisis dummy $Crisis_{it}$ is constructed using the method discussed in section 2. The fundamentals I include in the probit model are the real exchange rate (RER), inflation (INF), the ratio of M2 to reserves (MRES), and the growth rate of MRES (MREG). Many researchers have pointed out the significance of these variables in currency crises. As noted in section 3, Sachs, Tornell, and Velasco (1996) prove the RER, MRES, and MREG variables are important in the transmission of the Tequila effect. Tornell (2000) finds these variables are important fundamentals determining the spread of the Asian crisis. The importance of the real exchange rate is also emphasized in Dornbusch, Goldfajn, and Valdes (1995). Inflation is regarded as an important term in the objective function of the government. Its variation may affect the government's willingness to defend the exchange rate (Obstfeld 1994, 1996).

The result from the probit model is reported in the second column of Table 3. The model gives a log likelihood of –60.42. Three out of four fundamentals (MRES, MREG, and RER) are not significant. Only INF is significant at 5% level.

Can the ACH specification do better? Before switching to the ACH reciprocal specification, I compare its performance to that of the probit model. To make it a fair

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[2] The estimation of the ACH model requires setting a parameter Δ for a smoothing function. Following Hamilton and Jorda (2000), I set the parameter to be 0.1.

[3] We also explored an ACH (1,1) specification, but the beta coefficient was never significant. The ACH (1,0) model is adequate for the application here.

comparison, only the four "fundamental" variables are included in the ACH model. The specification is

$$P(Crisis_{i,t} = 1) = \frac{1}{C + \gamma_1 MRES_{i,t-1} + \gamma_2 MREG_{i,t-1} + \gamma_3 RER_{i,t-1} + \gamma_4 INF_{i,t-1}} \tag{5.1}$$

Notice that the duration variables are not included in model 5.1. Therefore, the discrepancy of the performances only comes from the difference in the functional forms. The result is reported in column 3 of Table 3. The ACH functional form has a better fit to the data. The log likelihood from the ACH model is -58.35, higher than that from the probit model. As for the parameters, the ACH model is consistent with the probit model. γ_1, γ_2, and γ_3 show large standard errors. Only the parameter for INF, γ_4, is significant.

Table 3. Results for One Probit Model and Five ACH (1,1) Models

Variable	Probit	ACH 5.1	ACH 5.2	ACH 5.3	ACH 5.4	ACH 5.5
C	-0.17 1/	-11.45	0.82	-0.87	-0.87	-0.19
	0.72 2/	9.29	0.33	0.69	N/A	3.69
$d_{i,t-1}$			0.15*		0	
			0.03		N/A	
d_{t-1}				1.57*	1.57	1.61*
				0.57	N/A	0.60
$MRES_{i,t-1}$	-0.02	-2.23				0.33
	0.14	1.73				0.32
$MREG_{i,t-1}$	0.06	-0.84				0.02
	0.08	0.47				0.25
$RER_{i,t-1}$	0.01	1.28				-0.31
	0.07	0.86				0.30
$INF_{i,t-1}$	-0.17* 3/	5.17*				-0.19
	0.08	2.53				0.36
Likelihood	-60.42	-58.35	-54.84	-45.37	-45.37	-44.77

1/ Large numbers are estimates for the coefficients. Small numbers are standard errors.
2/ Stared numbers are those with t-statistics greater than 2.
3/ The standard errors for ACH 5.4 are not available because the optimization run into the boundary (the parameters for duration variables are confined to be nonnegative).

The most important feature of the ACH specification is not the functional form, but the duration dynamics introduced by the d_t variables. How do the duration dynamics contribute to predicting the speculative attacks? First, I estimate an ACH (1,0) model without fundamentals.

The specification is

$$P(Crisis_{i,t} = 1) = \frac{1}{C + \alpha d_{i,t-1}} \tag{5.2}$$

where $d_{i,t}$ is the *country specific* duration variable which is constructed based on $Crisis_{i,t}$. The $d_{i,t}$ series is shown in Table 4.

This is a challenging test for the ACH specification because no fundamentals are included in model 5.2. The result is shown in column 4 of Table 3. The performance of the duration variables is impressive. Although Model 5.2 only contains one explanatory variable, it gives a better fit compared to the first ACH model. The log likelihood of the ACH (1,0) model is –54.84, higher than model 5.1 by 4. Since both models employ the same functional form, this improvement in the log likelihood shows that the duration dynamics explain the crisis significantly better than the fundamentals. The α parameter in model 5.2 is 0.15 and highly significant.

The duration variable $d_{i,t}$ is based on the speculative attacks only in country i. On the other hand, contagion theory suggests that the crisis in one country may be caused by the knowledge that a crisis happened in other countries. The corresponding ACH specification is as follows. Construct an aggregate dummy variable $Crisis_t$, which takes the value of one if any country in the sample was attacked in month t. Based on $Crisis_t$, construct the aggregate duration variable d_t. Intuitively, d_t measures the frequency of speculative attacks happened in a group of countries. It is also shown in Table 4. The ACH (1,0) model with d_t is

$$P(Crisis_{i,t} = 1) = \frac{1}{C + \alpha d_{t-1}} \tag{5.3}$$

The result is reported in Column 5 of Table 3. The performance of model 5.3 is significantly better than model 5.2. The log likelihood is –45.37, a more than 15% improvement over model 5.2. The α parameter in model 5.2 is 1.57 and highly significant.

Table 4. Duration Variables for ACH Models

	Indonesia		Korea		Philippines		Thailand		Sample	
	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{t}$	d_{t}
Jan-94	0	1	0	61	0	22	0	40	0	22
Feb-94	0	1	0	61	0	22	0	40	0	22
Mar-94	0	1	0	61	0	22	0	40	0	22
Apr-94	1	87	0	61	0	22	0	40	1	19
May-94	0	87	0	61	0	22	0	40	0	19
Jun-94	0	87	0	61	0	22	0	40	0	19
Jul-94	0	87	0	61	0	22	0	40	0	19
Aug-94	0	87	0	61	0	22	0	40	0	19
Sep-94	0	87	0	61	0	22	0	40	0	19
Oct-94	0	87	0	61	0	22	0	40	0	19
Nov-94	0	87	0	61	0	22	0	40	0	19
Dec-94	0	87	0	61	0	22	0	40	0	19
Jan-95	0	87	0	61	0	22	0	40	0	19
Feb-95	0	87	0	61	0	22	0	40	0	19
Mar-95	0	87	0	61	0	22	0	40	0	19
Apr-95	0	87	0	61	0	22	0	40	0	19
May-95	0	87	0	61	0	22	0	40	0	19
Jun-95	0	87	0	61	0	22	0	40	0	19
Jul-95	0	87	0	61	0	22	0	40	0	19
Aug-95	0	87	1	187	0	22	1	129	1	16
Sep-95	0	87	0	187	0	22	0	129	0	16
Oct-95	0	87	0	187	0	22	0	129	0	16
Nov-95	0	87	0	187	0	22	0	129	0	16
Dec-95	0	87	0	187	0	22	0	129	0	16
Jan-96	0	87	0	187	0	22	0	129	0	16
Feb-96	0	87	0	187	0	22	0	129	0	16
Mar-96	0	87	0	187	0	22	0	129	0	16
Apr-96	0	87	0	187	0	22	0	129	0	16
May-96	0	87	0	187	0	22	0	129	0	16
Jun-96	0	87	1	10	0	22	0	129	1	10
Jul-96	0	87	0	10	0	22	0	129	0	10
Aug-96	0	87	0	10	0	22	0	129	0	10
Sep-96	0	87	0	10	0	22	0	129	0	10
Oct-96	0	87	0	10	0	22	0	129	0	10
Nov-96	0	87	0	10	0	22	0	129	0	10
Dec-96	0	87	0	10	0	22	0	129	0	10

Table 4. Duration Variables for ACH Models (Continued)

	Indonesia		Korea		Philippines		Thailand		Sample	
	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_{i,t}$	$d_{i,t}$	$Crisis_t$	d_t
Jan-97	0	87	0	10	0	22	0	129	0	10
Feb-97	0	87	0	10	0	22	0	129	0	10
Mar-97	0	87	1	9	0	22	0	129	1	9
Apr-97	0	87	0	9	0	22	0	129	0	9
May-97	0	87	0	9	0	22	1	21	1	2
Jun-97	0	87	0	9	0	22	0	21	0	2
Jul-97	1	39	0	9	1	58	1	2	1	2
Aug-97	1	1	0	9	0	58	1	1	1	1
Sep-97	0	1	0	9	1	2	0	1	1	1
Oct-97	1	2	1	7	0	2	0	1	1	1
Nov-97	0	2	1	1	0	2	1	3	1	1
Dec-97	1	2	1	1	1	3	1	1	1	1

A comparison between model 5.2 and 5.3 indicates that the contagion effect can explain the attacks better than the country specific duration dynamics. A more explicit way to look at it is to include both d_t and $d_{i,t}$ in the same model and see which one plays a more important role. We estimate model 5.4

$$P(Crisis_{i,t}) = \frac{1}{C + \alpha d_{i,t-1} + \delta d_{t-1}} \tag{5.4}$$

The result is reported in column 6 of Table 3. It confirms the implication from model 5.3 and 5.4. The estimate of the parameter for $d_{i,t}$ is zero (the ACH model confines the value of α to be nonnegative). This makes model 5.4 and 5.3 equivalent. The explanatory power of country specific duration dynamics is completely dominated by that of regional duration process. Another way to see the "value added" of the duration dynamics is to add the duration variable into model 5.1 and check the improvement in the log likelihood. The specification becomes

$$P(Crisis_{i,t} = 1) = \frac{1}{C + \alpha d_{t-1} + Macro1_{t-1}'\gamma} \tag{5.5}$$

where $Macro1_{t-1} = \begin{bmatrix} MRES_{i,t-1}' \\ MREG_{i,t-1}' \\ RER_{i,t-1}' \\ INF_{i,t-1}' \end{bmatrix}$, $\gamma = \begin{bmatrix} \gamma_1 \\ \gamma_2 \\ \gamma_3 \\ \gamma_4 \end{bmatrix}$, $MRES_{i,t-1} = \begin{bmatrix} MRES_{ind,t-1} \\ MRES_{kor,t-1} \\ MRES_{phi,t-1} \\ MRES_{tha,t-1} \end{bmatrix}'$. Other vectors in $Macro1_{t-1}$ are defined in the same way as $MRES_{i,t-1}$. The result is reported in column 7 of Table 3. This model gives a log likelihood of –44.77, which is about 10 higher than the ACH model in equation 5.1. This gain in the log likelihood comes from the inclusion of only one extra variable, d_t. On the other hand, compared to model 5.3, model 5.5 introduces four more variables, but the log likelihood only goes up by about 1. The advantage of the duration dynamics over the fundamentals is overwhelming.

In Figures 1 to 4, I plot crisis dummy and the fitted values from the probit model, and the model from equation 5.5. The improvement from the probit model to the ACH model is striking. The fitted values from the probit model are almost flat for the whole period. On the contrary, the ACH model clearly illustrates the increase of speculative pressure in the four countries. The hazard rates from the ACH model show two obvious jumps, one in June and the other in September. Looking into the constructed regional duration variable d_t in Table 4, one can see why the two jumps happen. In May 1997, Thailand's currency was attacked. As a consequence, the duration variable changed from 9 to 2, which resulted in the jump in hazard for June. The change in hazard for September comes from the attacks happened in August. The duration variable in August became 1, which indicates the intensification of the crisis.

How well did the models capture the months that the crisis dummy $Crisis_{i,t}$ is one? Table 5 shows the hazard rates from the two models plotted in the graphs. The probit model missed all the attacks. For all the 16 months identified by the dummy, the highest hazard from the probit model is 0.22. The ACH model gives a much better description of the attacks happened in the second half of 1997. This is especially clear for Indonesia, Korea and Thailand. For the Philippines, the hazard rates are not as impressive as those in the other three countries, but the ACH model still outperforms the probit model by a large margin.

The contagion effect may come from fundamentals of other countries. To test if this is the case, I put fundamentals from other countries into equation 5.5 and use likelihood ratio test to check if these variables are redundant. All the foreign fundamentals fail the tests of significance. The "fundamental based" contagion is not supported by the ACH model.

Figure 1

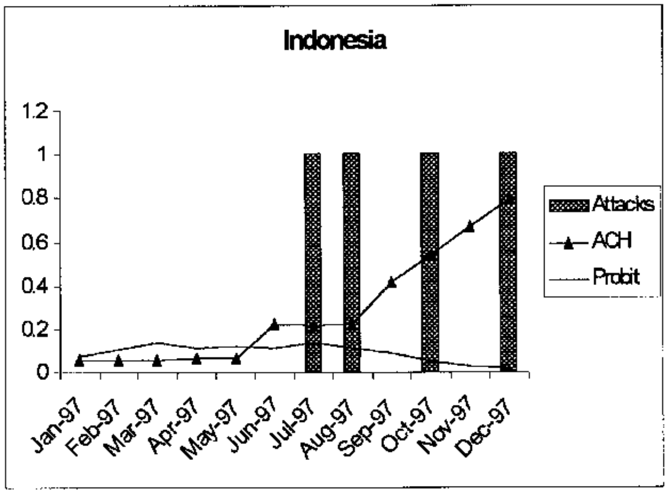

Figure 2

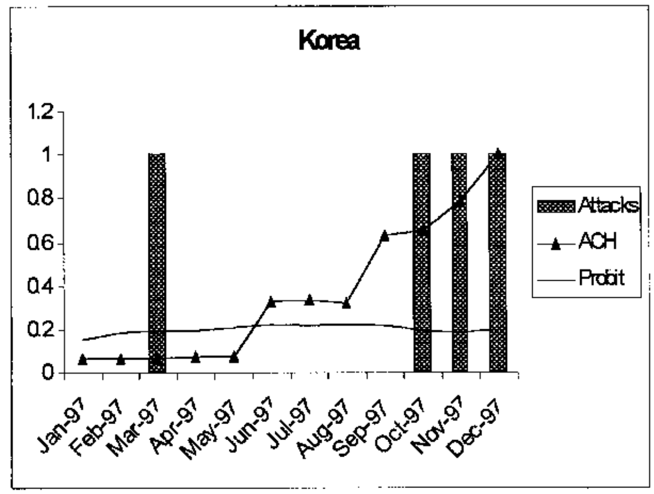

Figure 3

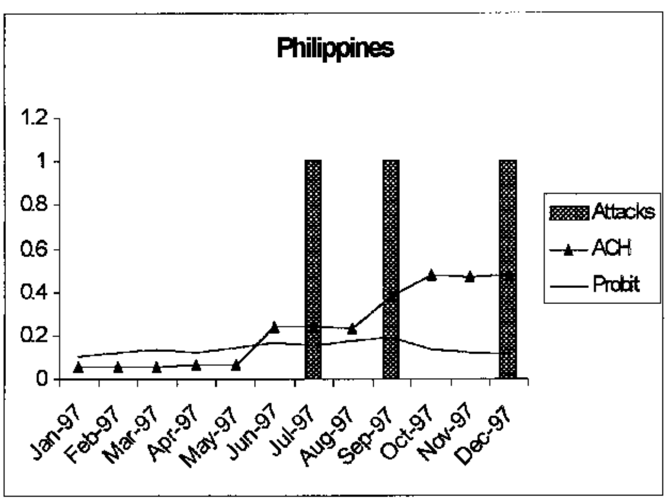

Figure 4

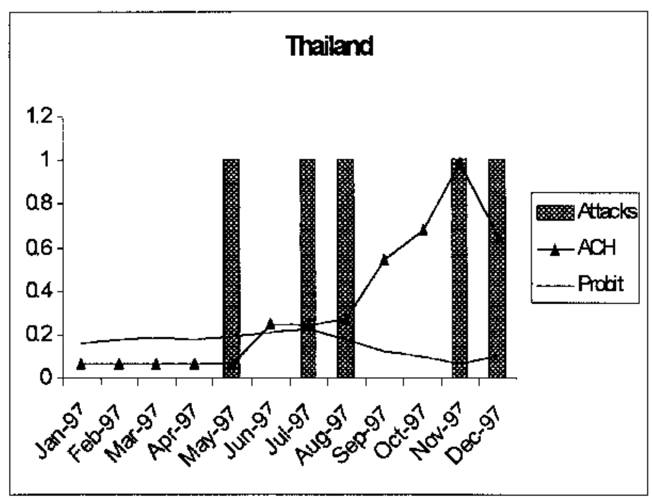

Table 5. Fitted Values from Three Models.

Attacks		Probit	ACH 1
Indonesia	Jul-97	0.13	0.21
	Aug-97	0.11	0.22
	Oct-97	0.05	0.53
	Dec-97	0.02	0.79
Korea	Mar-97	0.19	0.06
	Oct-97	0.19	0.64
	Nov-97	0.19	0.78
	Dec-97	0.19	0.99
Philippines	Jul-97	0.15	0.23
	Sep-97	0.19	0.38
	Dec-97	0.11	0.48
Thailand	May-97	0.19	0.06
	Jul-97	0.22	0.24
	Aug-97	0.17	0.27
	Nov-97	0.07	0.98
	Dec-97	0.10	0.64

1/ This table shows fitted values only for months that the crisis dummy is 1.

VI. Conclusion

This paper applies the ACH model to test for contagion effects in the Asian crisis in 1997. The result strongly supports the hypothesis that the probability of one currency being attacked in one period is influenced by the frequency of speculative attacks in other countries before that period. Why the attacks happened in the very beginning is still a puzzle, but the regional duration dynamics seem to explain the evolution of the crisis in Indonesia, Korea, and Thailand quite well.

Economists and policy makers have been interested in forecasting currency crisis for a long time. Research on this topic has focused on long run forecast (Kaminsy, Lizondo, and Reinhart, 1998). This paper shows that the ACH specification could be a powerful tool for them. Since the ACH model only uses lag variables on the right hand side, it can be used for forecasting. The out of sample performance of the ACH model is not tested in this paper. This is left for future research.

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