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Latin America and East Asia in the context of an insurance model of currency crises

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Abstract

This paper focuses on the 1995 Latin American and the 1997 East Asian crises using an insurance-based model of financial crises. First, the Dooley model is described. Second, some empirical evidence for an insurance model is presented. The key variables in this approach include the ratio of foreign exchange reserves to bank loans (domestic credit) extended to the private sector, the ability of the private sector to appropriate government assets, and appropriation as measured by capital flight. We argue that the insurance model is consistent with the observed evolution of these variables in the recent crises in Latin America and Asia. Finally, we examine the statistical evidence in favor of the model using panel regressions. We find that the econometric results are consistent with the insurance model, and tend to support this approach over some competing explanations. © 1999 Elsevier Science Ltd. All rights reserved.

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1. Overview

A comparison of the currency crises of East Asia and Latin America presumes that there is something different about East Asia that makes it a natural aggregate for analysis. Until quite recently the distinguishing characteristics relative to other regions have been strong administrative controls of domestic and international finan-

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cial markets, heavily managed exchange rates, rapid growth of international trade and an admirable degree of financial stability. However, the devaluation of the Thai baht and the related attacks on other East Asian currencies have reinforced the warnings from economists that there is nothing inherently special about international finance in East Asia. Indeed, Kaminsky and Reinhart (1998b) have argued that whatever differences may have existed in the past between these two regions, they are fast disappearing. Hence, policy conflicts associated with the implicit guarantees associated with a managed nominal exchange rate and other monetary and fiscal objectives are very likely to lead to speculative attacks whether in Korea or in Mexico.

Moreover, at least two new sets of reasons to worry about speculative attacks have emerged since 1992. Spectacular attacks on the European Exchange Rate Mechanism have suggested that the usual policy conflicts might not be necessary to trigger an attack. One idea is that financial policies of the government might make the regime vulnerable to "self-fulfilling" shifts in private expectations (e.g. Eichengreen and Wyplosz, 1993; Obstfeld, 1994). The other important development is the realization that international capital inflows might set the stage for attacks on recently liberalized domestic financial markets.

For emerging markets in East Asia these new reasons to worry about speculative attacks are particularly relevant. As discussed below, both theories suggest that a country can get all the usual "IMF fundamentals" right and still see its monetary arrangements destroyed by a successful speculative attack. The reasons to avoid such situations are clearly illustrated in recent revised projections for growth rates in countries that have suffered through recent attacks.

In Section 2 we lay out the intuition for the insurance model developed in Dooley (1997). This is a "first generation" model of speculative attacks in which the policy conflict is between a credit-constrained government's desire to accumulate liquid assets and its desire to insure domestic financial systems. Section 3 interprets the behavior of key variables in the context of this model—such as the size of the insurance pool and extent of capital inflows, the duration of the inflow and degree to which the private sector is able to appropriate government assets, and the observable manifestations of the gap between the government's state contingent assets and liabilities. Unfortunately, all these variables are difficult to measure directly. In the latter case, actual contingent assets and liabilities are not observable, so we take as proxies bank lending to the private sector, under the presumption that the government cannot afford to allow the banking system to collapse. Liabilities are at a first approximation equal to foreign exchange reserves. This approach also implies that one sees the first manifestation of an incipient currency crisis not in the usual macroeconomic observables (interest rates, exchange rate overvaluation, etc.), but in "capital flight".

In Section 4, we present a formal econometric examination of the data, focusing on five Latin American and six East Asian countries over the sample period 1980–1997. We find that there is substantial evidence consistent with our model. We also subject the empirical model to some robustness checks to see whether the results are sensitive to the inclusion of variables that other competing models imply should determine financial crises. The conclusions are presented in Section 5.

2. An insurance model

The argument is close to the spirit of that offered by Diaz-Alejandro (1985) and developed further by Velasco (1987). The policy conflict in the model to be tested below is between the desire of a credit-constrained government to hold reserve assets as a form of self-insurance and the government's desire to insure financial liabilities of residents. The first objective is met by the accumulation of foreign exchange reserves and lines of credit. The second objective generates incentives for investors to acquire the government's liquid assets when yield differentials make this optimal.

These ingredients provide a plausible capital inflow/crisis sequence. An important feature of our version of the model is that the capital inflow does not simply contribute to the vulnerability of the regime. Because the government is credit constrained, it cannot borrow against future tax receipts in order to delay a crisis. In this environment, credible free insurance raises the market yield on a set of liabilities issued by residents for a predictable time period. Yields rise because residents compete in order to exploit the insurance. This resulting yield differential between insured domestic "deposits" and the international risk free rate generates a private gross capital inflow (a sale of domestic liabilities to nonresidents) that continues until the day of attack. The private inflow is necessarily associated with some combination of an increase in the government's international reserve assets, a current account deficit and a gross private capital outflow. When the government's reserves are exactly matched by its contingent insurance liabilities, the expected yield on domestic liabilities falls below market rates and investors sell the insured assets to the government, exhausting its reserves. The speculative attack is fully anticipated and at the time of the attack nothing special happens to the fundamentals or expectations about the fundamentals.

This sequence of events is illustrated in Fig. 1.[1] The positive vertical axis in the top panel measures the stock of assets that the government, including the central bank, could liquidate *during a crisis* in order to redeem liabilities to the private sector. The negative vertical axis measures the government's total stock of contingent and noncontingent liabilities. We start from a situation in which the value of assets, A_0, is growing but is less than the value of debt, L_0.[2] A fall in international interest rates at t_1 reduces the value of the government's long-term liabilities from L_0 to L_1, but does not affect the contractual value of short-term assets. Part of the government's assets can now support additional liabilities.

In the middle panel we show the stock of insured private liabilities. At t_1, residents that can issue an insured liability will now offer to do so in order to appropriate some share of the proceeds, s.[3] Sellers of such liabilities are residents simply because only residents' liabilities are eligible for insurance. The government's contingent

[1] The dynamics of the model are set out more carefully in Dooley (1997).

[2] The market value of the debt would be equal to the collateral value. That is, there would be a secondary market price discount. See Dooley et al. (1996) for a model and evidence.

[3] A more realistic form of appropriation is state contingent. That is, insured residents exploit insurance by reaching for risk. They share returns earned in good states of the world and default in bad states of the world.

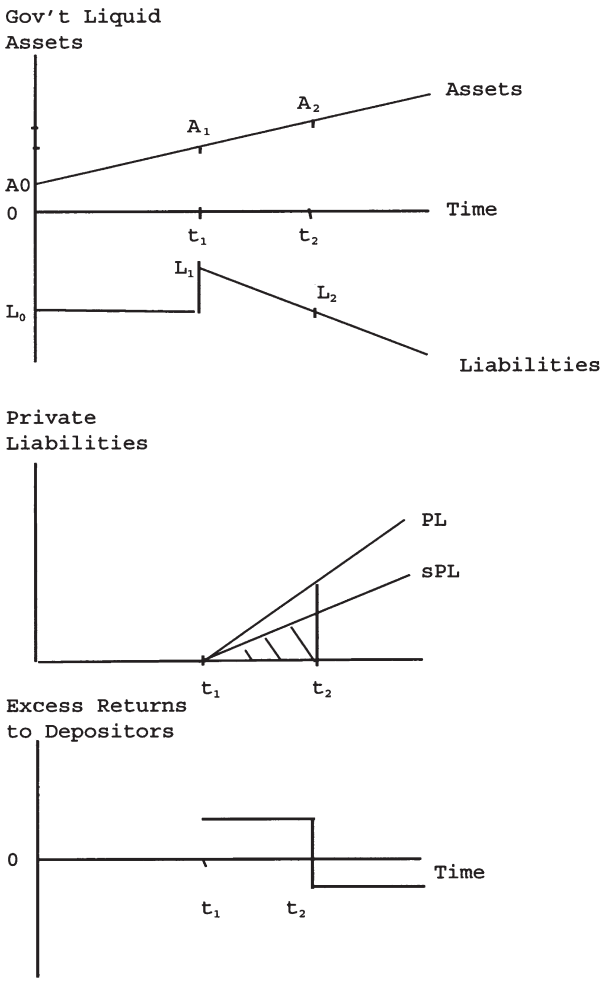

Fig. 1. Government guarantees on foreign borrowing.

liability is the same fraction of new insured liabilities (the shaded area in the middle panel).

The value of *s* is specific to the country and is small in a well regulated market and large in a poorly regulated market. The time derivative of the flow of new issues (the slope of *PL*) is also specific to each country and is also a function of the supervisory system in place. Relatively poorly regulated financial markets will see a relatively rapid increase in insured liabilities.[4]

[4] In the diagram it is assumed that implicit liabilities grow more rapidly than reserves. This does not follow from theory. In fact, an important difference between emerging markets and industrial countries is that the governments in industrial countries constrain domestic intermediaries before the government's net worth is exhausted. Thus, in the US savings and loan crisis there was no run on the government's

Investors are willing to buy residents' liabilities because they are insured and because competition among (resident) sellers will force them to share a part of their appropriation with (nonresident) creditors. This will take the form of above-market expected yields on residents' liabilities.[5] Yields will be the same for both domestic currency and foreign currency liabilities of residents as long as the insurance is expected to cover both types of domestic liabilities.[6]

As long as the "foreign" investors earn above-market yields there is a disincentive for an attack on the government's assets. Investors will prefer to hold the growing stock of high yield insured liabilities of residents and allow the government to hold reserves that earn the risk free rate. Private profits are realized before the attack. The attack itself is generated by competition to avoid losses. When the contingent liabilities of the government are just equal to liquid assets ($A_2 = L_2$), competition among investors will ensure that all will call the insurance option. The bottom panel of Fig. 1 reflects the assumption that nonresidents demand a constant premium in order to accumulate insured deposits. On the day of the attack the expected value of this premium becomes negative because each depositor's share of the insurance pool will begin to shrink. Resident borrowers will continue to appropriate a part of new loans and this will depress expected yields on deposits that after t_1 are only partially insured.

Following an attack, the regime returns to its initial equilibrium in which the government's net international reserves have returned to zero. The crisis does nothing to resolve the underlying policy conflict. Following the crisis, the government will once again attempt to accumulate liquid assets and unexpected capital gains and losses on the governments asset position will eventually generate a new inflow/attack sequence.

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reserves. Instead a binding constraint was established by re-regulating the financial system. The resulting loss to the government was substantial but well within its ability to provide credible insurance.

[5] The accounting is straightforward if we abstract from financial intermediation. Suppose a resident household can issue a $10 liability to a foreign investor. The household plans on repaying $5. The household shares its gain by paying the investor $2.50 and keeping $2.50. The investor expects the government to purchase the liability for $10 in 1 year. The government's contingent liability is $5.00. More realistic examples will involve one or more financial intermediaries in this process. The distribution of the rents among the participants will depend on their relative bargaining power. If investors' demand for claims on residents are very elastic, residents will capture most of the rents. This seems to us to be the most likely outcome. It is difficult to interpret historical evidence for deposit rates. As insurance became credible after 1989, deposit rates should have fallen as default risk was absorbed by the government. In Mexico real ex-post rates on domestic deposits (adjusted for actual changes in dollar exchange rates) fell from about 15% above US rates in 1990 to equality with US rates in late 1994. While this pattern in returns is consistent with our model, Mexico's stabilization program may have had important implications for this history of yield differentials. See Kaminsky and Leiderman (1996) for a discussion of stabilization plans and real interest rates.

[6] If the insurance is only available on domestic (foreign currency) liabilities an equilibrium covered interest differential will emerge in favor of domestic (foreign currency) liabilities. A fixed exchange rate regime is not crucial for the argument. Under floating exchange rates the nonresident investor plans to liquidate her position at the time of the anticipated attack. It follows that any spot foreign exchange transactions will be offset by a matching forward exchange transaction. Private interest arbitrage will ensure that there is no net change in spot or forward rates.

3. Empirical implications

3.1. Examples of changes in binding constraints

Three "insurance fundamentals" must be present in order to generate a private capital inflow followed by a speculative attack. The first is that a credit-constrained government must have positive net assets. Net assets are defined to include some contingent assets and liabilities but not the present value of future tax receipts. Second, the government's commitment to exhaust these net reserves to pay off an implicit or explicit insurance contract must be credible. That is, it must be consistent with the government's incentives and ability to mobilize and exhaust a well-defined set of assets *after* the attack begins. Third, private investors must have access to transactions that produce insured losses.

All three factors must be present to trigger a capital inflow and subsequent attack. One or more of these fundamentals are found in most countries most of the time, but as long as one ingredient is missing there will be no capital inflow and no crisis. Crisis episodes are associated with the relaxation of a *binding* constraint. It follows that there is no simple temporal ordering of changes in insurance fundamentals and crises.

A government with open financial markets, weak regulatory systems and a credible commitment to insure a well-defined set of residents' liabilities will not experience a sequence of capital inflows followed by an insurance attack unless it has net assets to expend during the attack. For middle income developing countries with substantial stocks of external debt the missing fundamental from 1982 through to 1989 was a stock of assets to support a credible insurance commitment. For this group of countries an important source of changes in the value of governments' net assets was changes in the market value of governments' external debt caused by changes in international interest rates.

Table 1 shows data for individual emerging markets that accounted for about 80% of the total capital inflow to emerging markets from 1990 to 1996. Column 1 shows the cumulative net private capital inflow to each country over the 7 year time period. Our hypothesis is that all these capital inflows were generated by credible insurance policies. The capital gain on external debt outstanding in 1989 was the product of outstanding debt and the change in the relevant interest rate on that debt. Because the currency denomination of the Latin American and Asian debt was quite different in 1990 we constructed a weighted average international interest rate for each debt stock (see Fig. 2). The sensitivity of the market value of Latin American debt is clear because it traded at substantial discounts before the drop in interest rates. The secondary market price for Latin American debt, also shown in Fig. 2, jumped from about 30 cents to near par as interest rates declined. There is no similar change in the market value of Asian debt because its market price was near par in 1990. Our interpretation of this data is that Latin American governments could not have had net assets to cover new insurance as long as existing government liabilities sold for substantially below par. The fall in international interest rates eliminated the claims of existing creditors in excess of government assets in 1990. The capital gain in

Table 1
Capital flows and assets for selected emerging markets (billions of US $)

	Private inflows 1990–1996[a] [1]	Debt 1989[b] [2]	Change in reserves 1990–1996[c] [3]	Rescue package [4]
China	217.2	44.9	77.5	
Mexico	112.5	95.6	15.2	47
Korea	79	33.1	19.2	52.8
Brazil	76	111.3	50.9	
Malaysia	60.1	18.6	17.3	
Indonesia	60.2	53.1	10.8	42.3
Thailand	47.8	23.5	24.4	17.2
Argentina	46.8	64.7	13.5	
India	27.8	62.5	18.6	
Russia	59.9	79	5.4	22.6
Turkey	23.2	41.6	10.4	
Chile	20.2	18.2	8.8	
Hungary	19.7	20.6	8.7	

[a] Russia and Indonesia private inflows cover 1990–1997.
[b] Russia's debt is from 1992.
[c] Russia's change in reserves is measured from 1993 to 1996; Mexico's change in reserves is measured from 1990 to 1993.
Sources: Debt—World Debt Tables from 1990–1991; Private inflows—Global Development Finance, except Korea—International Financial Statistics (IFS); Change in reserves—IFS; Rescue package—IMF.

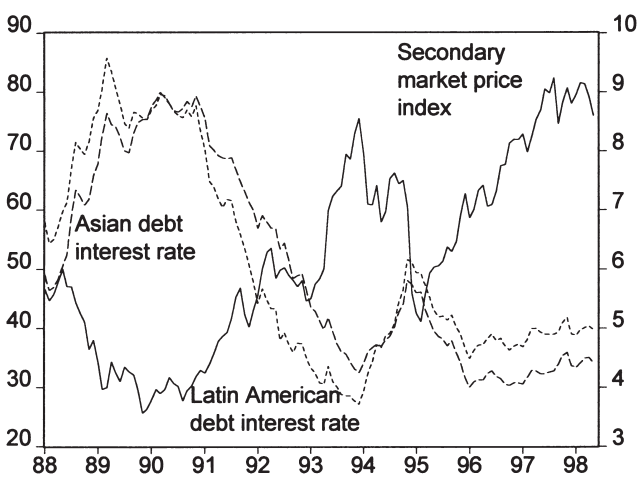

Fig. 2. Secondary market prices for Latin American debt and interest rates for Latin American and Asian debt.

1990 was about one half of the initial stock of floating rate external debt shown in column 2 of Table 1. From this point forward capital gains on debt and other asset accumulation provided a credible insurance pool. A similar capital gain for Asian governments created an immediate insurance pool.

Table 1 also provides evidence that liquid assets were accumulated. The change in international reserve assets and official rescue packages provide rough measures of assets available to support an inflow/crisis sequence. Column 3 shows the cumulative change in international reserve assets over the same time period. Reserves were augmented by contingent lines of credit from other governments and international organizations. Column 4 shows that a large share of the resources made available to private investors following the Mexican crisis, about $47 billion, came from loans from creditor governments and international organizations (Boughton, 1997). Official credits to Indonesia, Korea and Thailand following the Asian crisis totalled about $118 billion and in 1998 Russia received an additional $22.6 billion in official credit (IMF, 1998). While only suggestive, this data is consistent with the view that, except in the transition economies, capital gains on debt were an important contribution to net assets as international interest rates fell after 1989. Moreover, for all the emerging economies, the sum of capital gains on debt and the subsequent accumulation of international reserves and credit lines were of roughly the same magnitude as private capital inflows. Fig. 3 confirms the assertion that the actual inflows and the predicted inflows match almost one-for-one. An OLS regression of cumulative net inflows on the insurance pool figures implied by Table 1 yields a slope coefficient of 0.97, with a standard error of 0.04, and an R^2 of 0.92.

Another plausible sequence of events that would trigger an inflow/crisis sequence is economic reform in developing and transition economies. Reform involves both opening of domestic financial markets and improved access to international financial markets. These programs relax three constraints. First, they make domestic liabilities

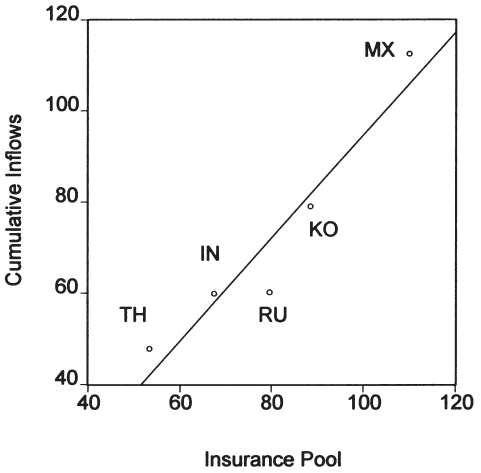

Fig. 3. Cumulative inflows versus implied insurance pool. *Source*: Table 1 and author's calculations. Indonesia (IN), Korea (KO), Mexico (MX), Russia (RU) and Thailand (TH).

available to foreign investors. Second, they make the existing regulatory framework less effective. These effects of liberalization are discussed and modeled in McKinnon and Pill (1996) and Krugman (1998). Kaminsky and Reinhart (forthcoming) offer empirical evidence that liberalization helps to predict banking crises. Our approach suggests that the third condition, positive net assets, is also an important constraint. Liberalization provides an insurance pool because, as noted above, creditor governments and international organizations provide generous lines of credit to support reform programs.

3.2. Bank credit, lending booms and reserves

It might appear that this model reduces, in its empirical manifestation, to the familiar Krugman (1979) speculative attack model, with slightly different definitions for assets and liabilities. In fact, the redefinition of liabilities and assets to include contingent ones is of key importance. The central variable in the Krugman model is the reserves to narrow money (M_1) ratio. Mexico does appear to conform to the suggested pattern, although it is interesting that the ratio does not decline until the end of the first quarter of 1994. Brazil and Argentina clearly do not fit the pattern (the actual series are displayed in the working paper version of this article).

Thailand and Malaysia do display a declining reserves to M_1 ratio for the year preceding the crisis. Indonesia, in contrast, exhibits rising ratios leading up to the crisis. The Korean reserves to money ratio peaks at the end of the second quarter of 1997, just a few months before the won devaluation.

In order to examine the implications of the insurance model, we redefine the contingent liabilities to equal the bank loans extended to the private sector (i.e. domestic credit). This implies that the government is generally unwilling to allow the banking system to collapse, and either explicitly or implicitly guarantees these deposits.[7]

One indicator of the fragility of the banking sector (and the proportion of bad loans) is the rate of growth of domestic credit. Various researchers (Kaminsky and Reinhart, 1998a, forthcoming) have shown a correlation between banking crises and currency crises. Kaminsky and Reinhart also show that rapid growth of domestic credit lagged for 2 years is a good predictor of a financial crisis. Rapid domestic credit growth also finds a role in various post-mortem accounts (e.g. BIS, 1998, Chapter VII). Chinn and Dooley (1997) find some evidence that rapid expansion of bank lending increases the riskiness of the marginal project in some Pacific Rim countries.

The model predicts that the reserves to domestic credit variable (*RS__DCR*) should be falling in the period leading up to a crisis, while the 2 year change in the log

[7] One criticism of this approach might be that one should use the ratio of nonperforming loans to adjust the size of contingent liabilities. However, such data are administrative in nature, and are probably uninformative regarding the true extent of banking sector problems. In fact, to the extent that regulators may hesitate to declare loans nonperforming for fear of forcing bankruptcy, the nonperforming loan ratio may exhibit perverse behavior. Corsetti et al. (1998) report positive results using a considerably modified measure of nonperforming loans.

real domestic credit variable, *LGBOOM*, should peak around 2 years prior to the crisis. Mexico approximately fits this pattern (4 years would be closer), as does Argentina. Brazil's ratio does not fit the pattern, although the lending boom variable does climb steeply 2 years prior to the Tequila Effect (the real stabilization plan in June of 1994 may complicate the interpretation of these data).

For Thailand, reserves to money declines in the first quarter of 1996 onward. The peak in the domestic credit growth is exactly 2 years before the crisis. Malaysia and, to a lesser extent, Singapore also fit this pattern. For the former the reserves to money ratio is declining over the entire period from 1994Q1 to 1997Q3, and for the latter, from 1994Q1 to 1996Q2, and stabilizing thereafter.[8]

Indonesia presents an interesting case. Bank lending exploded in the early 1990s, growing at an annualized rate of 38% per year, then plummeting to 5% in 1993 before accelerating again in 1993, when bank lending growth rose again. *RS__DCR* also peaks exactly 2 years before the crisis. Korea's *RS__DCR* ratio declines from its peak in 1996Q1, and then plummets again beginning in 1997Q2 as the other East Asian currencies fall. While the bank lending growth does not peak 2 years before the won crisis, it does decelerate.

The end result of these rapid expansions in bank lending is a large nonperforming loan problem in many of these countries. Indonesia, Malaysia and Thailand face serious property sector risks. Korea, on the other hand, faces a very serious corporate sector risk, in the form of nonperforming loans extended to the chaebols (Morgan Guaranty, 1998).

3.3. Property booms and capital flight

As noted above a wide variety of financial transactions might generate implicit liabilities for the government. Depending on the regulatory environment, financial institutions engaged in looting (Akerlof and Romer, 1993) will seek out transactions that are least likely to attract attention from the authorities. For example, our interpretation of the prevalence of real estate lending leading up to crises is that loans based on this type of collateral are generally favored by regulatory authorities. Given perfect foresight, the run up in property values preceding the crisis reflects property owners bargaining position in the game. Since property owners know that looting requires their cooperation in borrowing against property, a rising price for their property is necessary to compensate them. Notice that in this context the property owner is selling the property to the bank since both parties know that the crisis is coming and that the collateral will be forfeited. In fact, competition among property owners will ensure that all the property will be owned by the banks when the

[8] Galindo and Maloney (1998), drawing on Calvo and Mendoza (1996a, b) and Krugman models, find that the reserves/M_2 ratio predicts speculative pressure well for their sample, except for the East Asian countries (their sample did not include the 1997 crises, however). This reserves/M_2 ratio behaves similarly to the *RS__DCR* ratio in our sample, although the latter tends to fall more substantially prior to a crisis, in our dataset.

crisis occurs. The same argument explains the run up in the value of equities and other assets that typically serve as collateral for bank credit.

All the private participants in this game will look forward to conditions following the crisis. It seems quite likely that assets that are not insured might be vulnerable to taxation in order to offset the government's loss. Thus, while private capital inflows are observed residents will also export private capital in order to avoid post-crisis taxation. The empirical counterpart to this is unrecorded increases in gross private claims on nonresidents.

The fact that residents are trying to hide these assets from the domestic authorities makes measurement of capital flight difficult but a number of statistical procedures have proven useful. In this paper we use a residual method that exploits the country's balance of payments data but augments this with data for international lending to the country reported by other countries.[9] Fig. 4 presents comparative data for Latin America and East Asia over the 1978–1994 period. The striking feature of this data is that capital flight was not an important factor in Asia before or following the 1982 debt crisis. Our interpretation is that an insurance crisis in Latin America was not present in Asia in this time period. This, we think, accounts for the Asian emerging markets being little effected by the 1982 crisis. In contrast, capital flight was clearly a problem in Asia after 1993. In Fig. 5, estimates for capital flight (expressed as a proportion of GDP) are presented for the three largest Latin American countries over the 1990–1996 period. A positive number indicates "capital flight". Fig. 6 shows that there was substantial capital flight from East Asia in 1993–1995. Indonesia exhibits the most variable and substantial magnitudes of capital flight over the sample period. Although the peak capital flight for Indonesia is 10.8% of GDP in 1994, the figure for 1997 is still a considerable 2% of GDP. Korean capital flight is also substantial. Between 1993 and 1996, capital flight ranged between 2 and 3% of GDP

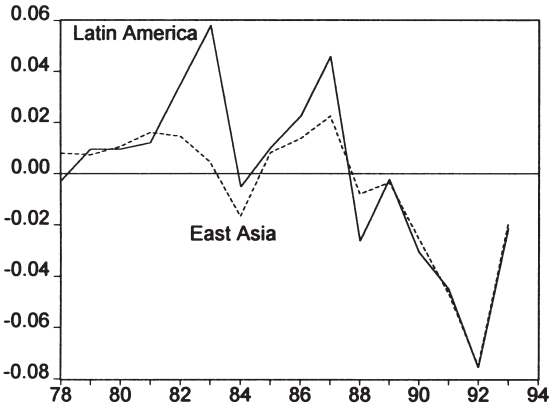

Fig. 4. Latin America and East Asia capital flight.

[9] These figures are the "adjusted World Bank" capital flight numbers. See Dooley (1988) for a discussion.

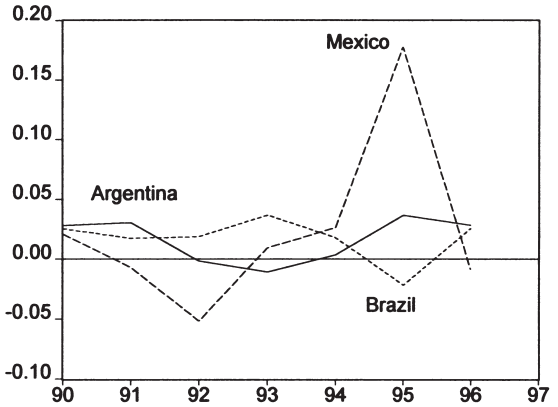

Fig. 5. Capital flight to GDP ratio, Argentina, Brazil and Mexico.

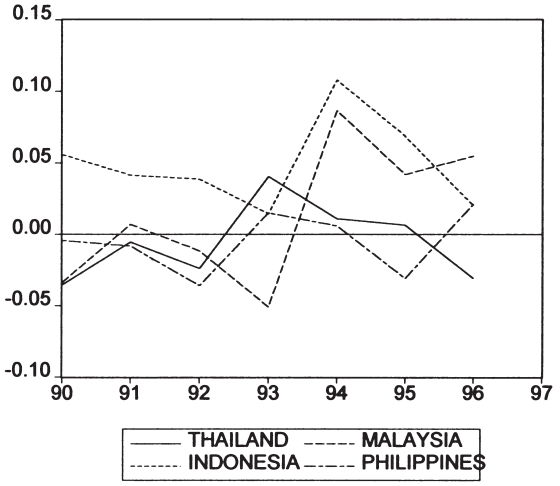

Fig. 6. Capital flight to GDP ratio, Thailand, Malaysia, Indonesia and the Philippines.

(Fig. 7). This suggests that residents saw trouble coming and moved their assets beyond the reach of the domestic government. It is simply implausible that the gross capital inflow over these same years was not largely motivated by the expectation that the government would back some set of insured positions when the crisis occurred.

3.4. Duration of capital inflows

Although a capital inflow/crisis sequence is likely to begin at about the same time for many indebted developing countries following a decline in international interest rates, the duration of the inflow and the timing of the expected crisis can vary widely. In fact, a crisis might never occur if the government reacts properly. The duration

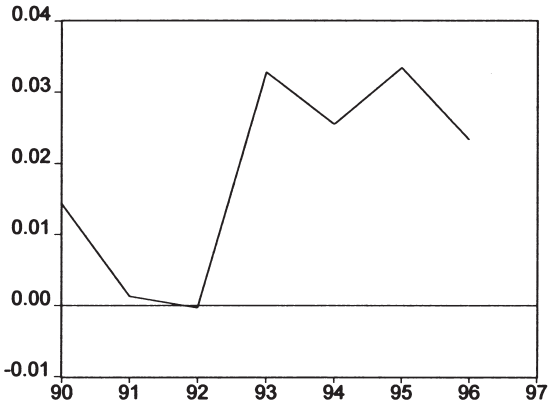

Fig. 7. Capital flight to GDP ratio, Korea.

of the capital inflow will depend on the rate at which banks, households and firms can sell insured liabilities and on the profitability of appropriation. If the share of each deposit appropriated is low because of regulatory constraints, appropriation may become unprofitable before the governments net assets are exhausted. The important implication is that crises will be spread over time and move from poorly to well-regulated financial systems.

A common negative shock to governments' net assets could truncate this process and generate a number of crises at the same time. In this case, a common fundamental has changed and crises are bunched in time for this reason not because events in one country alter expectations about events in others. As shown in Fig. 2, international interest rates did rise just before the Mexican crisis in 1994 and this common shock may explain the so-called Tequila Effect.

Crises might also be bunched in time because of revisions in expected values of official lines of credit. When an attack occurs, investors receive new information about the size and likely distribution of official credits available to cover insurance commitments. If the expected overall size of official lending is revised downward this can generate coincident runs in many countries. Following crises in Mexico and Russia, for example, there were considerable differences of opinion concerning the willingness or ability of the US government and international organizations to support additional loan programs in the face of congressional opposition. This may have reduced the expected value of official credits to other countries, perhaps to levels that made immediate attacks optimal. Even if the expected pool of loans is unchanged, the observation of loans to an individual country provides additional information about the expected distribution of loans over countries. On average, expectations for half of the countries will be revised downward and some of these might be pushed over the attack threshold. Ignoring the issue of interest rate increases, it still might be useful to consider the simple relationship between the duration of the capital inflow and some measure of how willing the government is to regulate financial markets. The model argues that ceteris paribus, a more transparent regulatory and financial

system will tend to extend the period of time before a crisis occurs. In Fig. 8, we assume that either the decline in US interest rates in 1990 or liberalization of the capital account triggers the beginning of capital inflows. *Duration* is the number of quarters from the beginning of inflows or liberalization to the crisis;[10] *transparency* is an *inverse* measure of corruption in 1996, described further below in Section 4. According to a truncated regression estimation procedure, there is a positive relationship between the (log) inverse of corruption and the duration, which is statistically significant.[11] A 10% decrease in corruption yields a point estimate implying a 2.5 quarter increase in the duration of the inflows (with the ± 2 standard error bands ranging from 2.0 to 2.9).

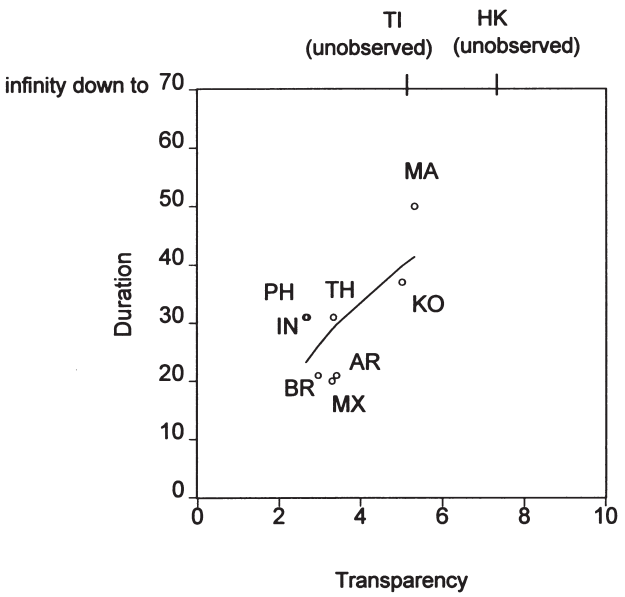

Fig. 8. Duration of inflows versus inverse of corruption. Regression line is for semilog specification. *Note*: *Duration* is the number of quarters from the beginning of inflows or liberalization. *Transparency* is an inverse measure of corruption. Argentina (AR), Brazil (BR), Mexico (MX), Indonesia (IN), Korea (KO), Malaysia (MA), the Philippines (PH), Thailand (TH), Taiwan (TI) and Hong Kong (HK).

[10] For most countries, the beginning of inflows is dated at 1990Q1, as US real interest rates began to fall. For Korea and Taiwan, opening is dated at 1989Q1, as suggested by Chinn and Maloney (1998), while opening for Singapore and Malaysia is dated by Chinn and Frankel (1994) at 1987Q1 and 1985Q2, respectively. The former two estimates are based on inverted quasi-money demand curves, while the latter two are based on covered interest differentials.

[11] Obviously, since some countries do not experience a crisis, OLS is not appropriate. We estimated a truncated regression in a semi-log specification; the slope coefficient is positive and statistically significant with or without inclusion of a constant (the latter is consistent with the view that the attack takes place instantaneously if a regime is completely unable to prevent the private sector from appropriating government assets). Since the constant is not statistically significant, we report in the text the results without the constant.

4. Panel regression analysis

4.1. Data and model specification

In this section we present some formal econometric results which bear upon the issue of which models describe the onset of crises in these countries. We do not attempt to replicate the comprehensive cross-country analyses, such as Frankel and Rose (1996), Kaminsky et al. (1998), Sachs et al. (1996), Corsetti et al. (1998) or Berg and Patillo (1998). Rather we focus on 11 countries that account for a large proportion of total capital flows to emerging markets during the 1980s and 1990s. Moreover, we attempt to more closely locate the timing of the crises and hence distinguish among competing hypotheses. The analysis is conducted on quarterly data for the period 1980Q1–1997Q4 for the Latin American countries of Argentina, Brazil, Chile, Colombia and Mexico, and the East Asian countries of Indonesia, Korea, Malaysia, the Philippines, Singapore and Thailand. Most of the data are drawn from the IMF's *International Financial Statistics*, June 1998 CD-ROM.

We use several measures of crises. The first is a binary variable defined using a threshold of a 20% quarterly change in the log bilateral real exchange rate, *CRISISR*. The second variable is defined using a weighted average of the log first differences of the real exchange rate and of international reserves, using a 20% cutoff (*CRISI2R*). A weight of three quarters is placed on the change in the bilateral exchange rate, and one quarter on the change in reserves. The third variable, *CRISIS_P*, is the same as *CRISI2R*, except that the cutoff value is 12.5%, and it defines a crisis period as the period in which the threshold is breached, plus the three subsequent quarters.

For the determinants of crises, we use a number of variables mentioned in the graphical assessment of the model. Ideally, we would like to directly measure the growth of governments' implicit liabilities. This is possible following a crisis since transfers to financial markets measure accumulated appropriation by the private sector. Such figures are reported by Corsetti et al. (1998) for 1 year—1996. However, it is impossible to obtain these variables for long time series as we require them. For the key variables we use the ratio of foreign exchange reserves to domestic credit extended to the private sector as a measure of the gap between contingent assets and liabilities (*RS_DCR*). As discussed above, the share of domestic credit that is the government's contingent liability is related to three variables that we can observe. The first is the *change* in log real domestic credit (*LDCD90*) over a 2 year period, *LGBOOM*; in the regressions, this variable is lagged by 2 years. The second is the volume of capital flight. The dollar amount of capital flight is measured using the World Bank approach. That is, capital flight is the sum of the current account surplus and increases in external debt, less recorded net private capital inflows and increases in official reserve assets. This variable is converted into domestic currency terms and expressed as a proportion of GDP, *CFLT*.[12] The third is the quality of

[12] Since the debt figures are available only at an annual frequency, we have generated quarterly series by using a HP filter.

regulation in domestic financial markets. We do not have direct observations on this variable, but we can proxy it with indices of corruption. The index we use is that of *Transparency International*, *TI*. *TI* reports corruption perceptions indices ranging in value from 10 (highly clean) to 0 (highly corrupt). The logged 1996 value of this index (*LTRANSPRNT96*) is used.[13] The US real interest rate, R^{US}, is also included as changes in this rate will revalue the external debt of these countries.

The regressions are estimated using probit,[14] in the following specification:

$$CRISIS_t = f(RS\_\_DCR_{t-1}, LGBOOM_{t-8}, R_{t-1}^{US}, TRANSPRNT96, CFLT_{t-8}; Z)$$

where Z is a set of other variables that are included with, or substitute for, the key variables implied by our model and can be thought of as controls or robustness checks. The set of variables for Z include the trade balance to GDP ratio, *TB__Y*, and the multilateral real exchange rate deviation from a linear trend, *DEV1* (Dornbusch et al., 1995). For purposes of comparison, we include a number of other indicators that are suggested by other models, such as the reserves to M_1 ratio, *RS__MRT* (Krugman, 1979) and reserves to M_2 ratio, *RS__M2R* (Calvo and Mendoza, 1996b). Note that no fixed effects are included in the panel regressions, so all countries are treated as identical. Obviously, a much better fit could be obtained merely by including country effects.

4.2. Empirical results

The results are reported in Tables 2–4. Table 2 contains the regression results using the *CRISISR* dependent variable. Since the absolute magnitudes of the probit regression coefficients have no simple economic interpretation, we only indicate the statistical significance of the coefficients. (Note that almost all the statistically significant coefficients are of correct sign; only in Table 4 do some incorrectly signed coefficients have some statistical significance, and then only at the 20% marginal significance level.) A baseline regression specification including only reserves to domestic credit, lagged lending boom, and the US real interest rate (column 1) indi-

[13] The *TI* Corruption Perceptions Indices are based on survey data from Economist Intelligence Unit, Gallup International, Institute for Management Development, Political and Economic Risk Consultancy, Political Risk Services, World Development Report and World Economic Forum. Details of construction of the indices are reported in Transparency International (1998). While period averages are given for the 1980–1985 and 1988–1992 periods, and a value for both 1996 and 1998, the data are not really comparable over time. In some regressions incorporating time varying values of the inverse corruption index, the coefficient is usually insignificant.

[14] In some other studies, such as Sachs et al. (1996) and Corsetti et al. (1998), continuous indicator variables have been used as the regressand. We also estimated OLS regressions incorporating the underlying data in *CRISI2R*. The results are weaker, with only the R^{US} and *DEV1* coefficients exhibiting statistical significance in the correct direction (*LGBOOM* is also significant but incorrectly signed). We view crises as discrete occurrences, in which case the limited dependent variable approach is the more appropriate one.

Table 2
Determinants of crises: 1980–1997. Dependent variable: *CRISISR*, dichotomous measure of depreciation

Variable	Basic (1)	Basic + Corruption (2)	Basic + Flight (3)	Basic + Corruption & Flight (4)	M1 instead of Domestic Credit (5)	Basic + Corruption & Flight & Disequil'm (6)	Basic w/ D90s (7)
RS_DCR	■■■	■■■	■■■	■■■		■■■	■■■
LGBOOM	■□□	■□□	■□□	■□□	■■□	□□□	■□□
R^{US}	□□□	●□□	●□□	●□□	●□□	□□□	■□□
LTRANSPRNT96		●□□		□□□	□□□	□□□	
CFLT(-8)			□□□	□□□	□□□	□□□	
RS_MRT(-1)					□□□		
DEV1						■□□	
TB_Y						□□□	
D90S							
D90S×RUS							■■□
McF. R^2	.12	.13	.13	.14	.07	.15	.16
N	621	621	522	522	522	507	621
# Crises	20	20	20	20	20	20	20

Notes: OLS regression results on panel without fixed country effects. Dependent variable is ***CRISISR*** takes on a value of unity when the log differenced quarterly real exchange rate is less than -0.20 (see text). ***RS_DCR*** is the foreign exchange to domestic credit ratio; ***LGBOOM*** is the 2 year change in the log real domestic credit; $\boldsymbol{R^{US}}$ is the real US Fed Funds rate. ***LTRANSPRNT96*** is the log of the inverse corruption index in 1996, ***CFLT*** is capital flight to GDP ratio. ***RS_MRT*** is the foreign exchange to *M*1 ratio; ***TB_Y*** is the trade balance to GDP ratio; ***DEV1*** is the log-deviation of the real exchange rate from linear time trend; ***D90S*** is a dummy variable for observations in the 1990's. McF R^2 is the McFadden R^2 statistic; *N* is the number of observations; ***# Crises*** is the number observations where the dependent variable takes a value of unity. ●□□ {■□□} (■■□) [■■■] denotes significance at 20% {10%} (5%) [1%] marginal significance level in the anticipated direction.

cates that the key variable exhibits statistical significance. This is a robust finding across all regressions.

The lagged lending boom also exhibits statistical significance at the 10% level, a result echoed in the other specifications. Finally, the real US interest rate is not significant in this simplest of specifications. Next, we augment the basic specification with either an inverse corruption measure (column 2) or capital flight (column 3). In the latter case, capital flight does not show up as important, while the real interest rate is now significant at the 20% level. The inverse corruption measure is also marginally significant. However, inclusion of both of these measures (which an attempt to proxy for the same factor) yields insignificant coefficients on both. Nonetheless, the key variables—reserves to domestic credit and the lagged lending boom—are still statistically significant.

In the next three columns we report results of three robustness checks. First, one might argue that reserves always decline prior to a crisis, and so it is not surprising that we find statistical significance for this variable. However, the evidence does not appear to bear out this assertion. The reserves to M_1 ratio does not show statistical significance (column 5); moreover, the McFadden R^2 drops from 0.22 to 0.15. This pattern of results is repeated elsewhere. The reserves to M_2 ratio works somewhat

better (results not reported), but is still inferior in performance to that of the reserves to domestic credit ratio.

We also examine whether our results are sensitive to the inclusion of variables that are not important in our model, but are in others—namely the trade balance to GDP ratio and the real exchange rate overvaluation, *DEV1*. Only the latter is significant in column 6, while the lending boom becomes insignificant. The *RS__DCR* coefficient is robust to the inclusion of these additional regressors.

How does one interpret the statistical significance of *DEV1*? Given that the dependent variable is based on the change in the real exchange rate, we view these results as confirming the Goldfajn and Valdes (1995) finding that large real appreciations are reversed by discrete depreciations. Furthermore, given that the trade balance does not enter significantly, we believe that these correlations do not speak to currency crises directly.

Finally, we ask if the 1990s are different from the 1980s. A number of specifications, including a dummy and slope interaction terms to account for the possibility of a structural change, are estimated. In column 7, the results for a representative specification are shown. The key variables show up as significant, while the only statistically significant change is in the slope coefficient for the US real interest rate. In words, this means that during the 1990s, crises are more likely to occur for a given US real interest rate increase than was the case during the 1980s.

In Table 3, we report the results using *CRISI2R*, based on an exchange market pressure variable. The results are largely in line with those in Table 2, with the following exceptions: the lending boom variable is nowhere significant, and the capital flight variable is usually significant when it is included in a regression. Further note that in column 5, the reserves to M_1 ratio is completely uninformative about crises; moreover the R^2 drops substantially. The only variable of significance is capital flight. In column 6, we find that the role of real exchange rate overvaluation is muted with this alternative definition of a currency crisis. The difference between the 1980s and the 1990s shows up only in the slope coefficient associated with the US real interest rate. In words, the estimates in column 7 indicate that real interest rates did not matter in the 1980s, while they did in the 1990s.

In the regressions underlying the results in Tables 2 and 3, we have set a fairly high standard—we seek to explain only the observations at the *onset of a crisis*. In most previous analyses, the data analysed has been annual, so that the crisis is implicitly assumed to last for a year. If we adopt a similar perspective, defining the three subsequent quarters as crisis periods also, the results in Table 4 are obtained. One notable finding is that the lending boom variable drops out of significance (and is incorrectly signed). However, the reserves to domestic credit ratio and the US real interest rate are always statistically significant; moreover, the inverse corruption and capital flight variables are also statistically significant when the *RS__DCR* ratio is used. (If one looks to the specification using the reserves to M_1 ratio in column 5, one finds that *LTRANSPRNT96* and R^{US} are no longer significant at the conventional levels, although capital flight is.)

Note that in column 6, the disequilibrium measures drop out of significance. In particular, the real exchange rate deviation is completely irrelevant. Finally, account-

Table 3
Determinants of crises: 1980–1997. Dependent variable: *CRISI2R*, dichotomous measure of change in exchange market pressure

Variable	Basic (1)	Basic + Corruption (2)	Basic + Flight (3)	Basic + Corruption & Flight (4)	M1 instead of Domestic Credit (5)	Basic + Corruption & Flight & Disequil'm (6)	Basic w/ D90s (7)
RS_DCR	■■■	■■■	■■■	■■■		■■□	■■■
LGBOOM	□□□	□□□	□□□	□□□	□□□	□□□	
R^{US}	□□□	□□□	□□□	□□□	□□□	□□□	
LTRANSPRNT96		□□□		□□□	□□□	□□□	
CFLT			■□□	■□□	■■□	●□□	■■□
RS_MRT					□□□		
DEV1						●□□	
TB_Y						□□□	
D90S							●□□
D90S×RUS							■□□
McF. R^2	.08	.09	.11	.11	.07	.12	.12
N	621	621	522	522	522	507	621
# Crises	17	17	17	17	17	17	17

Notes: OLS regression results on panel without fixed country effects. Dependent variable is ***CRISI2R*** takes on a value of unity when the log differenced quarterly exchange market pressure variable is less than -0.20 (see text). ***RS_DCR*** is the foreign exchange to domestic credit ratio; ***LGBOOM*** is the 2 year change in the log real domestic credit; $\boldsymbol{R^{US}}$ is the real US Fed Funds rate. ***LTRANSPRNT96*** is the log of the inverse corruption index in 1996, ***CFLT*** is capital flight to GDP ratio; ***RS_MR***T is the foreign exchange to *M*1 ratio; ***TB_Y*** is the trade balance to GDP ratio; ***DEV1*** is the log-deviation of the real exchange rate from linear time trend; ***D90S*** is a dummy variable for observations in the 1990's. McF R^2 is the McFadden R^2 statistic; *N* is the number of observations; *# Crises* is the number observations where the dependent variable takes a value of unity. ●□□ {■□□} (■■□) [■■■] denotes significance at 20% {10%} (5%) [1%] marginal significance level in the anticipated direction.

ing for a break at 1990, one finds that the US real interest rate effect is coming almost completely from the 1990s.

One link to the banking crisis literature is notable. Industrialized country real interest rates appear to be key determinants of banking crises (Eichengreen and Rose, 1998; Hutchison and McDill, 1998). Interestingly, these authors also find that real overvaluation is not central to the onset of a banking crisis.

4.3. Some robustness checks

In this subsection we check that our results are not being driven by: (i) the particular selection of time periods and countries; (ii) the use of dichotomous dependent variables; or (iii) the inappropriate imposition of coefficient restrictions.

The results in Tables 2–4 suggest that the 1990s were somewhat different from the earlier period, especially in terms of the role for US real interest rates. Those results assumed the same error variance in the 1980s as the 1990s. To relax this assumption, we stratified the sample, and estimated the "Basic + Corruption + Flight" specification over only data for the 1990s. We also estimated a specification augmented by the disequilibrium measures—the overvaluation measure and the trade

Table 4
Determinants of crises: 1980–1997. Dependent variable: *CRISIS_P*, dichotomous measure of change in exchange market pressure

Variable	Basic (1)	Basic + Corruption (2)	Basic + Flight (3)	Basic + Corruption & Flight (4)	M1 instead of Domestic Credit (5)	Basic + Corruption & Flight & Disequil'm (6)	Basic w/ D90s (7)
RS_DCR	■■■	■■■	■■■	■■■		■■□	■■■
LGBOOM	○□□	○□□	○□□	□□□	□□□	□□□	○□□
R^{US}	■■□	■■□	■■■	■■□	●□□	■■□	□□□
LTRANSPRNT96		■■■		■■■	□□□	■■□	■■□
CFLT			■■■	■■■	■■■	■■■	■■■
RS_MRT					■■■		
DEV1						□□□	
TB_Y						●□□	
D90S							□□□
D90S×RUS							■■□
McF. R^2	.17	.19	.21	.22	.15	.22	.25
N	621	621	522	522	522	507	522
# Crises	74	74	74	74	75	74	74

Notes: OLS regression results on panel without fixed country effects. Dependent variable is ***CRISIS_P*** takes on a value of unity when the log differenced quarterly exchange market pressure variable is less than -0.125, and for the subsequent three quarters (see text). ***RS_DCR*** is the foreign exchange to domestic credit ratio; ***LGBOOM*** is the 2 year change in the log real domestic credit; R^{US} is the real US Fed Funds rate. ***LTRANSPRNT96*** is the log of the inverse corruption index in 1996, ***CFLT*** is capital flight to GDP ratio. ***TB_Y*** is the trade balance to GDP ratio; ***RS_MR***T is the foreign exchange to *M*1 ratio; ***DEV1*** is the log-deviation of the real exchange rate from linear time trend; ***D90S*** is a dummy variable for observations in the 1990's. McF R^2 is the McFadden R^2 statistic; *N* is the number of observations; ***# Crises*** is the number observations where the dependent variable takes a value of unity. ●□□ {■□□} (■■□) [■■■] denotes significance at 20% {10%} (5%) [1%] marginal significance level in the anticipated direction. ○□□ indicates significance level at the 20% marginal significance level in the *un*anticipated direction.

balance. These results are reported in Table 5, for the dependent variables *CRISI2R* and *CRISIS__P*. The role of the reserves to domestic credit ratio is preserved in all cases. The US real interest rate shows up as significant in all cases save one—using *CRISI2R* as a dependent variable and including the disequilibrium measures. The real overvaluation also shows up as significant here. However, in explaining *CRISIS__P*, both US real interest rates and capital flight are statistically significant determinants. Interestingly, *LGBOOM* is also significant regardless of whether *DEV1* and *TB__Y* are included (neither of these are statistically significant).[15]

Finally, it has been popular to argue that East Asia is governed by a different set of economic laws than those that apply to the rest of the emerging markets. Econometrically, this proposition reduces to assessing whether imposing common slope coefficients across Latin America and East Asia is appropriate. Defining an *ASIA* dummy variable which takes on a value of unity for Indonesia, Korea, Malay-

[15] Although the standard overvaluation measure is not a robust indicator of crises, an alternative measure that takes into account issues of real exchange rate nonstationarity (Chinn, 1998) does perform better across all specifications.

Table 5
Determinants of crises: 1990–1997

	CRISI2R		CRISIS_P	
Variable	Basic + Corruption & Flight (1)	Basic + Corruption & Flight & Disequil'm (2)	Basic + Corruption & Flight (3)	Basic + Corruption & Flight & Disequil'm (4)
RS_DCR	■■□	■■□	■■■	■■■
LGBOOM	□□□	□□□	■■□	■■□
R^{US}	■■□	□□□	■■■	■■■
LTRANSPRNT96	□□□	□□□	□□□	□□□
CFLT	■■□	□□□	■■□	■□□
DEV1		■■□		□□□
TB_Y		□□□		□□□
D90S				
D90S×RUS				
McF. R^2	.22	.35	.25	.26
N	290	275	290	275
# Crises	8	8	30	30

Notes: OLS regression results on panel without fixed country effects. Dependent variable is either ***CRISI2R*** or ***CRISIS_P*** (see text for description). ***RS_DCR*** is the foreign exchange to domestic credit ratio; ***LGBOOM*** is the 2 year change in the log real domestic credit; $\boldsymbol{R^{US}}$ is the real US Fed Funds rate. ***LTRANSPRNT96*** is the log of the inverse corruption index in 1996, ***CFLT*** is capital flight to GDP ratio. ***TB_Y*** is the trade balance to GDP ratio; ***DEV1*** is the log-deviation of the real exchange rate from linear time trend; ***D90S*** is a dummy variable for observations in the 1990's. McF R^2 is the McFadden R^2 statistic; *N* is the number of observations; *# Crises* is the number observations where the dependent variable takes a value of unity. ●□□ {■□□} (■■□) [■■■] denotes significance at 20% {10%} (5%) [1%] marginal significance level in the anticipated direction.

sia, the Philippines, Singapore and Thailand, we re-estimated the specifications in Table 4 allowing for a mean shift, and slope interaction terms. The results (not reported) do not indicate any robust slope differences. The mean shift coefficient is often significant, but this probably reflects the 1990s (no East Asian country experienced a crisis in the 1980s). While in a specification including only *RS__DCR*, *LGBOOM* and R^{US}, an interaction term on *RS__DCR* may show up as significant, this finding disappears when estimating the complete equation ("Basic + Corruption + Flight"). Finally, a Wald test for the restriction that all the slope interaction terms are jointly zero fails to reject the null hypothesis.

5. Conclusions

We have argued that the predictions of the insurance model suggest a declining reserves to bank liabilities ratio (holding asset quality constant) as the crisis approaches. Asset quality may not be constant, and empirical evidence suggests that

it deteriorates after a large burst in domestic credit growth, as occurred in all the crises analysed here. Furthermore, the insurance model makes a prediction regarding capital flight, validated by the data, not explicitly made in the other theoretical frameworks. In summary, since we are interested in avoiding future crises, it would seem prudent to subject the insurance model to further empirical testing.

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