Currency Crises:
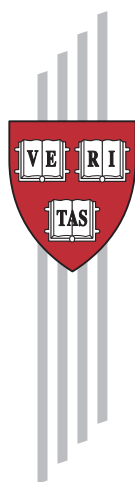# Is Central America Different?

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CURRENCY CRISES: IS CENTRAL AMERICA DIFFERENT?

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Abstract

In a recent paper we analyzed the determinants of currency crises in a sample of 30 high and middle income countries (Esquivel and Larraín, 1998). In this work we focus on Central America and analyze whether the determinants of currency crises in this region are different from those identified in our previous work. We conclude that they are not, and show that a small set of macroeconomic variables helps to explain the currency crises that took place in Central America between 1976 and 1996. The results of tests applied here support the empirical approach that attempts to explain currency crises by focusing on the behavior of a few macroeconomic indicators. Part of the interest of this result stems from the fact that the Central American countries had an exchange rate system markedly different from that prevailing in the economies that are usually analyzed in similar studies.

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CURRENCY CRISES: IS CENTRAL AMERICA DIFFERENT?

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1. Introduction

In a recent paper we analyzed the determinants of currency crises in a sample of 30 high and middle income countries for the period 1976-1996 (Esquivel and Larrain, 1998, E-L hereafter). In that work we found that high seignorage rates, large current account deficits, real exchange rate misalignments, low foreign exchange reserves, negative terms of trade shocks, negative per capita income growth, and a regional contagion effect, are highly significant in explaining the presence of currency crises in our sample. These results were robust to changes in the specification, definition of variables, method of estimation and country sample.

The purpose of this paper is to study whether the small group of macroeconomic variables that we identified in our previous work is also useful in explaining the presence of exchange rate crises in Central American (CA) countries. There are several reasons to be interested in studying the determinants of currency crises in this group of countries. First, CA countries are poorer relative to the sample in our previous work. Thus, results based in our expanded sample may therefore shed light on the validity of extrapolating our previous conclusions to other economies. Second, CA countries belong to a single geographical region and they tend to have strong commercial relationships. In fact, these countries have a trade agreement that goes back to the 1960s (the Central American Common Market). This characteristic allows us to test again for the existence of a regional contagion effect.

Third, currency crises in Central America are also interesting because the countries of this region (with the exception of Nicaragua) seem to have achieved greater exchange rate stability than most of their Latin American neighbors. Indeed, some Central American countries managed to keep a fixed exchange rate *vis-a-vis* the American dollar for a very long period of time.[1] However, these periods were often accompanied by large black market exchange rate premiums (Gaba, 1990). The presence of a black market in this context may therefore seriously undermine the explanatory and predictive power of any model that attempts to explain sudden changes in the nominal exchange rates. In this sense, we believe that this exercise provides one of the most challenging tests that we can impose on the whole approach of trying to explain currency crises based on the behavior of a small set of macroeconomic variables

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[1] In Honduras, for example, the fixed parity lasted for more than seven decades!

In addition to this introduction, the rest of the paper is organized as follows. Section 2 provides a brief historical description of the nominal and real exchange rate trends in Central America. In Section 3 we establish the criteria that we use to identify the presence of currency crises in our sample. Section 4 describes the data and the econometric methodology. In Section 5 we employ an econometric methodology to estimate the one-step-ahead probability of a currency crisis in our expanded sample. There we also evaluate whether or not including the Central American countries in our sample makes a difference in our results. In section 6 we evaluate the in-sample explanatory power of our estimated model for the specific case of the Central American countries. Section 7 concludes.

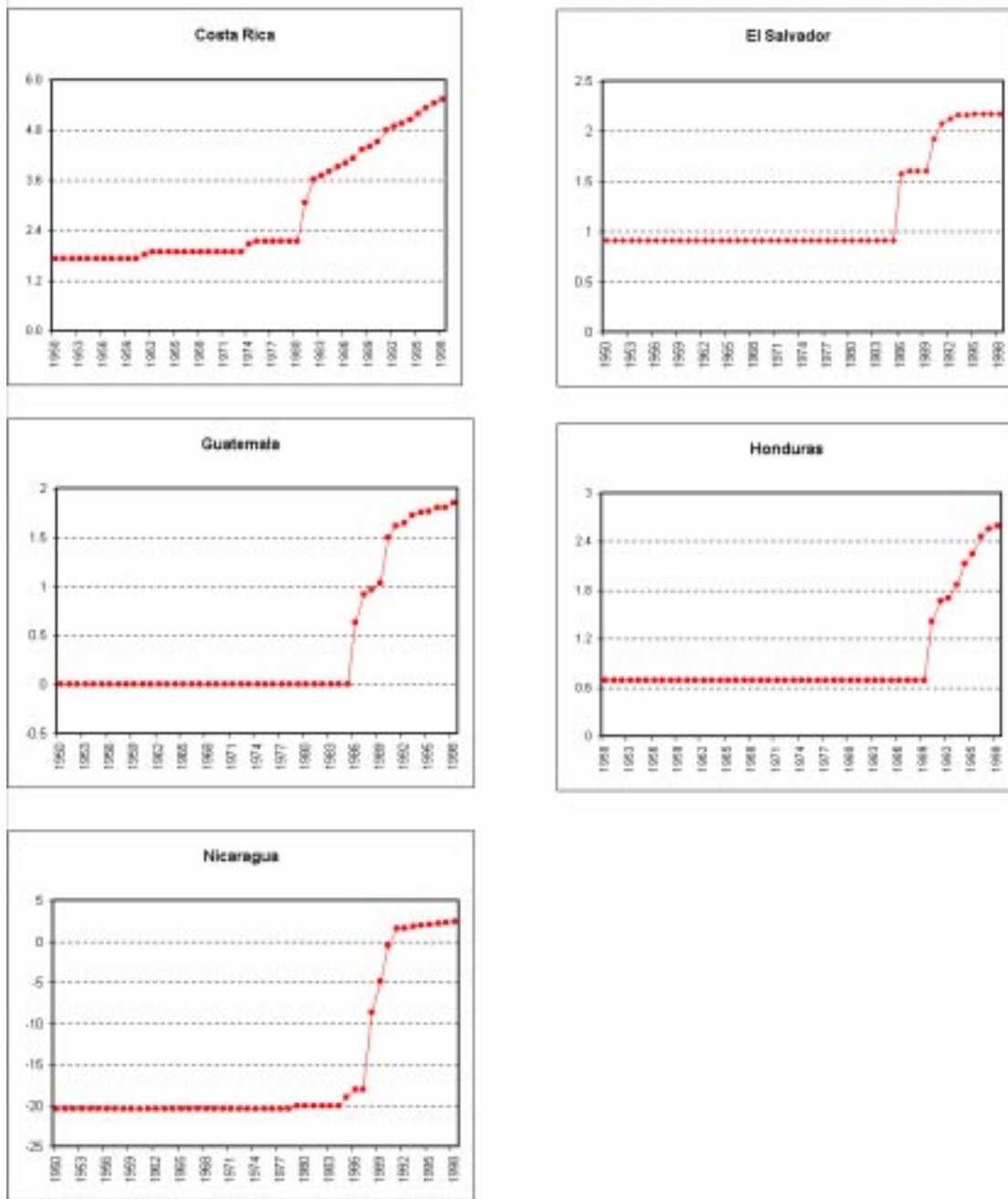## 2. Exchange Rates in Central America

Nominal Exchange Rates

Figure 1 shows the evolution of the nominal exchange rate for the five Central American countries between 1950 and 1998. The exchange rate is in local currency units per U.S. dollar and Figure 1 uses logs to illustrate proportional changes in the exchange rate. As can be seen in this figure, the exchange rate history of Central America is somewhat different from that of the rest of Latin America. Unlike most Latin American countries, the small economies of Central America were able to keep a fixed exchange rate parity vis-a-vis the U.S. dollar for a very long period.

The fixed parity of the Honduran currency was the longest in Central America and it survived intact until 1990, when the government was forced to devalue. El Salvador and Guatemala were able to sustain a fixed parity only until the mid-eighties, when their currencies collapsed in the midst of internal civil conflicts and when the debt crisis in Latin America was at its worst. Nicaragua, on the other hand, devalued its currency in early 1979 at a time of tremendous civil unrest that culminated with the fall of the Somoza regime and the Sandinista takeover. In the region, Costa Rica had the least stable parity against the dollar in the pre-debt crisis period. In fact, Costa Rica had to adjust the value of its currency as early as 1961 and then again in 1974 and 1981.[2]

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[2] For more information on the exchange rate history of Central American countries see Bulmer-Thomas (1987), Edwards (1995), Edwards and Losada (1994), and Gaba (1990).

Figure 1   Central America: Annual Nominal Exchange Rates
(in logs of local currency units per dollar)

Source: International Financial Statistics, IMF.

Figure 1 also provides hindsight on the exchange rate regimes that followed the collapse of the fixed parities in Central America. On the one hand, both Costa Rica and Honduras steadily moved towards more flexible exchange rate arrangements. Costa Rica, after repeated unsuccessful attempts to stabilize its currency (none of which lasted more than a few months), began to target its real exchange rate in mid-

1985 with relative success (see Figure 2). Honduras decided to implement a more flexible exchange rate policy in late 1992, after almost two years of attempting to restore the stability of its currency at 5.4 *lempiras* per U.S. dollar (up from the exchange rate of 2 *lempiras* per dollar prevalent until February of 1990). Since then, the Honduran currency has fluctuated relatively freely.

On the other hand, El Salvador, Guatemala and Nicaragua attempted to confront the collapse of their currencies with a new fixed parity against the U.S. dollar. The shortest-lived of these experiments was in Guatemala, where the attempt to establish a fixed parity of 2.5 *quetzales* per dollar in 1986 (up from a one-to-one parity) was abruptly abandoned 2 years later. In El Salvador and Nicaragua, the new pegs lasted longer. In 1986, El Salvador devalued its currency from 2.5 to 5 *colones* per dollar. The new parity lasted for approximately four years until it collapsed again in May 1990. Since 1993, the currency of El Salvador has remained fixed against the dollar, this time at a rate of 8.755 pesos per dollar. On the other hand, in 1979 the recently installed Sandinista government of Nicaragua chose to keep its exchange rate fixed against the dollar. The parity was sustained until 1985, when the accumulated domestic inflation and the external conditions made inevitable the adjustment of the exchange rate.[3] Since 1992, Nicaragua has implemented a managed float with pre-announced daily changes of the exchange rate.

Real Exchange Rates

Figure 2 shows the monthly multilateral real exchange rates of the five Central American countries for the period of 1970 to 1998.[4] The real exchange rate indexes are trade-weighted and they are depicted using December 1990 = 100. A rise in the exchange rate represents a real depreciation. Two interesting facts emerge from this figure. First, the end of the fixed exchange rate parities in Central America is easily identifiable by the pronounced peaks that are observed in these graphs. As can be observed in Figure 2, the end of the fixed parities in El Salvador, Guatemala, and Honduras was preceded by a substantial appreciation of their real exchange rate. This phenomenon is less evident in the case of Costa Rica and was definitely not a factor in the abandonment of the fixed parity in Nicaragua in 1979. In the next section, we return to the link that seems to exist between nominal and real exchange rates movements.

A second interesting aspect that emerges from Figure 2 is the clearly differentiated pattern followed by the real exchange rates of the Central American countries in the post-crisis period. On the one hand, Costa Rica, Honduras and, to a lesser extent, Guatemala, all ended with a substantially depreciated real exchange rate relative to their eighties level. On the other hand, El Salvador and

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[3] For a detailed account of the exchange rate experience of Nicaragua see Ocampo (1991) and Gibson (1993).
[4] The exception is Nicaragua, for which we have information from 1973 onwards.

Nicaragua ended with an appreciated real exchange rate. In both cases, their real exchange rate in 1998 is slightly above one third of the one that prevailed until 1979.

Figure 2 — Central America: Monthly Real Exchange Rates
(December 1990=100)
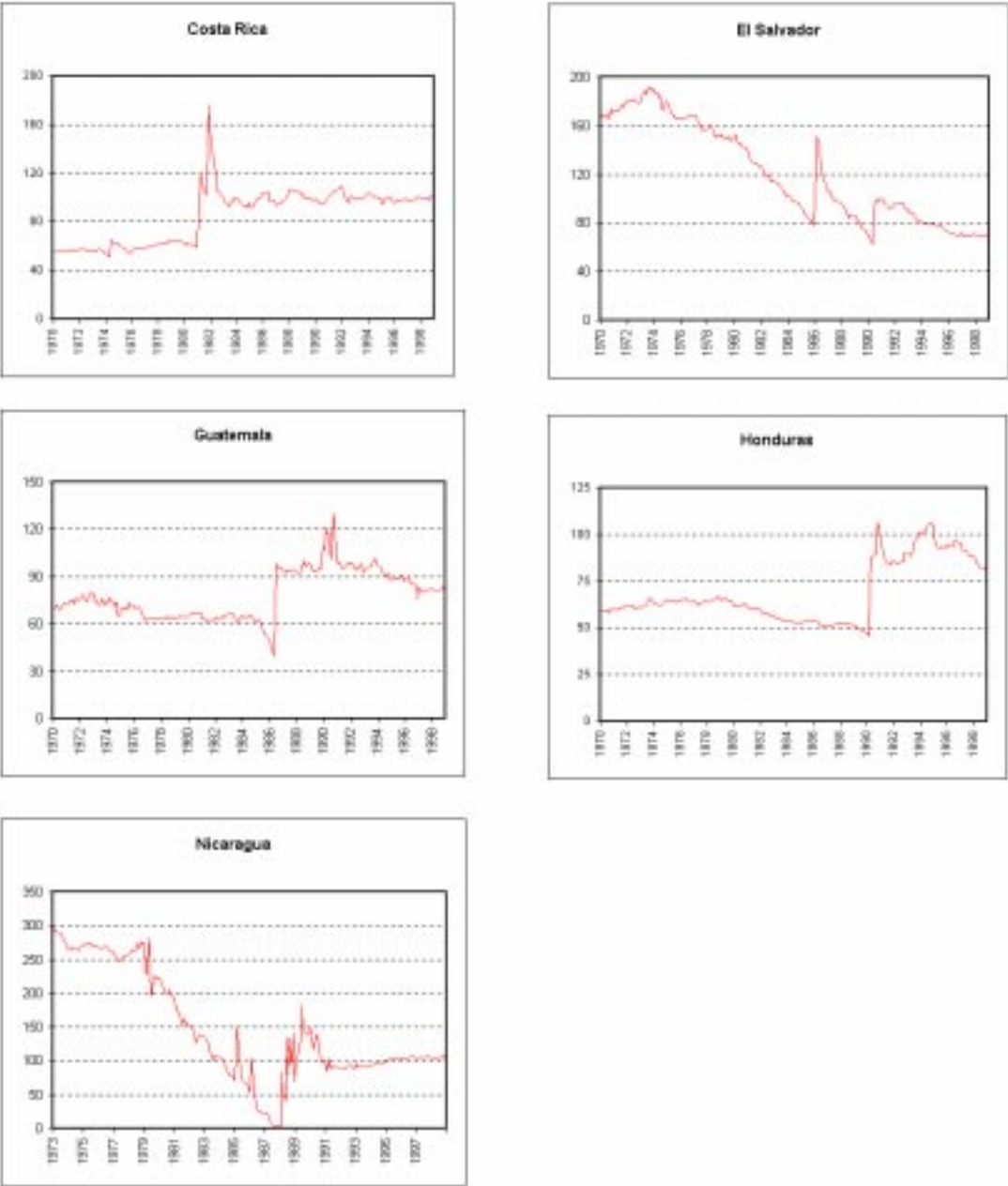
Note: Multilateral real exchange rates. Estimates are trade-weighted.

Source: Authors' calculations based on IMF and UNCTAD data.

3. Definition of Currency Crisis

In this section of the paper we present our definition of currency crisis and the criteria we use to identify these situations in our sample.[5]

In our view, a currency crisis exists only when there is an important change in the nominal exchange rate. Thus, unlike some of the previous studies on the topic, we exclude unsuccessful speculative attacks from our definition of crisis.[6] We exclude these episodes from our definition of crisis because we consider identifying unsuccessful speculative attacks a very difficult and subjective task.[7]

For a nominal devaluation to qualify as a currency crisis, we use two criteria. First, the devaluation rate has to be large relative to what is considered standard in a country (we will be more precise about this later). Second, the nominal devaluation has to be meaningful, in the sense that it should affect the purchasing power of the domestic currency. Thus, nominal depreciations that simply keep up with inflation differentials are not considered currency crises even if they are fairly large. Our definition of crises therefore excludes many of the large nominal depreciations that tend to occur during high-inflation episodes.

By putting these two considerations together we conclude that a currency crisis exists only if a nominal devaluation is accompanied by an important change in the real exchange rate (at least in the short run). If we assume that the price level reacts slowly to changes in the nominal exchange rate then, in practical terms, we can detect a currency crisis simply by looking at the changes in the real exchange rate. Before doing so, however, we need to define how large the real exchange rate (RER) movement must be in order to be considered as a crisis.

We consider that a currency crisis has occurred when at least one of the following conditions is met:

Condition A:
The accumulated three-month real exchange rate change is 15 percent or more
or,

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[5] This section draws on Esquivel and Larraín (1998). The reader is referred to that work for further details on the methodology.
[6] Some of the papers that prefer to include these events in their definition of crises are Eichengreen, Rose and Wyplosz (1995), Kaminsky and Reinhart (1999), and Sachs, Tornell and Velasco (1996).
[7] See, for example, the discussion on the "speculative pressure index" in Flood and Marion (1998).

<u>Condition B</u>:

The one-month change in the real exchange rate is higher than 2.54 times the country specific standard deviation of the RER monthly growth rate, provided that it also exceeds 4 percent, i.e.:

$$\Delta\varepsilon_{it} > 2.54\ \sigma_i^{\Delta\varepsilon} \text{ and } \Delta\varepsilon_{it} > 4\%,$$

where ε_{it} is the real exchange rate (RER) (RER) in country i in period t, $\Delta\varepsilon_{it}$ is the one-month change in the RER, and $\sigma_i^{\Delta\varepsilon}$ is the standard deviation of $\Delta\varepsilon_{it}$ in country i over the whole period.

Condition A guarantees that any large real depreciation is counted as a currency crisis. The threshold value of 15 percent is certainly somewhat arbitrary, but sensitivity analysis shows that the precise threshold is largely irrelevant for our results.[8] Condition B, on the other hand, attempts to capture changes in the RER that are sufficiently large relative to the historical country-specific monthly change of the RER.[9]

4. Data and Econometric Methodology

Data

In E-L (1998) we estimated a model on the determinants of currency crises for a group of 30 high and middle income countries from 1975 through 1996. [10] In this paper we add Costa Rica, El Salvador, Guatemala, Honduras, and Nicaragua to our original sample. The real exchange rate measure that we use for the CA countries was described in section 2, with remaining variables obtained from various international sources (IMF, The Word Bank, and UNCTAD).

Since we are interested in both explaining and forecasting currency crises, most explanatory variables enter in lagged form. Thus, explanatory variables run from 1975 to 1995, whereas our dependent variable goes from 1976 to 1996. We then have a panel dataset of 21 years for 35 countries, which makes a total of 735 potential observations. Since our dependent variable is dichotomous and takes

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[8] Other authors have also used thresholds in their definition of *crisis*. Frankel and Rose (1996), for example, use a 25 percent *nominal* exchange rate change as a threshold value. Eichengreen, Rose and Wyplosz (1995), Goldfajn and Valdes (1998), and Kaminsky, Lizondo and Reinhart (1998) have instead used a definition that is closer to our condition B.

[9] Assuming that changes in the RER are normally distributed, condition B is defined as to capture changes in the RER that lie in the upper 0.5% of the distribution for each country.

[10] The original sample consisted of the following countries: Argentina, Australia, Belgium, Brazil, Chile, Colombia, Denmark, Ecuador, Finland, Greece, Indonesia, Ireland, Italy, Korea, Malaysia, Mexico, Morocco, New Zealand,

the value of 1 when there is a crisis and 0 otherwise, our fitted values may then be interpreted as the one-step-ahead probabilities of a currency crisis.

Episodes of Crises in Central America

When our two conditions are applied to the Central American countries during the period 1970 trough 1998, we identify fourteen episodes of currency crises in the region. The dates and months of these events are shown in Table 1. As discussed above, the application of our two criteria to identify instances of crisis produces results that coincide very well with situations of abrupt movements in the nominal exchange rates in Central American countries.

Table 1. Currency Crises in Central America

Country	**Date**
Costa Rica	• May 1974 • January 1981
El Salvador	• January 1986 • May 1990
Guatemala	• June 1986 • August 1990
Honduras	• March 1990
Nicaragua	• March 1979 • February 1985 • February 1986 • February 1988 • March 1989 • May 1990 • March 1991

Norway, Peru, Philippines, Portugal, Singapore, Spain, Sweden, Switzerland, Thailand, Turkey, United Kingdom and Venezuela.

Number of Crises in the Sample

Through the imposition of conditions A and B to our monthly real exchange rate dataset we have identified 124 episodes of crisis, of which 111 correspond to countries in our original sample and 13 to Central American countries.[11] Figure 3 shows the number of currency crises per country in the Latin American countries in our sample.[12] This Figure confirms that, with the exception of Nicaragua, Central American countries have had much more stable currencies than the other countries in the subcontinent.

Figure 3. Number of Currency Crises in Latin American Countries, 1976-96

Number of Currency Crises (1976-96)

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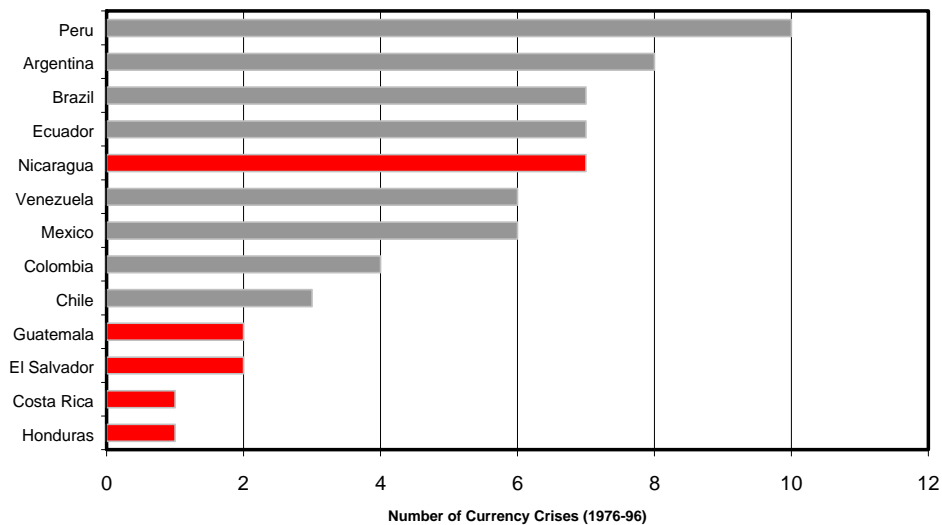
[11] Note that we "lose" an episode of crisis in Central America, because the 1974 crisis in Costa Rica does not lie within our estimation period.

[12] For a more detailed description of the currency crises for other countries in our sample, the reader is referred to Esquivel and Larraín (1998).

Estimation Methodology

We now describe our approach in estimating the determinants of currency crises. The variable to be explained (y_{it}) is dichotomous, and takes the value of 1 if a currency crisis occurred during year t and 0 otherwise. We estimate a *probit model* of the form:

$$\text{Prob (Crisis}_{it}) = \text{Prob } (y_{it}=1) = \Phi \ (\boldsymbol{b'x}_{it-1})$$

where $\boldsymbol{x}_{it-1}$ is a vector of explanatory variables for country i in period t-1, $\boldsymbol{b}$ is a vector of coefficients to be estimated, and Φ is the normal cumulative distribution function.

Note that in our estimation we are implicitly assuming the existence of a an unobservable or latent variable (y_{it}^*) which is described by

$$y_{it}^* = \boldsymbol{b'x}_{it-1} + u_{it}$$

where $\boldsymbol{x}_{it-1}$ and $\boldsymbol{b}$ are as before, u_{it} is a normally distributed error term with zero mean and unit variance, and the observed variable y_{it} behaves according to $y_{it} = 1$ if $y_{it}^* > 0$, and $y_{it} = 0$ otherwise. Please note that in this regard we depart slightly from E-L (1998), since in that paper we used a probit model with random effects.[13]

Explanatory Variables

The explanatory variables that we use in this paper are the same that we used in E-L (1998). For completeness, here we present a brief description of each one of them.[14]

Seignorage. This variable, defined as the annual change in reserve money as a percent of GDP, attempts to capture Krugman's original insight that monetization of the government deficit is key to explaining exchange rate collapses. We expect this variable to have a positive effect on the probability of a crisis.

 Real Exchange Rate Misalignment. This variable is defined as the negative of the percentage deviation of the real exchange rate from its average over the previous 60 months.[15] This definition makes our variable easily comparable across both time and countries. An increase in the RER misalignment is expected to increase the risk of a currency crisis.

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[13] In E-L (1998) we showed that there are no substantial differences in the results obtained with alternative estimation methods. For this reason, and to simplify the exposition of our results, in this work we prefer to use a more standard econometric methodology.

[14] For more details in the construction of these variables and for a justification of their inclusion in the empirical analysis see E-L (1998).

[15] Note that the RER misalignment variable is defined so that RER appreciation (or overvaluation) with respect to the previous 5-year average enters with a positive sign. An *increase* in the misalignment variable then represents a larger appreciation and a higher risk of a crisis.

Current Account Balance. A deterioration of the current account balance is expected in anticipation of a currency crisis. Therefore, we expect to find a negative relationship between the current account balance and the probability of crisis. This variable enters as a percentage of GDP.

M2/Reserves. This variable is the ratio of a broad definition of money to official foreign exchange reserves. It attempts to capture the vulnerability of the central bank to possible runs against the currency. This variable is in logs and we expect to find a positive association between this ratio and the probability of a crisis.[16]

Terms of Trade Shock. This variable is defined as the annual percentage change in the terms of trade and we expect a negative relationship between this variable and the probability of crisis.

Per Capita Income Growth. A negative per capita income growth is assumed to increase the policymaker's incentives to switch to a more expansionist policy, which can be achieved through a nominal devaluation of the currency. Our variable is dichotomous and takes the value of 1 if per capita income growth is negative in a given year and 0 otherwise. Consequently, we expect a positive coefficient associated to this variable.

Contagion Effects. It has recently been argued that crises can be transmitted across countries through many different channels (Drazen, 1998). Most of the likely explanations, however, suggest that contagion effects tend to occur at the regional level.[17] In consequence, in order to capture the possibility of a contagion effect, we first define geographical regions. Next, we specify a dichotomous variable that takes the value of 1 for countries belonging to a region where at least one other country has had an exchange rate crisis in the current year, and 0 otherwise. [18]

5. Empirical Results

Table 2 shows the results that we obtain when we apply the econometric methodology described above to our data. We use the individual and joint significance of the coefficients and a pseudo-R^2 measure (the so-

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[16] Other authors have suggested to use the short-term debt/Reserves ratio to capture this effect (see, for example, Sachs and Radelet, 1998). However, cross-country information for this variable is available only starting in 1986. We therefore prefer the M2/Reserves variable.

[17] See Glick and Rose (1998) and the models discussed in Drazen (1998).

[18] We have defined the following regions: Europe, Asia, Oceania, Central America and the rest of Latin America. See E-L (1998) for an explanation of how did we proceed in the cases of Turkey and Morocco.

called McFadden's R^2) to evaluate the goodness-of-fit of our model.[19] All regressions include annual dummies and a constant,[20] and the estimated parameters have been transformed so that the reported coefficients can be interpreted as the change in probability associated to a unit change in the explanatory variables.

The numbers in parentheses in Table 2 are z-statistics that test the null hypothesis of no significance of the parameters associated to the explanatory variables. We use asterisks to identify the coefficients' level of significance. The next-to-bottom part of each table includes a chi-square statistic (and its associated p-value) which tests for the joint significance of all coefficients other than the constant and the time dummies. The bottom part of the table shows the results of a Lagrange test that is described below.

Column (1) in Table 2 presents the estimates when we use our original sample of 30 high and middle income countries. As in E-L (1998), all the coefficients have the expected signs and they are statistically significant at conventional levels. Moreover, they are jointly significant at the one-percent level.

Column (2) of Table 2, on the other hand, shows the results when we expand our dataset to include the five Central American countries. There are major differences between regressions (1) and (2) in Table 2. Three coefficients present an important reduction in their absolute value relative to column (1) (those associated to the seignorage, real exchange rate misalignment and current account variables); whereas the coefficient associated to the real exchange rate misalignment loses its statistical significance when we add the CA countries. Furthermore, when we implement a Lagrange test to evaluate whether the coefficients associated to the CA countries are statistically different to the rest of our sample we obtain a statistic of around 20. This result, as indicated by its very low *p*-value, means that we strongly reject the null hypothesis that the coefficients associated to the CA countries are no different from those of the rest of our sample.

The next step in our empirical analysis is to investigate what drives these results. Is it the case that all Central American countries are different from the rest of the countries in our sample? Or, is it just

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[19] In the next section we discuss the in-sample prediction performance of our results as an alternative goodness-of-fit measure.

[20] These annual dummies are intended to capture any worldwide effect that may have an impact on the likelihood of currency crises in our entire sample of countries. Thus, these variables may capture not only effects on world interest rates but they may also be reflecting any other similar worldwide phenomenon.

that our estimates are very sensitive to the inclusion of a country, Nicaragua, whose behavior is clearly different from the rest of our sample (see Figures 1 and 2 and Table 1)? To investigate this question, column (3) in Table 2 drops Nicaragua from the sample. The new estimates present some noticeable changes. For example, now all the coefficients are statistically significant at the 5 percent level. Moreover, the absolute value of all the coefficients is now much closer to those of column (1). It is therefore clear that the reduction in the absolute value and statistical significance of some of the coefficients in column (2) was largely driven by the inclusion of Nicaragua in our sample. However, although the new empirical estimates are closer to those of our shorter sample, the Lagrange statistic that tests the null hypothesis that coefficients associated to Central American countries are no different is still rejected at a 10 percent level of significance.

One possible explanation for the latter result is that our contagion variable for Central America in regression (3) still considers the possibility that Nicaragua's currency crises may have an effect on other Central American countries. However, it is very likely that allowing for such an effect is not necessarily correct (at least for most of the period under study). It must be recalled that between 1979 and 1989 Nicaragua was ruled by the *Sandinistas*, which implemented non-market oriented economic policies. As a result, the Nicaraguan *cordoba* became strongly appreciated in real terms (see Figure 2), and international trade with non-socialist countries decreased dramatically. These circumstances suggest that practically none of the channels that are usually called in to explain the occurrence of a crisis contagion were actually effective:[21] interregional trade was limited and likely investors were clearly able to distinguish between the policies implemented by Nicaragua and those of its Central American neighbors.

The next step in our empirical analysis is to introduce a minor modification in the definition of the contagion effect for the Central American countries. The new variable is defined in such a way that the contagion effect in this region can only occur within the four CA countries included in the analysis: Costa Rica, El Salvador, Guatemala, and Honduras. The results, when we include this modification of the regional variable for Central America, are displayed in column (4) of Table 2. All of the new coefficients, with the exception of the contagion effect, are very close to those obtained in regressions (1) and (3) and they are all significant at the 5 percent level. The new estimated coefficient of the contagion effect is much larger than in previous regressions. In the new results, if a country belongs to a region where at least one other country has recently experienced a currency crisis, its probability of also having a currency crisis increases, on average, by about 11 percentage points (up from a previous estimate of about 8 percentage points). The observed increase in the contagion effect can certainly be attributed to the fact

that CA countries tend to be more closely integrated among themselves, and this in turn increases the possibility of contagion across these countries.

The most important result of column (4), however, is that with the modification of the regional variable for CA countries we can now accept the null hypothesis that coefficients for Central America are no different from those of other countries in the sample. The Lagrange test statistic for this specification takes a value of only 8.5 and it is accepted at any significance level below 25 percent. In what follows we use regression (4) as our *benchmark*. As mentioned above, coefficients are shown as marginal effects on the probability of crisis, and they are evaluated at the mean values of the explanatory variables. In consequence, the fitted values can then be interpreted as the one-step-ahead probability of a currency crisis. In the case of explanatory dummy variables, coefficients have been computed as the actual change in probability that occurs when the dummy variable switches from 0 to 1, assuming that all the other explanatory variables remain at their mean values.[22]

The first coefficient in column (4) shows that a one-percentage point increase in the rate of seignorage to GDP increases the probability of crisis in about 1.8 percentage points. Likewise, a real exchange rate misalignment of about 10 percent translates into an increase in the probability of a currency crisis of about 3.3 percentage points. Although this effect seems to be relatively small, it is important to keep in mind two considerations. First, this result is obtained after controlling for the current account balance (which is strongly associated with the RER misalignment variable). Second, RER misalignments as large as 30 percent often occur in our sample, which therefore represents an increase in the probability of a currency crisis of about 10 percentage points.

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[21] See Drazen (1998) and Glick and Rose (1998).
[22] This procedure is standard in situations with discrete explanatory variables and a qualitative dependent variable. See Greene (1996) for more details.

Table 2: Determinants of Currency Crises

Variables	Regression Coefficients (z-statistics)			
	(1)	**(2)**	**(3)**	**(4)**
Seigniorage	0.0214*	0.0149*	0.0191*	0.0180*
(as a percent of GDP)	(4.10)	(3.99)	(4.14)	(3.97)
RER Misalignment	0.0035**	0.0017***	0.0033*	0.0033*
	(2.45)	(1.86)	(2.98)	(2.98)
Current Account Balance	-0.0103*	-0.0028	-0.0070**	-0.0076**
(as a percent of GDP)	(-2.85)	(-1.28)	(-2.24)	(-2.41)
Log(M2/Reserves)	0.0411**	0.0442*	0.0477*	0.0495*
	(2.22)	(2.68)	(2.87)	(2.97)
Terms of Trade Shock	-0.0046***	-0.0048*	-0.0046**	-0.0048**
	(-1.84)	(-2.58)	(-2.25)	(-2.34)
Negative Growth Dummy	0.0810**	0.0704**	0.0641**	0.0626**
(1 if per capita income growth <0)	(2.31)	(2.26)	(2.08)	(2.05)
Contagion Effect	0.0804**	0.1029*	0.0882*	0.1163*
(1 if at least one country in the region had a crisis)	(2.39)	(3.60)	(3.08)	(3.89)
Number of Observations	629	734	713	713
Number of Countries	30	35	34	34
Log Likelihood	-242.2	-276.5	-264.4	-261.1
McFadden's R^2	0.17	0.17	0.17	0.18
Chi-square (Ho: Coefficients = 0)	45.92	52.85	51.66	53.83
(p-Value)	(0.00)	(0.00)	(0.00)	(0.00)
Lagrange Test (Ho: Central America is no different)		20.27	12.36	8.47
(p-value)		(0.005)	(0.089)	(0.293)

Note: All regressions include time dummies and a constant.
     * indicates statistical significance at a 1% level,
    ** indicates statistical significance at a 5% level.
   *** indicates statistical significance at a 10% level.

Column (4) in Table 2 also shows that a one-percentage-point increase in the current account deficit to GDP increases the probability of crisis in slightly less than 1 percentage point. The coefficient associated with the current account deficit is always negative and strongly significant in the four regressions in Table 2. As mentioned in our previous work, our current account results are of special

interest because other empirical studies summarized by Kaminsky, Lizondo and Reinhart (1999) and Glick and Moreno (1999) found this variable to be non-significant as a determinant of currency crises. It is also worth emphasizing the empirical relevance of this result since the current account has often been interpreted by analysts and practitioners as an indicator of an economy's vulnerability to a currency crisis. Therefore, our results can be seen as providing support for such interpretation.

Our fourth explanatory variable is the log of (M2/reserves); column (4) shows that a doubling of this ratio increases the probability of crisis by around 9 percentage points. This effect reflects the widely documented result that this ratio rises very quickly during the months preceding currency crises.

Column (4) also shows that a 10 percent terms-of-trade decline translates into a 5 percent increase in the probability of crisis. Additionally, a period of negative per capita income growth increases the probability of crisis by more than 6 percentage points. The magnitude of these two effects confirms the relevance of models that characterize the devaluation decision as the result of balancing conflicting policy objectives. In cases where the exchange rate is a policy variable, these results may be interpreted as providing some support to the escape-clause models developed by Obstfeld (1996) as well as to other models that stress the "political" nature of some currency crises.[23]

6. An In-sample Evaluation of the Model's Predictive Power

In this section we present an evaluation of our model's ability to predict the in-sample presence of currency crises in our sample, with special emphasis on the fitted values for Central American countries.

In E-L (1998) we assessed the overall explanatory performance of a similar empirical model by employing a standard *hits-and-misses* approach. By applying such an evaluation method to the benchmark regression of our previous study, we concluded that our estimated model was able to predict accurately more than 50 percent of all the crises events in our sample. As discussed in that work, such a rate of success is much higher than previous studies have found. The application of the hits-and-misses technique to the results presented in regression (4) in Table 2 lead to conclusions similar to those presented in E-L (1998) and therefore we will not discuss them in more detail here. Instead, in this paper we present an alternative method to evaluate the predictive performance of our empirical model.

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[23] See Drazen (1998) for a brief review of this literature.

Testing the Predictive Performance of the Model

In this section we evaluate the predictive performance of our model based on the application of a simple non-parametric test proposed by Pesaran and Timmermann (1992). This test asks whether a set of predictions for a binary event (in this case, crisis period *versus* tranquil period) is statistically better than pure random guesses. Before applying the Pesaran-Timmermann test to our model, however, we need to define a prediction (or classification) rule. Following E-L (1998) we have chosen the following prediction rule:[24]

    a) If $P_{it} > P*$ a crisis is predicted (i.e. an alarm is issued)
    b) Otherwise, a tranquil period is predicted

where $P*$ is a threshold value that ranges from 0.20 to 0.50.

Table 3 shows the result of calculating the Pesaran-Timmermann (P-T) statistic for a range of threshold values that goes from 0.20 to 0.50. Since the statistic is distributed as a standard normal, results in Table 3 suggest that we can strongly reject the null hypothesis that our predictions are no better than random guesses. That is, our classification rule has some value from a purely predictive perspective regardless of our thresholds value. Interestingly, the threshold value that maximizes the Pesaran-Timmerman statistic is $P*=0.30$, the same value selected in E-L (1998) based on an *ad-hoc* criterion.

Table 3. A Test of the Predictive Performance

P*	0.20	0.25	0.30	0.35	0.40	0.45	0.50
P-T	10.02*	10.05*	10.20*	9.89*	8.90*	8.63*	8.01*

Notes: Results are obtained using regression (4) in Table 2. The P-T statistic is distributed as a standard normal. The null hypothesis is that predictions are no better than random guesses. * Indicates that we reject the null hypothesis at 1% level of significance.

Predicting Currency Crises

Yet another method to evaluate the predictive performance of our model is by comparing the average one-step ahead probabilities of crisis in both tranquil and crisis periods. In principle, if our empirical results contain valuable information about the likely occurrence of a crisis in the near future, the average predicted probability of a crisis should be higher in periods when a crisis actually occurs in the next year than in periods when there is none.

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[24] See E-L (1998) for a discussion of the relevance of choosing an appropriate threshold value.

Figure 4 and Table 4 show the average one-step ahead probabilities of crisis in tranquil and crisis periods for all of the countries in our sample and for the Central American countries (both as a group and individually). The most obvious fact that emerges from Figure 4 is that the average predicted probability of crisis is, in all cases, higher in crisis periods than in tranquil ones. This result suggests that there is indeed some valuable information in our forecasts since they tend to anticipate a higher probability of crisis when these events occur. Interestingly, the average predicted probabilities of a crisis in the years immediately preceding a crisis event for the cases of Guatemala, El Salvador, and Honduras were 5, 4 and

Figure 4: Average One-step ahead probability of a crisis
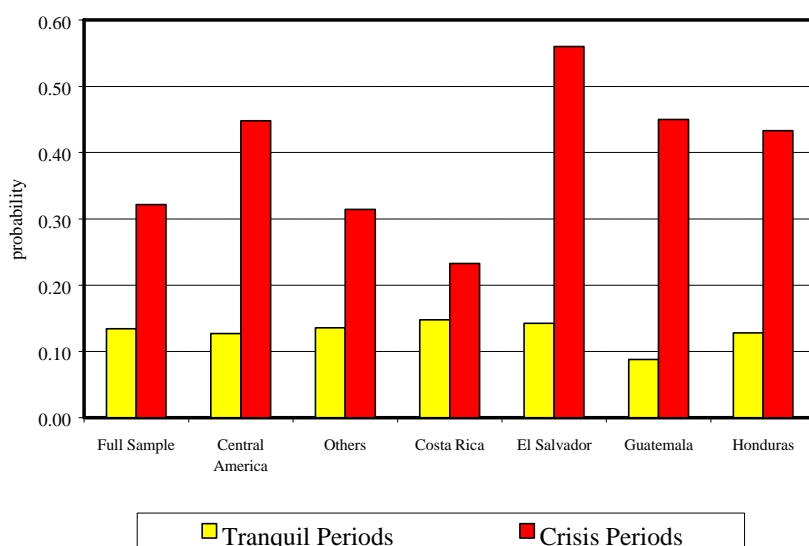
3.5 times, respectively, the average predicted probabilities in tranquil periods. These results, although suggestive, are not conclusive since we have only shown that average predicted probabilities are *numerically* higher in crisis periods than in tranquil ones, but we have not yet shown that these differences are *statistically significant.*

Table 5 shows the results of a difference-between-means test that investigates whether the average predicted probability of a crisis is statistically higher in years that precede the occurrence of a crisis or not. The test is applied to two groups of countries: Central America and to the other countries in the sample. The results of the test strongly support the conclusion that predictions about the probability of a crisis occurring next year are statistically higher in periods when a crisis actually occurs. Therefore, this result also supports our previous conclusion that forecasts based on our empirical estimates have indeed some valuable information about the likely occurrence of a currency crisis.

Table 4: Average One-step Ahead Probability of a Currency Crisis

	Tranquil Periods		Crisis Periods	
	Probability	Observations	Probability	Observations
Full Sample	0.135	596	0.321	117
Central America	0.127	78	0.448	6
Others	0.136	518	0.315	111
Costa Rica	0.148	20	0.233	1
El Salvador	0.142	19	0.560	2
Guatemala	0.088	19	0.450	2
Honduras	0.128	20	0.433	1

**Table 5.   Tests of Difference between Average
One-step Ahead Probability of Crisis**

	t	P-value
Crisis *versus* Tranquil Periods		
Central America	3.45*	0.009
Other countries	8.70*	0.000
Central America *versus* other Countries		
Crisis Periods	1.41	0.212
Tranquil Periods	-0.63	0.528

Note: * indicates that we reject the null hypothesis that means are equal
at 1 percent level of significance.

Are Predictions for Central America Different?

In Section 5 we showed that our empirical estimates for Central America are not different from the rest of the countries in our sample. Now, we will try to answer the following question: Are predicted probabilities different for Central American countries? In order to respond to this question we implement another difference-of-means test. The lower part of Table 5 shows the results that we obtain when we test whether average predicted probabilities for CA countries are the same as those of the other countries in the sample during both tranquil and crisis periods. The results of Table 5 show that we cannot reject the null hypothesis that predictions for CA are similar to those made for other countries in either tranquil or crisis periods.

7. Conclusions

In this paper we have examined whether the variables that had been found to determine the presence of currency crises in a broad sample of countries have also been important for Central American countries. Our empirical results have shown that coefficients for CA countries, once we exclude Nicaragua from the empirical estimation and as a likely source of contagion, are no different from the other countries in our sample. We have also shown that our estimates for two groups of countries (Central America and the other countries in our sample) provide valuable information as predictors of crises: the average one-step-ahead probabilities of crises tend to be statistically higher when a crisis actually occurs in the next period than otherwise. Finally, we have also shown that our average predicted probabilities for Central American countries are not significantly different from predictions made for other countries.

The results of the various tests applied in this paper support the empirical approach that attempts to explain currency crises by focusing on the behavior of a small set of macroeconomic variables. This result is relevant because the Central American countries studied in this paper had an exchange rate system markedly different from that of the economies that are usually analyzed in similar studies.

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