The Asian currency crash: were badly driven fundamentals to blame?☆

Steven Husteda,*, Ronald MacDonaldb

aDepartment of Economics, University of Pittsburgh, Pittsburgh, PA 15260, USA
bDepartment of Economics, University of Strathclyde, Glasgow G4 OLN, UK

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Abstract

This paper examines the extent to which a number of currencies central to the Asian currency crisis were misaligned at the end of 1996. A well-known fundamentals-based exchange rate model, the monetary approach to exchange rate behavior, is used to produce estimates of equilibrium exchange rates for a number of Asian currencies. The estimates, calculated using panel methods, are shown to be consistent with the underlying model. Most significantly, very little evidence of misalignment is found to exist in 1996. This suggests that the cause of the Asian crash cannot be attributed to traditional fundamentals. © 2000 Elsevier Science Inc. All rights reserved.

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1. Introduction

It has now been more than two years since the onset of the Asian financial crisis, and much research has been devoted to its causes and consequences.1 Certainly, one of the most dramatic features of this crisis has been the sharp deterioration in the exchange rates of many Asian countries. Even countries, such as Australia, India, and New Zealand, that have not usually been identified with the crisis saw sharp depreciations in their exchange rates

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* Corresponding author. Tel.: +1-412-624-6094; fax: +1-412-624-6855.
E-mail addresses: husted@fcas.pitt.edu (S. Husted), r.r.macdonald@strath.ac.uk (R. MacDonald).

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vis-à-vis the U.S. dollar during 1997. Many have cited the rapid deterioration in regional exchange rates as a principal causative factor in the financial crisis that has followed. Radelet and Sachs (1998b, p. 11) argue that the “combination of real exchange rate depreciation and sharply higher interest rates lead to a rapid rise in non-performing loans in the banking sectors of the Asian economies.”

Many commentators have focused on the behavior of nominal (and real) exchange rates prior to the collapse. Bergsten (1998) argues that the “sharp swings in the yen-dollar rate contributed to the outbreak of the Asian crisis in the first place”. The Reserve Bank of Australia (1998) also argues that weakness in the yen-dollar rate has led to deterioration in the exchange rates of other Asian economies. The logic behind this reasoning is hard to dispute. Prior to 1997, many countries in Asia linked the value of their currencies to the U.S. dollar, either by nominal or real exchange rate pegging. As the dollar rose in value against the yen, these countries experienced a loss of competitiveness in world markets relative to Japan.

Given these movements in exchange rates, and the dire implications these movements have had for the countries involved, a natural question to ask is whether or not the exchange rates of these countries were out of line with economic fundamentals prior to the crash. Surprisingly, only limited research has been done on this issue.2 This is probably due to two reasons. First, it is now accepted that the relationship between exchange rates and fundamentals is long-run. Given the short spans of data that exist since the collapse of Bretton Woods, testing these models is difficult. Added to the problem in this case is the fact that most empirical modeling of long-run exchange rate behavior has focussed on U.S. dollar exchange rates. In Asia, where many countries have tied their exchange rate to the dollar, empirical models have little relevance.

This paper seeks to address both problems described above. It does so by implementing a new tool for the analysis of exchange rate behavior, panel data estimators, on data sets involving the yen exchange rates for a set of Asian economies. Tests of long-run behavior necessarily involve long spans of data. In the context of exchange rate modeling, this requires sampling over various exchange rate regimes, which may cause hypothesized relationships to break down. An alternative way of increasing the span of the data in testing equilibrium exchange relationships involves utilizing panel data sets for the recent float (see Chinn & Johnston, 1996; Frankel & Rose, 1996; and MacDonald, 1988, 1996). Such tests have re-examined purchasing power parity (PPP) and the behavior of real exchange rates.34 In terms of panel PPP tests, researchers find that homogeneity restrictions hold and that the implied real exchange rate is strongly mean-reverting with around one-half of a disturbance to PPP being extinguished after four years.5 Most recently, Husted and MacDonald (1998, 1999) and Mark and Sul (1998) have used panel estimators to test the monetary approach to exchange rate determination. All three papers report strong evidence in favor of the monetary approach, using panels of data from OECD countries and dollar, DM, and yen exchange rates.

This paper follows the lead of these earlier studies. We employ a panel of data from nine Asia/Pacific countries and test the monetary approach as a determinate of long-run equilibrium exchange rates. The evidence in favor of this hypothesis is strong. The results indicate that the collapse of some of these exchange rates in 1997 was not due to exchange rates being
out of line with fundamentals. In other words, the evidence suggests that these currencies were not “over-valued” so that their subsequent falls were not market corrections.

2. Exchange rate models

The monetary model of exchange rate behavior serves as the basic construct for equilibrium nominal exchange rate behavior in a variety of macroeconomic models. The model is a straightforward extension of absolute PPP:

\[ s_t = p_t - p_t^* \]  

where \( s_t \) denotes the spot exchange rate (home currency price of a unit of foreign exchange), \( p_t \) denotes a consumer price level and an asterisk denotes a foreign magnitude. Lower case letters in this and future expressions indicate that a variable is written in logs (the exception being interest rates). On using standard money market equilibrium conditions, the relationship of Eq. (1) may be transformed into a monetary approach equation. The money market equilibrium conditions for the home and foreign country are given by Eq. (2) and Eq. (3), respectively:

\[ m_t - p_t = \kappa + \alpha y_t - \beta i_t \]  
\[ m_t^* - p_t^* = \kappa^* + \alpha^* y_t^* - \beta^* i_t^* \]  

where \( m \) denotes the money supply, \( y \) is real income, \( i \) is the interest rate, \( \kappa \) is an arbitrary constant, \( \alpha \) is the income elasticity of demand for money and \( \beta \) is the interest rate semi-elasticity. On rearranging Eq. (2) and Eq. (3) for the home and foreign price levels and substituting into Eq. (1) a monetary approach equation may be derived which is the long-run solution for the flex-price monetary model of Bilson (1978), Frenkel (1978) and Hodrick (1978) and the sticky price monetary model of Dornbusch (1976) and Frankel (1979). This equation is reported here as Eq. (4):

\[ s_t = \psi + m_t - m_t^* - \alpha y_t - \beta i_t - \beta^* i_t^* \]  

Eq. (4) asserts that an exchange rate in equilibrium is driven by relative excess money supplies. This equation may be re-expressed as a reduced form, as:

\[ s_t = \omega_0 + \omega_1 y_t + \omega_2 m_t^* + \omega_3 y_t^* + \omega_4 i_t^* + \omega_5 i_t + \omega_6 i_t^* + u_t \]  

where the \( \omega \)s are the parameters to be estimated and \( u_t \) is a disturbance term. The strict form of the monetary model implies \( \omega_1 = -\omega_2 = 1, \omega_3 \) and \( \omega_4 \) should take on values which are close to estimated income elasticities from money demand functions and \( \omega_5 \) and \( \omega_6 \) should take on values close to interest rate semi-elasticities from the demand for money. The hypothesized values of \( \omega_1 \) and \( \omega_2 \) would be those expected, in equilibrium at least, by proponents of both the flex-price and fix-price monetary models. However, it is important to note that even within the monetary tradition there is a view which indicates that these coefficients may be less than one in absolute value; this view is attributed to Mussa (1976).
Glassman (1987) all estimate Eq. (5), or simplified versions thereof, using the Engle-Granger two-step methods and find no evidence of a long-run relationship in the sense that the residual series is an I(1) process. However, subsequent to these studies, MacDonald and Taylor (1993, 1994) have demonstrated that using the multivariate methods of Johansen (1988) a long-run version of the pure monetary model Eq. (5) can be defined for the recent float in the sense that the residuals are I(0) and point estimates are close to their a priori values. However, the speed of adjustment in such relationships proves to be painfully slow.

One feature of the PPP hypothesis, and therefore the monetary approach, is that the real exchange rate is assumed to be stationary. Several authors have recently noted that this assumption may be violated in certain countries, especially in Asia (see Isard & Symansky, 1996; and Chinn, 1997, 1998). To deal with this problem, Chinn (1998) suggests adding an additional variable to the empirical specification. This variable allows for movements in relative prices of tradables to nontradables within and across countries that may either be due to relative differences in productivity in tradable and nontradable sectors (see Balassa, 1964; and Samuelson, 1964) or international differences in preferences for nontradables (see DeGregorio & Wolf, 1994).

Since nontradable prices are not directly observable, it is necessary to construct an empirical proxy for the relative price term. Chinn (1998) suggests using

$$rp_t = \log \left( \frac{(PPI/CPI)}{(PPI^*/CPI^*)} \right),$$

where PPI is the producer price index, assumed to be comprised largely of tradables, and CPI is the consumer price index, assumed to include many nontradable goods prices. Chinn argues that the sign of this term should be positive.

### 3. Econometric methods

The approach adopted in this paper to estimate Eq. (5) involves exploiting panel data sets defined with respect to the Japanese yen. The use of panel data in the current context has a key advantage over single equation estimates. It is now widely accepted that low frequency data, such as annual data, have a much higher signal-to-noise ratio than monthly or quarterly data in the context of estimating an equilibrium relationship (Shiller & Perron, 1985). However, in trying to estimate, say, a PPP relationship for the recent floating period with annual data, a researcher will have only approximately 23 useable observations. Although this may be sufficient to obtain homogeneity, it is unlikely that it will produce evidence of a cointegrating relationship (i.e., a stationary residual series). A way to increase the power of such tests involves increasing the span of the data (see MacDonald, 1995). One possibility simply involves lengthening the historical time series of the data (see Diebold, Husted, & Rush, 1991). However, such an approach is fraught with difficulty—there may be several exchange rate regime changes throughout the period and the definition of the explanatory variables may have changed beyond all recognition from the beginning to the end of the sample. An alternative way to increase the span of the data involves expanding the panel dimensions. In a univariate context, Levin and Lin (1993) have demonstrated that a sub-
stantial increase in power may be obtained by testing for unit roots in a panel context. Pedroni (1995) has made a similar case for estimating cointegrating relationships in a panel context. We therefore propose using panel methods to estimate the monetary model of Eq. (5).

Our testing methods may be illustrated by starting with the following standard panel framework:

\[ s_{it} = \alpha + \beta' x_{it} + \{ \Sigma_i \gamma_i D_i \} + \{ \Sigma_t \delta_t D_t \} + u_{it} \quad (7) \]

where the \( i \) subscript indicates that the data have a cross sectional dimension (running from 1 to N), \( x_{it} \) is a matrix, consisting of any of the sets of explanatory variables from the models noted above, \( D_i \) and \( D_t \) denote, respectively, country-specific and time-specific fixed effect dummy variables (although not noted here, it is straightforward to incorporate random effects into Eq. (7)). In a standard panel setting several modeling strategies are available for the disturbance term: it may be assumed random, heteroscedastic, autoregressive (with either a common autoregressive terms across individual panel members or different autoregressive terms across members), it may be spatially correlated or some combination of these assumptions may be used.

However, the key feature of recent work with panel data sets relates to the potential non-stationarity of the data. If the dependent and explanatory variables in Eq. (7) are stationary then Eq. (7) would represent an appropriate way to estimate the desired relationship. If, however, the variables in Eq. (7) are non-stationary then it has to be determined if they form a cointegrating set or not. If they are I(1) and not cointegrated, an appropriate way to proceed simply involves recasting Eq. (7) with the dependent and independent variables in first differences. If, however, \( s_{it} \) and \( x_{it} \) are cointegrated then simply specifying Eq. (7) in difference form would be inappropriate since information pertaining to the long-run relationship would be lost.

In the single equation setting, testing for stationarity and determining the existence of cointegration amongst a vector of non-stationary variables is relatively straightforward since there are a number of alternative estimators available (see Campbell & Perron, 1991) for a discussion of the relative merits of a range of different estimators). Unfortunately, cointegration in the panel context is not well developed. Several approaches to this problem have been attempted in the recent literature. Husted and MacDonald (1998) employ a two-step approach wherein they use a variety of panel estimators to estimate a cointegrating vector and then test for the stationarity of the residuals from these specifications using the Levin-Lin panel unit root test. In so doing, they acknowledge the problem that they are using generated residuals in their second stage so that the conventional critical values for the Levin-Lin procedure do not apply. Consequently, they also supplement their results with alternative specifications such as error correction models that do not employ non-stationary data in the setup. Pedroni (1997) has proposed a test for panel cointegration that is somewhat similar to the procedure described above. However, his test is difficult to implement empirically, especially in the case of a common cointegrating vector. Finally, Mark and Sul (1998) also propose a two-step procedure that uses the panel unit root tests of Im, Pesaran, and Shin (1997) in the second stage.
The Levin-Lin method involves assuming that the $x_{it}$ matrix in Eq. (7) is simply equal to a scaler, defined as the lagged dependent variable. The stationarity of the exchange rate may be gauged by the estimated $t$-ratio on the lagged dependent variable. In circumstances where all of the deterministic elements in Eq. (7) are excluded apart from the single constant term, $\alpha$, Levin and Lin demonstrate that the $t$-statistic converges to a standard normal distribution. Including individual-specific effects (either or both) but excluding time-specific intercepts, Levin and Lin (1992) show that the $t$-ratio converges to a non-central normal distribution, with substantial impact on the size of the unit root test. However, Levin and Lin (1993) argue that unless there are very strong grounds for exclusion, time specific intercepts should always be included in this kind of panel test. The reason for this is that the inclusion of such dummies is equivalent to subtracting the cross section average in each period. This subtraction may be dispensed with in cases where the units in the panel are independent of each other, however in the case of panels involving bilateral exchange rates this would clearly be inappropriate.\footnote{7}

In testing for cointegration, we apply Levin-Lin tests to the residuals from Eq. (5). In particular, we report adjusted (ALL) and unadjusted (LL) Levin-Lin statistics, both constructed on the basis of no deterministic elements and a country-specific constant. The use of Levin-Lin $t$-ratios, of course, leaves the results reported below open to the generated residual criticism (although it should be noted that Frankel and Rose (1996) in their panel tests of PPP use a more primitive version of the Levin-Lin $t$-ratio, which does not adjust the data for cross sectional means). To address the generated residual issue, we follow Husted and MacDonald (1998) by using a method proposed by Hansen (1990). In that paper, Hansen describes what he refers to as the curse of dimensionality which is that in a residual-based cointegration test the cointegrating vector is random, imparting randomness into the generated residual series, and that this randomness is a function of the number of variables in the system. The particular implication that this has is to give dramatic reductions in power as the number of variables in the system increases, rendering cointegration tests ineffective in moderate sample sizes with moderately large systems. The reason that the test distributions depend on the dimensionality of the system is because they involve estimation of all $n$ unit roots in the system. Hansen’s proposition is to first adjust the non-stationary variables using a Cochrane-Orcutt estimator, and his Theorem 1 demonstrates that under the null hypothesis of no cointegration, the Cochrane-Orcutt estimate of the cointegrating coefficient converges to a constant rather than a random variable as it does using least squares. This suggests, therefore, that a test for cointegration constructed from the Cochrane-Orcutt estimates will not display the curse of dimensionality. In particular, Hansen’s test statistics have the advantage that their limiting distributions are always the same as those of univariate unit root tests. Since almost all of the traditional panel estimators implemented in this paper use a Cochrane-Orcutt type correction this should help provide confidence in the results of the panel cointegration tests. More generally, our approach involves using several panel estimators to act as a robustness check on the simplest model. Since one of the estimators involves a correction for contemporaneous correlation, it directly addresses the point made by O’Connell (1996) that a failure to control for cross sectional dependence in a panel unit root test can have a dramatic effect on statistical size.
4. Data and empirical results

The data for this study are collected from the International Monetary Fund’s *International Financial Statistics* CD-ROM disc and are for the sample period 1974–1996 for nine Asian countries. The exchange rate is defined as the home currency price the Japanese yen (line ae for the individual countries divided by line ae for Japan), money supplies are M1 (line 34), income terms are real GDP (line 99b.p), and interest rates are short rates (line 60 or 61). All variables, apart from the interest rate terms, have been converted into natural logarithms. The nine countries that make up the panel are Australia, India, Indonesia, Korea, Malaysia, New Zealand, Philippines, Singapore, and Thailand.

### 4.1. Single country two-step cointegration results

As a motivation for the panel results, we first report, in Table 1, individual country estimates of cointegrating vectors for the monetary model. Slightly more than half of the coefficients are correctly signed. The variable that performs least well is the home interest rate, with only two of 18 coefficients correctly signed. Foreign (i.e., Japanese) income has the most correctly signed coefficients (15 of 18). Money supplies have correctly signed coefficients in half the cases. Evidence against the no-cointegration is limited. ADF and PP t-ratios are large in absolute value. But, few are smaller than the 5% critical value (of \(-5.53\), which is based on six right hand side variables. Critical values for larger specifications, such as the equations that include \(r_p\) are not available). Overall the model does not perform as well as PPP. This should not be surprising. The monetary model requires not only that PPP holds, but also that price levels are determined solely by excess supplies of money. Both of these

<table>
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<th>Country</th>
<th>M</th>
<th>M*</th>
<th>Y</th>
<th>Y*</th>
<th>i</th>
<th>i*</th>
<th>RP</th>
<th>R^2</th>
<th>ADF</th>
<th>PP</th>
<th>k</th>
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<td>-4.38</td>
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propositions seem sensible as long run constructs, but may not occur with sufficient frequency to be detected in relatively short spans of data. This reasoning provides the motivation for the panel estimates that follow.

4.2. Panel estimates

Panel estimates of Eq. (5) were obtained using a variety of techniques and are reported in Table 2. Before discussing these estimates, we provide some detail on the estimators used in this section. They are: OLS with fixed effects, that is individual intercepts for each unit (denoted in the Table by $g_{0i}$); Maximum Likelihood estimates with fixed effects and a common AR1 process specified for the error term (this is denoted in the Table as $\text{MAX}_1(r_i = r_j)$); and Maximum Likelihood with fixed effects and an AR1 process which differs across individual units (represented in the Table by $\text{MAX}_2(r_i \neq r_j)$); finally, we report a Maximum Likelihood estimator which allows fixed effects, differing AR1 processes across individual units, contemporaneous residual correlation and a correction for heteroskedasticity (this is labeled ‘POOL’ in the Table). Each of these estimators is used over the entire panel of data, and slope coefficients are assumed to be equal across countries.

The above-noted ‘traditional’ panel models facilitate testing for cointegration in the various nominal models. As an alternative check on the existence of cointegration, we also provide estimates from two error correction models (ECMs) in a panel context. The ECMs are constructed with the current change in the exchange rate on the left-hand-side and two lagged changes of all the variables in the model as well as the lagged equilibrium error on the right. The equilibrium error is the residual taken from a first-stage regression based on OLS fixed-effects estimates of Eq. (5). The table reports OLS estimates of the ECM model (labeled ‘ECM (OLS)’) and full pooling estimates using the Pool estimator referred to above (ECM(POOL)). The focus of interest in ECM models is the coefficient on the lagged equilibrium error, reported in the table under the heading $\mu$. The value of $\mu$ provides information on whether or not cointegration exists and, if so, the speed at which the exchange

<table>
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<th>Method</th>
<th>M</th>
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<th>Y</th>
<th>Y*</th>
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<td>ECM (POOL)</td>
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<td>-0.16</td>
<td>.30</td>
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<td>(0.06)</td>
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rate responds to disturbances in the equilibrium relationship. Consequently, we report both the coefficient estimate and in parentheses below its estimated standard error.

4.3. Panel monetary approach regressions

Table 2 reports estimates of the monetary approach model, Eq. (5), using several alternative panel estimators. The results in support of the monetary model are striking. First, a very high proportion of the coefficient estimates have the correct sign, and many of these parameters have magnitudes consistent with the standard monetary model. In particular, the coefficients on home money, and home and foreign output are all correctly signed. The fixed effects OLS estimates are all correctly signed, and the fit is very good. However, simple GLS corrections for residual autocorrelation (i.e., MAX1 and MAX2) lead to sign changes on two variables. When these corrections are accompanied by GLS corrections for heteroskedasticity and contemporaneous correlation with the POOL estimator, the hypothesized signs reemerge.

As before, the LL and AL statistics provide information on whether or not these are long-run relationships in the data. And again, with limited exceptions, the statistics are much lower than the critical value of $-1.96$. This suggests relatively strong evidence of cointegration, most especially with the POOL estimates. The error correction results reinforce this conclusion. The estimates of $\mu$ are statistically significant, although relatively small. They suggest that the half-life of a shock is 4.6 years. This evidence provides further support for the use of a panel of data rather than data from a single country to test the monetary model.

In Figs. 1 through 5 we present an ocular impression of a representative set of equilibrium
and actual exchange rates. In particular we plot the equilibrium, or fitted, values derived from the OLS fixed effects panel estimates of the monetary model, for five of the currencies most closely associated with the Asian crises.¹⁰ We believe these figures are interesting. For three of the currencies – the Korean won, Thai baht and Phillipine peso (Figs. 1, 2 and 3, respectively), we see:

- Fig. 1. Korean won
- Fig. 2. Thai baht
- Fig. 3. Philippines peso
respectively) – there is a tight lock between actual and fitted values just before the crises occurred suggesting these currencies were not misaligned. Indeed, there is a remarkably close correspondence in the actual and equilibrium values for these two currencies throughout the sample. For the Malysian ringgit and Indonsian rupiah (Figs. 4 and 5) there is more
evidence of misalignment at the end of 1996. The former currency appears overvalued in the 1990’s while the rupiah seems persistently undervalued from the mid-1980s onwards, there being a very striking undervaluation just before the crash.

5. Summary and conclusions

This paper has sought to provide information on the long-run behavior of Asia-Pacific nominal exchange rates vis-a-vis the yen over the post Bretton Woods period. In particular, we report tests of the monetary approach to exchange rates over the period 1974–1996, individually for nine countries and for these countries as a panel. Evidence in favor of the monetary model emerges strongly from the panel estimates, but not in the individual country estimates. Again, a cointegrating relationship between exchange rates and money market fundamentals is evident in the panel tests. To supplement these tests, we also report estimates from error correction models. This exercise suggests that the half-life to an exchange rate shock may be as long as 4.5 years.

The results in this paper argue that sensible statements can be made about equilibrium values of Asian/yen exchange rates. Is there any evidence from this analysis that actual exchange rates at the end of 1996 were substantially overvalued? To address that question, we compare actual and estimated values from the panel OLS monetary model. For most of the countries, the model does a very good job of replicating actual behavior. In virtually all cases, the model fits the data more closely at the end of the period than in the beginning. In particular, for many of the countries most closely associated with the Asian crisis, Korea, Malaysia, the Philippines, and Thailand, the model works especially well. With the exception of Malaysia, the model suggests that equilibrium exchange rates at the end of 1996 were very close to actual values. In Malaysia, the ringgit appreciated strongly in 1996, especially when compared to long run equilibrium. For Indonesia, the model suggests that the rupiah has been undervalued vis-a-vis the yen for most of the past ten years. These findings suggest that the currency falls experienced by these countries represent shifts in long run mean values unrelated to underlying fundamentals.

Notes

1. See, for example, Radelet and Sachs (1998a) and the references therein.
3. See Frankel and Rose (1996), and MacDonald (1988).
5. These results are identical to those obtained with long runs of historical data.
7. Subtracting the cross section mean reduces the potentially problematic effect that contemporaneous correlations can have in a panel unit root context; see O’Connell (1996).
8. Several interest rate values were missing from the data. Simple interpolations were used to construct these values.
9. Data problems for both the People’s Republic of China and Taiwan led us to exclude them from the panel.
10. Residuals from the other estimated models reported in the paper represent errors from generalized transformations of the raw data.

References


