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A grayscale photograph of the Bank of Canada building, a large modern structure with a prominent classical-style portico and a flagpole on the roof. The building is surrounded by trees and a street with a lamp post.

# **An Empirical Analysis of Foreign Exchange Reserves in Emerging Asia**

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**Marc-André Gosselin and Nicolas Parent**

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The views expressed in this paper are those of the authors.  
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## Abstract

Over the past few years, the ability of the United States to finance its current account deficit has been facilitated by massive purchases of U.S. Treasury bonds and agency securities by Asian central banks. In this process, Asian central banks have accumulated large stockpiles of U.S.-dollar foreign exchange reserves. How far is the current level of reserves from that predicted by the standard macroeconomic determinants? The authors answer this question by using Pedroni's (1999) panel cointegration tests as the basis for the estimation of a long-run reserve-demand function in a panel of eight Asian emerging-market economies. This is a key innovation relative to the existing research on international reserves modelling: although the data are typically I(1), the literature ignores this fact and makes statistical inference based on unadjusted standard errors. While the authors find evidence of a positive structural break in the demand for international reserves by Asian central banks in the aftermath of the financial crisis of 1997–98, their results indicate that the actual level of reserves accumulated in 2003–04 was still in excess relative to that predicted by the model. Therefore, as long as historical relationships hold, a slowdown in the rate of accumulation of reserves is likely. This poses negative risks for the U.S. dollar. However, both the substantial capital losses that Asian central banks would incur if they were to drastically change their holding policy and the evidence that the currency composition of reserves evolves only gradually mitigate the risks of a **rapid** depreciation of the U.S. dollar triggered by Asian central banks.

*JEL classification: C23, F31, G15*

*Bank classification: Econometric and statistical methods; International topics; Financial stability*

## Résumé

Ces dernières années, la capacité des États-Unis à financer le déficit de leur balance des paiements courants a été favorisée par les achats massifs d'obligations du Trésor américain et de titres d'agences américaines par les banques centrales asiatiques. Celles-ci ont ainsi amassé d'énormes réserves de dollars É.-U. Dans quelle mesure le niveau actuel de leurs réserves de change diffère-t-il de celui que justifient les déterminants macroéconomiques habituels? Les auteurs répondent à cette question en recourant aux tests de cointégration sur données de panel proposés par Pedroni (1999) pour estimer la fonction de demande à long terme de réserves d'un groupe de huit économies émergentes d'Asie. Leur démarche novatrice se distingue de l'approche privilégiée jusqu'ici dans les travaux consacrés à la modélisation des réserves de change, où les auteurs tirent leurs inférences statistiques sans tenir compte du fait que les données sont généralement de type I(1) et, donc, sans corriger les écarts-types. Bien que l'étude montre qu'une

rupture structurelle positive soit survenue dans la demande de réserves internationales émanant des banques centrales asiatiques au lendemain de la crise financière de 1997-1998, il reste que le niveau des réserves en 2003 et en 2004 demeure supérieur aux projections du modèle utilisé. En conséquence, dans la mesure où les relations déduites des données historiques sont toujours valables, on peut s'attendre à un ralentissement du rythme d'accumulation des réserves. Cette évolution serait défavorable au dollar américain. Néanmoins, comme une révision radicale de la politique que les banques centrales asiatiques suivent en matière de réserves leur ferait subir de lourdes pertes en capital et que la composition en devises des réserves a tendance à ne se modifier que graduellement, le risque qu'elles déclenchent une dépréciation **rapide** du dollar américain est limité.

*Classification JEL : C23, F31, G15*

*Classification de la Banque : Méthodes économétriques et statistiques; Questions internationales; Stabilité financière*

## 1. Introduction

Over the past few years, the ability of the United States to finance its current account deficit has been facilitated by massive purchases of U.S. Treasury bonds and agency securities by Asian central banks. In this process, Asian central banks have accumulated large stockpiles of U.S.-dollar foreign exchange reserves.

Central banks cannot accumulate reserves indefinitely. Excessive reserve hoarding entails significant sterilization costs, since the negative spread between the interest earned on reserves and the interest paid on the country's public debt increases with reserve accumulation. Moreover, if capital flows are not sterilized, sustained reserve accumulation will, at some point, generate inflationary pressures that could increase the risk of domestic financial crises. On the other hand, if these central banks decide to stop accumulating U.S.-dollar reserves, they could trigger an abrupt depreciation of the U.S. dollar and a sharp rise in interest rates.<sup>1</sup> Given the potential impact on global interest rates, growth, and financial stability, the issue of Asian reserve accumulation is of considerable importance.

How far is the current level of reserves from that given by the standard macroeconomic determinants? In this paper, we answer this question by using Pedroni's (1999) panel cointegration tests as the basis for the estimation of a long-run reserve-demand function in a panel of eight Asian emerging-market economies: China, India, Indonesia, Korea, Malaysia, the Philippines, Singapore, and Thailand. This is a significant econometric improvement relative to the existing research on international reserves modelling: although the data are typically I(1), the literature ignores this fact and makes statistical inference based on unadjusted standard errors.

In line with the literature, we find that the level of reserve holdings can be explained by a few key macroeconomic factors. However, while our model accounts for a positive

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<sup>1</sup> For instance, Warnock and Warnock (2005) estimate that, had foreign official flows been zero over the past twelve months, long rates would currently be 60 basis points higher in the United States.



structural break in the demand for international reserves by Asian central banks in the aftermath of the financial crisis of 1997–98, it cannot explain the large accumulation of international reserves by these institutions in 2003–04.

This result suggests that, *ceteris paribus*, a slowdown in the speed of accumulation of reserves is likely (even if exchange rate policies in this area remain unchanged). Taken alone, this factor implies potential downward pressures on the U.S. dollar. However, the risks of large capital losses on Asian central banks' balance sheets mitigate somewhat the risks of a **rapid** depreciation of the U.S. dollar triggered by Asian central banks.

This paper is organized as follows. In section 2, we review some key stylized facts regarding the accumulation of international reserves in emerging Asia. In section 3, we discuss recent findings from the empirical literature on foreign exchange reserves. In section 4, we present our empirical model and results. In section 5, we discuss our findings. Section 6 offers some conclusions.

## **2. Stylized Facts**

Global reserve purchases topped \$441 billion in 2003. According to Bank for International Settlements (BIS) estimates, these purchases financed 83 per cent of the U.S. current account deficit in that year, with private investors funding the remainder (Higgins and Klitgaard 2004). At the end of 2003, central bank holdings of dollar assets were equivalent to more than half of marketable U.S. Treasury debt outstanding. As Figure A1 in Appendix A shows, developing Asia almost doubled its level of reserves between 1999 and 2003, holding more reserves in 2003 than all developed countries taken together. According to the International Monetary Fund (IMF), of the roughly \$1.2 trillion increase in global reserves from the end of 1999 to the end of 2003, \$582 billion reflects purchases by developing countries in Asia. The share of global reserves held by emerging-market countries rose from 37 per cent in 1990 to 61 per cent in 2002, with emerging Asia accounting for much of the increase.

Figure A2 shows international reserves deflated by GDP in selected Asian emerging-market economies. Reserves show an upward trend even after economic size is taken into account. As the empirical analysis will show, increasing openness to international trade is a key factor behind this tendency. Reserves increased significantly in economies with limited exchange rate flexibility and those with managed floating exchange rates.<sup>2</sup> In absolute terms, the increase is most significant in China and Korea; between 1980 and 2003, reserves rose by \$405 billion for China and \$152 billion for Korea. Within emerging Asia, the share of international reserves held by China and Korea rose sharply from 8 to 45 per cent and from 9 to 17 per cent, respectively (Figure A3).

Several indicators have been used to evaluate whether reserve holdings are sufficient.<sup>3</sup> Scaled against short-term external debt, output, or imports, international reserves in emerging Asia have increased markedly (Table 1).

**Table 1: International Reserves in Emerging Asia**

| Reserves                      | 1980 | 1990 | 2000 | 2001 | 2002 | 2003 |
|-------------------------------|------|------|------|------|------|------|
| Total minus gold <sup>a</sup> | 33   | 129  | 593  | 654  | 790  | 1010 |
| Scaled by:                    |      |      |      |      |      |      |
| Short-term debt               |      | 0.7  | 2.9  | 3.3  | 4.2  | 4.5  |
| GDP                           | 4%   | 10%  | 22%  | 24%  | 27%  | 31%  |
| Imports <sup>b</sup>          | 2    | 4    | 6    | 7    | 8    | 9    |

a. In billions of U.S. dollars.

b. In months of import coverage.

The ratio of reserves to short-term external debt measures the capacity of a country to service its external liabilities in the forthcoming year, should external financing conditions deteriorate sharply. According to the Greenspan-Guidotti rule, a ratio above one signals that a country holds an adequate level of reserves to face the risk of a financial crisis, while a ratio below one may suggest a vulnerable capital account.<sup>4</sup> The

<sup>2</sup> See Table A1 in Appendix A for exchange rate regime classifications in emerging Asia.

<sup>3</sup> See Bird and Rajan (2003) for a discussion.

<sup>4</sup> See Greenspan (1999) and BIS (2000).

rationale is that, if reserves exceed short-term debt, then a country can be expected to meet its obligations in the coming year and thus avoid rollover problems stemming from liquidity concerns. Figure A4 shows that in Indonesia, Korea, the Philippines, Singapore, and Thailand the ratio of reserves to short-term external debt was either around one or significantly below one before 1997. By 2003, however, with the exception of the Philippines, where the ratio remained fairly stable throughout the sample, all countries experienced an improvement in their capacity to face short-term liabilities. China and India display the highest ratio in 2003, with, respectively, 10.2 and 5.4, up from 3.6 and 2.6 in 1997.

The ratio of reserves to imports is considered as a proxy for a country's current account vulnerability. The ratio measures the number of months a country is able to finance its current level of imports. Normally, a ratio of 3 and 4 would be considered adequate (Fisher 2001). The current ratio of reserves to imports shows that emerging Asia is able to finance nine months of imports, pointing again to substantial reserve accumulation. When examining similar ratios, Mendoza (2004) concludes that reserve management in many countries in emerging Asia is motivated by a desire to self-insure against a financial crisis.

Although these ratios provide a good measure of reserve adequacy in terms of a country's resiliency when facing a potential financial crisis, they do not provide an upper bound for reserve holdings. Nevertheless, such high ratios suggest that the reserve buildup in emerging Asia may have reached a point where some slowdown in the rate of accumulation could be warranted. Is it possible to explain the current level of foreign exchange reserves with standard macroeconomic determinants? The remainder of this paper examines this issue.

### **3. Review of the Empirical Literature**

Emerging-market economies hold reserves as a buffer stock to smooth unexpected and temporary imbalances in international payments. In determining the optimal level of reserves, the monetary authority will seek to balance the macroeconomic adjustment

costs incurred if reserves are exhausted (crisis-prevention motive) with the opportunity cost of holding reserves (Heller 1966).<sup>5</sup> In theory, a country can decide to accumulate foreign exchange reserves to eliminate all or some of its consumption volatility. In this case, the level of reserves will increase with a country's risk aversion and output volatility.

Empirical research on international reserves (Heller and Khan 1978; Edwards 1985; Lizondo and Mathieson 1987; Landell-Mills 1989; and Lane and Burke 2001) establishes a relatively stable long-run demand for reserves based on a limited set of explanatory variables. The determinants of reserve holdings reported in the literature can be grouped into five categories: economic size, current account vulnerability, capital account vulnerability, exchange rate flexibility, and opportunity cost. Table 2 lists potential explanatory variables for each of these categories.

**Table 2: Empirical Determinants of Reserve Holdings**

| Determinants                  | Explanatory Variables   |
|-------------------------------|---|
| Economic size                 | GDP, GDP per capita   |
| Current account vulnerability | Share of imports or exports in output, volatility of export receipts  |
| Capital account vulnerability | Financial openness: ratio of capital flows or broad money to GDP, short-term external debt, foreigners' equity position |
| Exchange rate flexibility     | Volatility of the exchange rate   |
| Opportunity cost              | Interest rate differentials   |

In theory, the volume of international financial transactions, and therefore reserve holdings, should increase with economic size. In the literature, GDP and GDP per capita are used as indicators of economic size. The vulnerability of the current account can be captured by such measures as trade openness and export volatility. In the long run, central banks will increase their reserves in response to a greater exposure to external shocks.

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<sup>5</sup> There is considerable evidence in the literature on early warning systems that higher reserves reduce both the likelihood of a crisis and its depth if it does occur (Berg and Patillo 1999). Note that, if macroeconomic policies are not sustainable, reserve holdings may only postpone the inevitable crisis (Krugman 1979).

For this reason, the level of reserves should be positively correlated with an increase in both exports and imports.<sup>6</sup> Capital account vulnerability increases with financial openness and potential for resident-based capital flight from the domestic currency. Consequently, reserves should be positively correlated with such variables as the ratio of capital flows to GDP and the ratio of broad money to GDP (which signals the potential demand for foreign assets from domestic sources). Exchange rate flexibility is usually important: it reduces the demand for reserves, since central banks no longer need a large stockpile of reserves to manage a pegged exchange rate. Because there is a “fear of floating,” flexibility is generally measured by the actual volatility of the exchange rate. There is an opportunity cost of holding reserves, because the monetary authority swaps high-yield domestic assets for low-yield foreign ones. It corresponds to the difference between the yield on reserves and the marginal productivity of an alternative investment. This variable is, however, often insignificant in the empirical literature, likely reflecting measurement problems (Edwards 1985).

The IMF (2003) recently studied a simple empirical model that incorporates various determinants of reserve holdings. The model is estimated using a large panel that covers 122 emerging-market economies with annual data from 1980 to 1996. In the study, real GDP per capita, the population level, the ratio of imports to GDP, and the volatility of the exchange rate are found to be statistically significant determinants of real reserves.<sup>7</sup> Predicted values from this model over the 1997–2002 period reveal that international reserves in Latin America are not excessive, while those in emerging Asia have increased more than warranted by the determinants since 2001. The IMF concludes that foreign exchange reserves in emerging Asia have reached a point where some slowdown in the rate of accumulation is needed.

As Mendoza (2004) shows, it is reasonable to assume that most Asian countries increased their level of reserves for self-insurance purposes in the aftermath of the Asian financial

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<sup>6</sup> In the shorter run, however, foreign exchange intervention may be required to maintain a currency peg so that reserves are positively correlated with the trade balance (i.e., positively correlated with exports but negatively correlated with imports).

<sup>7</sup> Measures of capital account vulnerability and opportunity cost were insignificant.

crisis.<sup>8</sup> Similarly, Lizondo and Mathieson (1987) find that the debt crisis of the early 1980s in Latin America produced a structural break in the demand for reserves. Given that the Asian crisis is likely to have led to significant changes in the relationship between variables, a question that naturally stems from the IMF study is that of structural stability. The IMF's predictions for the 1997–2002 period could be questionable, given that they are based on parameter estimates from the 1980–96 sample. When we run various robustness tests using the IMF's data set, however, we do not find any statistical evidence of a break in the patterns of the correlations among the variables.<sup>9</sup> This could reflect the fact that the IMF's data set covers a wide array of monetary regimes for which the average coefficients are stable.

Still, the idea that there have been structural breaks in the demand for international reserves for some countries in emerging Asia is confirmed by Aizenman, Lee, and Rhee (2004). Using data for Korea, they find evidence of a break in the pattern of hoarding of international reserves in the post-1997 period. The authors claim that the self-insurance motive became stronger following the crisis. More specifically, they find that trade openness is significant in explaining international reserves before the crisis, but that it loses significance after the crisis. They argue that this is consistent with the increased relative importance of financial openness. To examine whether increased external financial exposure is a driving factor behind reserve buildup in the post-crisis period, they consider foreigners' equity positions and short-term external debts as additional explanatory variables. They find that coefficients on these variables become significant after the crisis, supporting the view that Korea raised its level of reserves to increase its insurance against sudden stops of capital flows.

Aizenman and Marion (2002, 2004) investigate the interpretations of the relatively high demand for reserves by countries in emerging Asia and the relatively low demand by some other developing countries (e.g., Latin America). In addition to the variables listed

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<sup>8</sup> Lee (2005) develops a quantitative framework to bring forward the insurance motive for holding international reserves. In his model, the self-insurance value of reserves can be approximated by the cost of obtaining an equivalent insurance in the market.

<sup>9</sup> When estimating the IMF's model through 2002, we do not find any significant time effects in the intercept or slope coefficients.

in Table 2, they examine the role of political uncertainty and corruption as determinants of reserve holdings.<sup>10</sup> Using a theoretical model, they show that sovereign risk, costly tax collection to cover fiscal liabilities, and loss aversion (defined as the tendency of agents in an economy to be more sensitive to reductions in consumption than to increases) lead to a relatively large precautionary demand for international reserves. They further conclude that the recent large buildup of international reserves in emerging Asia is motivated by the experience of the Asian crisis.<sup>11</sup>

Another popular explanation for the high level of reserves is export competitiveness as a development strategy. Dooley, Folkerts-Landau, and Garber (2004) argue that reserve accumulation reflects the intervention of Asian central banks who want to prevent their currency from appreciating against the U.S. dollar in order to promote export-led growth. However, using lagged export growth and deviations from predicted purchasing-power parity in addition to the standard determinants, Aizenman and Lee (2005) find limited support for the mercantilist motive. Rather, their overall results are in line with the precautionary demand, even for China.

#### **4. Empirical Analysis**

The empirical literature focuses on examining equations for foreign exchange reserves from either very large panels or single countries. We estimate a long-run equation that applies to emerging Asia as a whole. Our sample covers China, India, Indonesia, Korea, Malaysia, the Philippines, Singapore, and Thailand.<sup>12</sup> Focusing on these countries is of greater relevance to the issue of global reserves accumulation, since the recent buildup in reserves is concentrated in that part of the world. This focus also addresses the question of reserve accumulation more thoroughly than by studying a sample of countries with

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<sup>10</sup> Political uncertainty is measured by the estimated probabilities of leadership change from a multinomial logit. The political corruption index is from the International Country Risk Guide.

<sup>11</sup> The fact that some countries will choose to hold more reserves in the aftermath of a crisis could also be explained by starting-point differences. The low level of reserves in Latin America is consistent with a desire to reduce inflation in that region. In emerging Asia, inflation is already low, so the emphasis is on bringing output back to potential.

<sup>12</sup> We do not include Japan in the analysis, since studies generally find that the behaviour of emerging-market economies diverges from that of industrial economies, with external variability being a more important factor of reserve demand for the former. Hong Kong and Taiwan are excluded due to lack of data.

fundamentally different policy regimes, as the IMF (2003) does. Given that the Asian crisis appears to have led to structural breaks in the patterns of the correlations among the variables, we also allow for breaks in the intercept and slope coefficients in the post-1997 period. The data are annual and span from 1980 to 2003; the model includes country fixed effects. The estimated model is:

$$y_{i,t} = \mathbf{a} + \mathbf{d}_i + \sum_{k=1}^K \mathbf{b}_k x_{k,i,t} + e_{i,t}, \quad (1)$$

where  $y_{i,t}$  is the dependent variable,  $x_{i,t}$  contains the explanatory variables,  $k$  is the number of regressors,  $i$  the number of countries,  $t$  the number of time periods, and  $e_{i,t}$  a stationary disturbance term.

#### 4.1 Methodology

Although the data are typically I(1), the existing literature ignores this fact and makes statistical inference based on unadjusted standard errors. It is well known in time-series econometrics that  $t$ -statistics of spurious regressions are biased upwards. We formally address the issue of non-stationarity by using panel cointegration tests.

Kao (1999) and Pedroni (1999) derive residual-based panel cointegration tests. Kao's tests allow for no heterogeneity among the regressors, meaning that the covariance matrices must be the same across countries. Although Pedroni's tests relax this assumption, Gutierrez (2003) finds that, for panels with a short time dimension, Kao's tests have higher power than Pedroni's. Using the Newey-West (1987) method to estimate the asymptotic covariance matrices for each country, we compute an  $F$  test for the equality of covariance matrices across countries. For some countries, we reject the null hypothesis of equality of matrices. This justifies the use of Pedroni's tests.

Pedroni (1999) derives seven different panel cointegration statistics. Of these seven residual-based statistics, four are based on pooling along the within-dimension (panel  $v$ ,  $\rho$ , Phillips-Perron (PP), and augmented Dickey-Fuller (ADF) stats), and three are based



on pooling along the between-dimension (group  $\rho$ , PP, and ADF stats).<sup>13</sup> Appendix B reports the precise form for each of these seven statistics. In all cases, these statistics can be constructed using the residuals of the cointegrating regression (1) in combination with various nuisance parameter estimators that can be obtained from these ( $\hat{L}_{11i}^2, \tilde{\mathbf{S}}_{N,T}^2, \hat{\mathbf{I}}_i, \tilde{\mathbf{S}}_{N,T}^{*2}$ , and  $\hat{s}_i^{*2}$ ; see Appendix B). We compute these statistics by following the five steps suggested by Pedroni (1999):

1. Estimate the panel cointegration regression (i.e., equation (1)) and collect the residuals, denoted  $\hat{e}_{i,t}$ .
2. Difference the original series for each member ( $\Delta y_{i,t}, \Delta x_{i,t}$ ) and compute the residuals from the differenced regression, denoted  $\mathbf{h}_{i,t}$ .
3. Calculate the long-run variance of  $\mathbf{h}_{i,t}$  ( $\hat{L}_{11i}^2$ ) using the Newey-West (1987) estimator.
4. Using the residuals,  $\hat{e}_{i,t}$ , of the original cointegrating regression, estimate the appropriate autoregression choosing either of the following forms: (a) for the non-parametric statistics, estimate  $\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \hat{\mathbf{m}}_{i,t}$  and use the residuals to compute  $\hat{\mathbf{I}}_i$  and  $\hat{s}_i^2$ ; the individual and panel long-run variances of  $\hat{\mathbf{m}}_{i,t}$ , denoted  $\hat{\mathbf{S}}_i^2$  and  $\tilde{\mathbf{S}}_{N,T}^2$ , respectively, can then be computed. Or, (b) for the parametric statistics, estimate  $\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \sum_{k=1}^{K_i} \hat{\mathbf{g}}_{i,k} \Delta \hat{e}_{i,t-k} + \hat{\mathbf{m}}_{i,t}^*$  and use the residuals to compute the simple variance of  $\hat{\mathbf{m}}_{i,t}^*$ , denoted  $\hat{s}_i^{*2}$ , as well as the contemporaneous panel variance estimator, denoted  $\tilde{\mathbf{S}}_{N,T}^{*2}$ .

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<sup>13</sup> As Pedroni (1999, 657–58) notes: “A consequence of this distinction arises in terms of the autoregressive coefficients,  $\gamma_i$ , of the estimated residuals under the hypothesis of cointegration. For the within-dimension statistics the test for the null of no cointegration is implemented as a residual-based test of the null hypothesis  $H_0: \gamma_i=1$  for all  $i$ , versus the alternative hypothesis  $H_1: \gamma_i=\gamma < 1$  for all  $i$ , so that it presumes a common value for  $\gamma_i=\gamma$ . By contrast, for the between-dimension statistics the null of no cointegration is implemented as a residual-based test of the null hypothesis  $H_0: \gamma_i=1$  for all  $i$ , versus the alternative hypothesis  $H_1: \gamma_i < 1$  for all  $i$ , so that it does not presume a common value for  $\gamma_i=\gamma$  under the alternative hypothesis.”

5. Using each of these parts, construct the statistics from Appendix B and apply the appropriate mean and variance adjustment terms from Pedroni (1999, Table 2, p. 666) to standardize the tests to an  $N(0,1)$  distribution.<sup>14</sup>

#### 4.2 Results: cointegration analysis

Since only the  $I(1)$  variables can be considered as potential regressors in the cointegrating space, we first check the order of integration of the variables commonly used in the literature. To do so, we use the Im, Pesaran, and Shin (IPS 2003) panel unit root test. Based on the mean of the individual ADF statistics of each country in the panel, the IPS test assumes that all series are non-stationary under the null hypothesis. Lags of the dependent variable are introduced to allow for serial correlation in the errors.

The dependent variable is the log of total reserves minus gold divided by nominal GDP (*res*). As Figure A1 shows, controlling for economic size is not sufficient to remove the upward trend in reserves. One potential reason for this is increasing openness to trade, which renders the economy more vulnerable to external shocks, since imbalances in payments can be more substantial. As such, we consider import propensity (imports divided by GDP, *imp*) and the volatility of export receipts (10-year moving standard deviation,  $v(x)$ ) as explanatory variables that capture current account vulnerability.<sup>15</sup> In the capital account vulnerability category, we consider the ratio of short-term external debt to GDP (*debt*) and the ratio of broad money to GDP (*M2*). The other potential explanatory variables we consider include exchange rate volatility (12-month moving standard deviation of per cent change,  $v(er)$ ) and opportunity cost (interest rate differentials, *cost*). Table 3 reports the results for the IPS test.

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<sup>14</sup> Pedroni (2004) shows that, following an appropriate standardization, each of these statistics will be distributed as standard normal when both the time series and cross-sectional dimensions of the panel grow large.

<sup>15</sup> We use the same definition of the volatility of export receipts as the IMF (2003). This allows us to better isolate the effect of the different methodologies.

**Table 3: Im, Pesaran, and Shin Panel Unit Root Test<sup>a</sup>**

| Variables   | Number of lags | $t$ -bar test | Order of integration |
|-------------|----------------|---------------|----------------------|
| <i>res</i>  | 4              | -1.362        | I(1)                 |
| <i>imp</i>  | 0              | -1.483        | I(1)                 |
| $v(x)$      | 0              | -0.961        | I(1)                 |
| <i>M2</i>   | 4              | -1.144        | I(1)                 |
| $v(er)$     | 4              | -1.880        | I(0)                 |
| <i>cost</i> | 4              | -2.396        | I(0)                 |
| <i>debt</i> | 0              | -1.301        | I(1)                 |

a. The critical value at the 10 per cent level is -1.880.

We reject only the null hypothesis of non-stationarity for the exchange rate volatility and interest rate differentials. Therefore, these two variables will not be considered in the panel cointegration tests.

Since the  $t$ -statistics associated with equation (1) are not valid, we determine the specification of the equation for our long-run reserve demand using the criterion that all variables in the cointegrating space are necessary for cointegration in *all* seven of the tests at the 5 per cent level.<sup>16</sup> To select the model, we first test whether we reject the null of no cointegration when all I(1) variables from Table 3 are included in the cointegrating space (i.e., *res*, *imp*, *M2*,  $v(x)$ ).<sup>17</sup> We also take into account a potential structural break in the demand for reserves following the Asian financial crisis by allowing for breaks in the intercepts and slope coefficients (i.e., in the post-1997 period). If the null hypothesis of no cointegration is rejected, we alternatively remove variables and check whether cointegration is maintained. If cointegration is preserved, the variable is superfluous and is consequently dropped. We proceed with this strategy until we reach the most parsimonious cointegrating space. Table 4 reports results from the Pedroni (1999) tests based on various cointegrating spaces.

<sup>16</sup> Finding cointegration with all seven tests is a form of robustness check.

<sup>17</sup> We also try the ratio of short-term external debt to GDP as a measure of potential external drain (e.g., non-rollover of debt, capital reversal). This variable was never necessary for cointegration.

**Table 4: Panel Cointegration Tests (1980–2003)<sup>a</sup>**

| Cointegrating space                 | Panel<br>v-stat | Panel<br>ρ-stat | Panel<br>pp-stat | Panel<br>adf-stat | Group<br>ρ-stat | Group<br>pp-stat | Group<br>adf-stat |
|-------------------------------------|-----------------|-----------------|------------------|-------------------|-----------------|------------------|-------------------|
| 1. $res, imp^*, M2_{t-1}^*, v(x)^*$ | -2.346          | 3.396           | 2.870            | 3.001             | 4.484           | 3.856            | 4.108             |
| 2. $res, imp^*, M2_{t-1}^*, v(x)$   | -1.935          | 2.614           | 2.016            | 2.164             | 3.730           | 3.013            | 3.277             |
| 3. $res, imp^*, M2_{t-1}^*$         | -1.809          | 1.854           | 1.019            | 1.170             | 2.705           | 1.552            | 1.885             |
| 4. $res, imp^*, M2_{t-1}, v(x)$     | -1.302          | 1.730           | 0.891            | 1.044             | 2.810           | 1.729            | 1.999             |
| 5. $res, imp, M2_{t-1}^*, v(x)$     | -1.642          | 2.149           | 1.662            | 1.760             | 3.247           | 2.686            | 2.927             |

a. All reported values are distributed  $N(0,1)$  under the null hypothesis of no cointegration. Critical values at the 10, 5, and 1 per cent level are 1.645, 1.96, and 2.575, respectively.  $res$  is total reserves minus gold deflated by nominal GDP (in logs),  $imp$  is the ratio of imports to GDP (in logs),  $M2_{t-1}$  is the lagged ratio of broad money to GDP (in logs), and  $v(x)$  is the volatility of export receipts. “\*” means that a structural break in the slope coefficient is allowed. The underlying regression model also includes country fixed effects.

With specifications 1 and 2, we reject the null hypothesis of no cointegration at the 5 per cent level. The break in the coefficient of  $v(x)$  is therefore superfluous and this variable can be excluded from the cointegrating space. From specifications 3, 4, and 5, we note that the other variables and the breaks are not redundant, since removing any of them reverses the conclusion of the test. Structural breaks in the coefficients of imports to GDP and money to GDP are therefore necessary to reject the null hypothesis of no cointegration. This supports the hypothesis of post-crisis break and points to evidence of a change in central bank behaviour.

Thus, the most parsimonious specifications yielding cointegration include the ratio of imports to GDP, the ratio of broad money to GDP (lagged, due to potential endogeneity), the volatility of export receipts, and breaks in the coefficients of imports to GDP as well as the ratio of broad money to GDP in the post-crisis period (specification 2). The fact that each of these variables is required to reject the null hypothesis of no cointegration is evidence that the regressors are statistically significant long-run determinants of reserve holdings.<sup>18</sup> From this, we obtain the following ordinary least squares estimates ( $d_{1997}$  is a dummy variable that is equal to 1 after 1997 and 0 otherwise):

<sup>18</sup> We find that, in addition to being stationary, the residuals from this specification, except for those from India, are all normally distributed according to the Shapiro-Wilk and Shapiro-Francia tests for normality.

$$res_{i,t} = 0.51 \cdot imp_{i,t} - 0.33 \cdot imp_{i,t} \cdot d_{1997} + 0.89 \cdot M2_{i,t-1} + 0.78 \cdot M2_{i,t-1} \cdot d_{1997} + 0.15 \cdot v(x)_{i,t} \quad (2)$$

We find a positive coefficient on the ratio of imports to GDP and the ratio of broad money to GDP. The volatility of export receipts also exhibits a positive coefficient. The potential for resident-based capital flight from the domestic currency seems to play an increasingly important role in determining reserve holdings in emerging Asia, since the coefficient associated with the ratio of broad money to GDP rises by 0.78 in the post-1997 period. This is consistent with an increasing role for the self-insurance motive against potential internal drain. Current account developments have, however, less of an impact on reserve holdings following the Asian crisis, as the coefficient associated with import propensity declines in the second subperiod. These results are in line with those of Aizenman, Lee, and Rhee (2004), who find evidence that the rapid integration of Korea with the global financial system has increased the weight of financial openness and reduced the weight of trade openness in accounting for the patterns of international reserves.

We use this demand equation to assess the extent to which international reserves diverge from their usual economic determinants. The level of reserves predicted by this set of explanatory variables should not be considered an optimal or desirable level. Rather, it corresponds to the level of reserves consistent with macroeconomic conditions. Appendix C provides graphs of the actual and predicted values for each country.

In China (Figure C1), reserves topped US\$408 billion in 2003. This is surprisingly *lower* than the US\$422 billion predicted by the determinants. Elsewhere in Asia, reserves are slightly above the level suggested by the demand equation: by about US\$32 billion in India, US\$2 billion in Korea, US\$6 billion in Indonesia, US\$11 billion in Malaysia, US\$3 billion in the Philippines, US\$6 billion in Thailand, and US\$6 billion in Singapore. On balance, with a positive gap of US\$52 billion, reserve holdings in emerging Asia as a whole are slightly above the level given by their determinants in 2003.

Table 5 summarizes the actual and predicted figures for 2004.

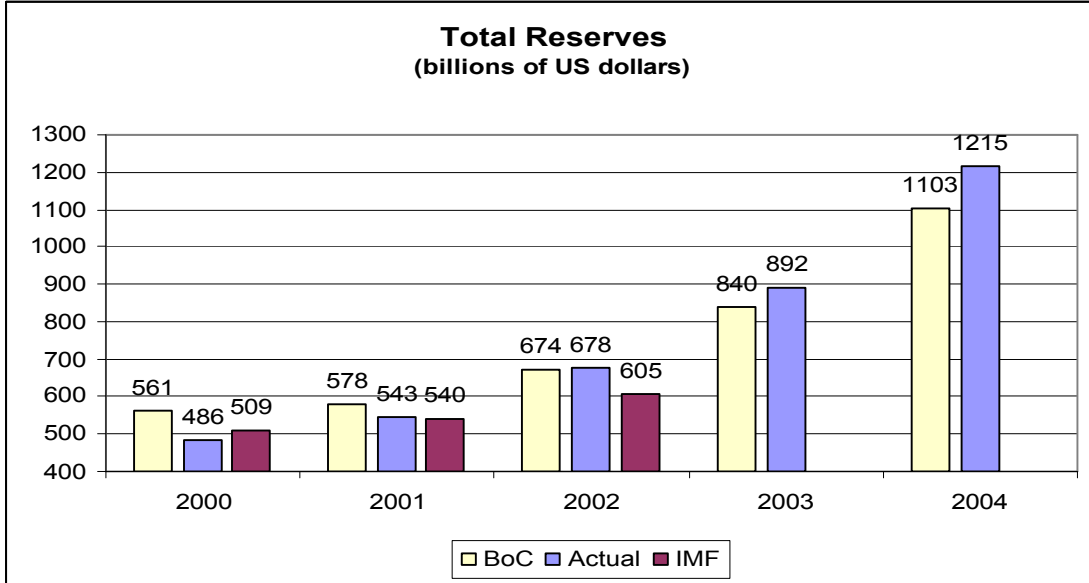
**Table 5: Foreign Exchange Reserves in Emerging Asia: 2004 Forecast (US\$ billions)**

|             | Actual | Predicted | Difference |
|-------------|--------|-----------|------------|
| China       | 614.5  | 562.4     | 52.1       |
| India       | 126.6  | 88.9      | 37.7       |
| Indonesia   | 35.0   | 28.2      | 6.8        |
| Korea       | 199.0  | 218.6     | -19.6      |
| Malaysia    | 66.3   | 40.9      | 25.4       |
| Philippines | 13.1   | 11.4      | 1.7        |
| Singapore   | 112.2  | 111.4     | 0.8        |
| Thailand    | 48.7   | 41.3      | 7.4        |
| Total       | 1215.4 | 1103.2    | 112.2      |

Overall, the positive gap in reserves that occurred in 2003 more than doubled in 2004: emerging Asian reserves exceeded the predicted level by US\$112 billion. As Table 5 shows, the difference comes mainly from China, where the reserve gap swings from -US\$14 billion to +US\$52 billion. With the exception of Korea, actual values surpassed predicted values everywhere in emerging Asia in 2004.

Figure 1 reports the actual and IMF predictions of total emerging Asian central bank reserves along with our results over the period 2000–04. Comparison with the IMF results is possible only up to 2002. Recall that the main differences between our approach and that of the IMF are: a more appropriate sample of countries, more formal time-series econometric handling, and allowance for structural breaks. This yields some key differences in predicted values. In 2000, both models indicate that reserves were below their fundamentals. However, while the IMF finds that reserves came into line with their determinants in 2001, our model suggests that they were still relatively weak compared with their long-run level. Again, there is a significant difference in the 2002 predictions: the IMF concludes that reserves were in excess of their long-run level by US\$73 billion, whereas we find that reserves were essentially in line with their determinants in that year.

**Figure 1: Predicted Reserves in Emerging Asia (2000–04)**



Overall, our model allows for a higher long-run level of reserves in the post-crisis period. Still, the very strong pace of reserve accumulation over the past two years has put the actual level of reserves in excess of US\$52 and US\$112 billion in 2003 and 2004, respectively.

### 4.3 Results: error-correction model

We further examine the cointegrating vector by estimating a fixed-effects panel error-correction model for the per cent change in the ratio of reserves to GDP:

$$\Delta y_{it} = \mathbf{f} + \mathbf{w}_i - V(y_{i,t-1} - \hat{y}_{i,t-1}) + \sum_{k=1}^K \mathbf{y}_k \Delta \mathbf{x}_{k,it} + \mathbf{x}_{it}. \quad (3)$$

This also allows us to examine the explanatory power of stationary variables that were dismissed from the cointegration analysis; i.e., exchange rate volatility and opportunity cost. In addition to these determinants, we consider changes in the variables of the cointegrating space. Results provide additional evidence of cointegration, with a strongly significant error-correction term of -0.56 (Table 6). Four lags of the dependent variable are needed to account for the autoregressive persistence of the series. The residuals are white noise according to the Ljung-Box Q-statistics. At 56 per cent per year, adjustment is fairly rapid. These results are broadly consistent with those of Ford and Huang (1994), who obtain a rate of adjustment of 52 per cent per year for China. Quite surprisingly,

aside from lagged values of the dependent variable, none of the aforementioned variables are statistically significant.

**Table 6: Panel Error-Correction Model**

|  | Coefficient | <i>t</i> -stat |
|--|-------------|----------------|
| <i>ecm</i> <sub><i>t</i>-1</sub>       | -0.56       | -5.58          |
| <i>dlres_gdp</i> <sub><i>t</i>-1</sub> | 0.18        | 1.89           |
| <i>dlres_gdp</i> <sub><i>t</i>-2</sub> | 0.08        | 0.87           |
| <i>dlres_gdp</i> <sub><i>t</i>-3</sub> | 0.15        | 1.95           |
| <i>dlres_gdp</i> <sub><i>t</i>-4</sub> | -0.10       | -1.45          |
| <i>dum</i> <sub><i>crisis</i></sub>    | 0.36        | 3.66           |
| <i>R</i> <sup>2</sup>                  | 40%         |                |
| <i>LB-Q</i> (1)                        | 0.595       |                |
| <i>LB-Q</i> (4)                        | 0.857       |                |

## 5. Discussion

In this section, we put our results into a broader perspective by examining other costs of holding reserves that our model cannot take into account. We then discuss the implications for the U.S. dollar.

Even if reserves had been in line with the identified long-run determinants up to 2004, there would still be a risk that some key elements are missed by our empirical approach. Indeed, overaccumulation of reserves entails domestic costs that our model cannot take into account. These include exchange rate misalignment, loss of monetary control, and sterilization costs. Taking these costs into account would reduce the desired level of international reserves, such that the gap relative to actual reserves would be even larger.

Examining the empirical role of such domestic costs is difficult for various reasons. First, instances of excessive reserves are virtually nonexistent over history. Indeed, until relatively recently, it was common wisdom that a country with conditional access to global capital markets would benefit from holding international reserves that were as large as possible to insure against the risk of capital reversals or sudden stops. Thus, it is very difficult to empirically establish a causal link between excessive reserve accumulation and the probability of domestic financial crises based on an historical data set that reflects an entirely different paradigm. Second, from an econometric perspective,



these costs are endogenous. They are the consequence, not the cause, of excessive reserves. As such, any proxy for these costs cannot be used to explain reserves unless they have some kind of forward-looking or expected component. To address these issues, we would need to perform policy or welfare analyses under a general-equilibrium framework.

### **5.1 Exchange rate misalignment**

Excessive reserves may be harmful when the rapid accumulation is a consequence of adopting and maintaining a fixed exchange rate, as seems to be the case for some Asian countries (Osakwe and Schembri 1998). Fixed exchange rates and higher reserves may indeed increase the vulnerability to a crisis, since economic agents may perceive them as an implicit guarantee and may not take sufficient insurance against exchange rate variability. Moreover, an undervalued exchange rate can have harmful effects on growth and welfare, by reducing consumption, inducing overinvestment in the traded goods sector, lowering domestic investment, and excessively increasing exposure to external shocks.<sup>19</sup>

### **5.2 Loss of monetary control**

The rapid accumulation of reserves can generate inflation. The authorities' ability to sterilize capital inflows is limited. For instance, in China in 2004, only about half of the liquidity arising from the increase in international reserves was sterilized. Rapid credit expansion and higher inflation could also lead to speculative bubbles, which might jeopardize domestic financial stability. The longer the authorities attempt to resist market pressures, the more difficult it becomes for them to retain monetary control. At some point, Asian central banks will have to either abandon their efforts to peg their currencies or lose control of monetary and financial expansion. Prasad, Rumbaugh, and Wang (2005) highlight these different costs and argue that it is typically better to allow the required adjustment to take place through changes in the nominal exchange rate than

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<sup>19</sup> Global savings could be misallocated, with overinvestment in export industries in Asia and underinvestment in the traded goods sector in the United States.

through inflation. Inflationary dynamics can pose serious risks because expectations of rising inflation can become entrenched, particularly in developing economies.

### **5.3 Sterilization costs**

Sterilization is costly, being roughly equal to the interest paid on the country's public debt minus the interest earned on reserves (typically, the interest rate on the U.S. Treasury debt). To the extent that domestic interest rates are above the rate of return on the reserve asset, the holding of reserves entails quasi-fiscal costs. According to the IMF, the cost of sterilizing a reserve accumulation of 10 per cent of GDP can range from zero to 1 per cent of GDP, depending on the interest spread and the expected exchange rate depreciation. In order to continuously sterilize inflows, the monetary authority has to offer ever-increasing interest rates, which would dampen domestic demand. Consequently, economic growth would likely be reduced following an episode of massive and prolonged sterilization.

### **5.4 Implications for the U.S. dollar**

Overall, our empirical results for 2003 and 2004, combined with the aforementioned cost considerations, indicate that, *ceteris paribus*, a slowdown in the speed of accumulation of reserves is likely. This implies negative risks for the U.S. dollar. Although the error-correction model suggests that adjustment can be relatively quick, changes in holding policies might actually be very gradual in the current context. Indeed, the amount of reserve assets held by Asian central banks is so large that any change in holding policies could have a substantial impact on the U.S. dollar and consequently on the balance sheets of Asian central banks. To avoid large capital losses, Asian central banks will be very cautious when slowing the rate of reserve accumulation. The recent decision by the Bank of China to adopt a new basket to which to peg its currency reflects this cautious approach. As a result, the chances of Asian central banks triggering a **rapid** depreciation of the U.S. dollar are not very high.

The currency composition of reserve stocks may pose an additional risk for the U.S. dollar. It is possible that, from the standpoint of international diversification, the

portfolios of Asian central banks would reveal an overweight in dollar assets.<sup>20</sup> *Ceteris paribus*, diversifying away from the dollar would reduce capital losses in the event of a reduction in reserve holdings (autonomous or coming from a currency revaluation). There are signs that central banks are currently losing their appetite for U.S. debt, as they appear to be reducing their dollar holdings in favour of the euro. For instance, a recent survey of 59 central banks by Pringle and Carver (2005) shows that, in 2004, 70 per cent of respondents increased their euro holdings, while 52 per cent reduced their dollar holdings. This is, however, only anecdotal evidence. Data on the currency breakdown of central banks' balance sheets is confidential, precluding any formal analysis of the currency composition of reserves.

Using unpublished IMF-World Bank data, a study by Dooley, Lizondo, and Mathieson (1989), updated by Eichengreen and Mathieson (2000), examines the determinants of the currency composition of reserves for developing countries. They find that the composition of reserves is responsive to the choice of currency peg, the identity of the dominant trading partner, and the composition of foreign debt. More importantly, they find that the currency composition of reserves is remarkably stable over time. This is not surprising, given that variables such as trading relationships and the composition of foreign debt give evidence of substantial inertia. Thus, the currency composition of reserves appears to evolve only gradually, so that a **radical** currency reallocation of reserves is not very likely to happen within a short period of time. Therefore, although the outlook for the U.S. dollar may not be favourable from the standpoint of the currency composition of reserves, risks of an **abrupt** dollar depreciation *coming from this factor* remain limited.

## 6. Conclusion

We have examined the issue of reserve accumulation by central banks in emerging Asia by estimating a reserve-demand function in a panel of eight Asian emerging-market economies. Although we accounted for a structural break in the demand for reserves in

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<sup>20</sup> BIS data reveal that dollar-denominated securities accounted for roughly 70 per cent of total reserves at the end of 2003.

the aftermath of the Asian financial crisis, our model could not explain the very strong pace of reserve accumulation in the past two years. This suggests that, as long as historical relationships continue to hold, a slowdown in the pace of reserve accumulation is likely. This finding implies negative risks for the U.S. dollar. However, the substantial capital losses that Asian central banks would incur if they were to drastically change their holding policy mitigate the risks of a **rapid** depreciation of the U.S. dollar triggered by Asian central banks.

As a future step to this research, it would be useful to have a more rigorous framework in which to characterize the potential dilemma facing monetary authorities when accumulating large stockpiles of reserves; i.e., abandon the currency peg or lose control of monetary and financial expansion. To develop this type of framework, we would need to construct a theoretical model in which central banks seek to simultaneously minimize the risks of external as well as internal crises. The empirical counterpart of such a model could be based on finding an interval for the level of reserve holdings such that the probability of external and internal crises is simultaneously minimized.

Another area of research could be to examine the interactions between the financial system and the process of foreign exchange reserve accumulation. Aizenman and Lee (2005) investigate the micro foundations of the precautionary demand for reserves along the lines of the Diamond and Dybvig (1983) bank-run model. Also, Velasco (1987) argues that foreign exchange reserves can be used to support failing banks and other financial institutions. This suggests that some financial variables, like average non-performing loans, or the number of bank closures/mergers, could have predictive power for reserves.

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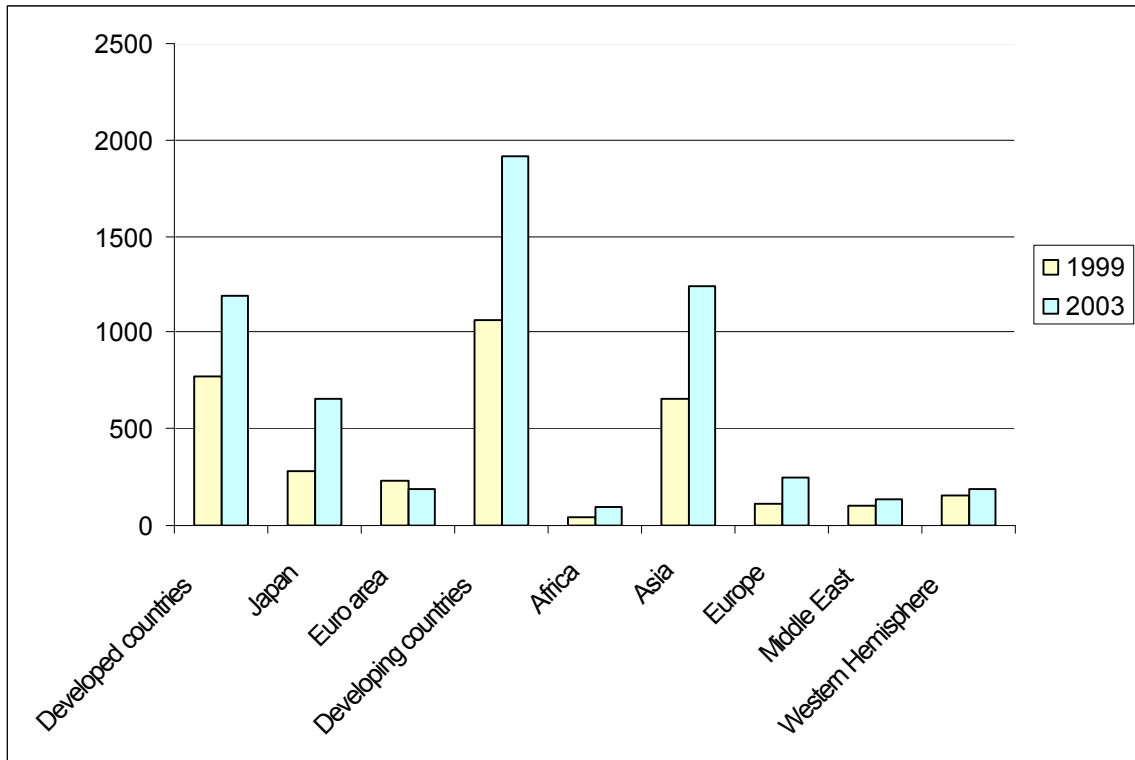
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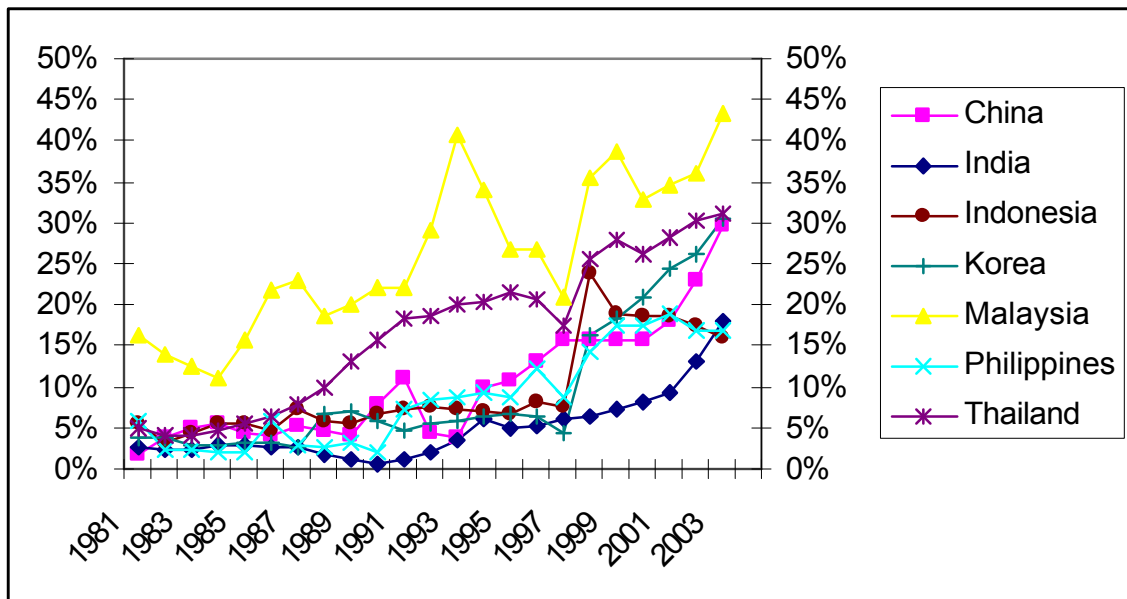
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**Appendix A: Stylized Facts**

**Figure A1: Global Reserve Stocks (billions of U.S. dollars)**

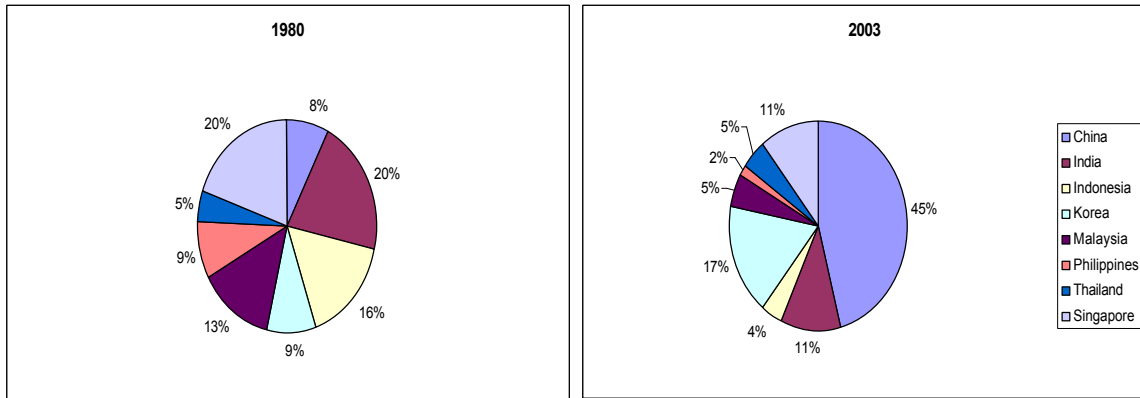


**Figure A2: International Reserves as a Percentage of GDP**

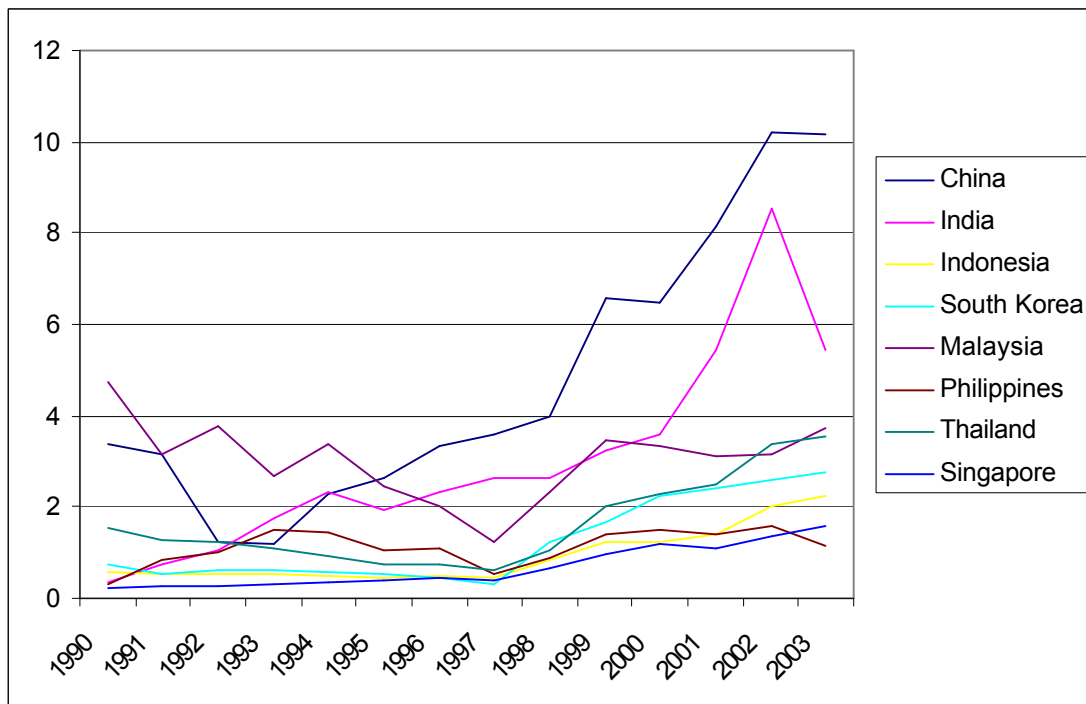




**Figure A3: Regional Distribution of Reserves**



**Figure A4: Ratio of Reserves to Short-Term External Debt**



**Table A1: Exchange Rate Regime Classifications**

| ER regime classification | IMF classification   |
|--------------------------|--|
| Fix                      | Currency boards<br>Single-currency pegs<br>Basket pegs                                   |
| Intermediate (Int)       | Crawling pegs<br>Target zones  |
| Flexible (flex)          | Float with some intervention<br>(no predetermined range for intervention)<br>Pure floats |

|  | 1980 | 1981 | 1982 | 1983 | 1984 | 1985 | 1986 | 1987 | 1988 | 1989 | 1990 |
|--|------|------|------|------|------|------|------|------|------|------|------|
|--|------|------|------|------|------|------|------|------|------|------|------|

|             |      |      |      |      |      |      |      |      |      |      |      |
|-------------|------|------|------|------|------|------|------|------|------|------|------|
| China       | NA   | NA   | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  |
| India       | Int  | Int  | Int  | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Indonesia   | Int  | Int  | Int  | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Korea       | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Malaysia    | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  |
| Philippines | Flex | Flex | Flex | Int  | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Singapore   | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Int  | Int  | Int  | Int  |
| Thailand    | Fix  | Fix  | Int  | Int  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  |

|  | 1991 | 1992 | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 | 2001 |
|--|------|------|------|------|------|------|------|------|------|------|------|
|--|------|------|------|------|------|------|------|------|------|------|------|

|             |      |      |      |      |      |      |      |      |      |      |      |
|-------------|------|------|------|------|------|------|------|------|------|------|------|
| China       | Fix  | Fix  | Fix  | Int  | Int  | Int  | Int  | Int  | Fix  | Fix  | Fix  |
| India       | Fix  | Fix  | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Indonesia   | Flex | Flex | Flex | Flex | Flex | Int  | Int  | Flex | Flex | Flex | Flex |
| Korea       | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Malaysia    | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  |
| Philippines | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex | Flex |
| Singapore   | Int  | Int  | Int  | Int  | Int  | Int  | Int  | Int  | Int  | Int  | Int  |
| Thailand    | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  | Fix  |

## Appendix B: Panel Cointegration Statistics (Pedroni 1999)

1. Panel v-stat: 
$$T^2 N^{3/2} \left( \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1}$$
2. Panel r-stat: 
$$T \sqrt{N} \left( \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{I}_i)$$
3. Panel pp-stat: 
$$\left( \tilde{\mathbf{S}}_{N,T}^2 \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{I}_i)$$
4. Panel adf-stat: 
$$\left( \tilde{\mathbf{S}}_{N,T}^{*2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^*$$
5. Group r-stat: 
$$TN^{-1/2} \sum_{i=1}^N \left( \sum_{t=1}^T \hat{e}_{i,t-1}^2 \right)^{-1} \sum_{t=1}^T (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{I}_i)$$
6. Group pp-stat: 
$$N^{-1/2} \sum_{i=1}^N \left( \hat{\mathbf{s}}_i^2 \sum_{t=1}^T \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{t=1}^T (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{I}_i)$$
7. Group adf-stat: 
$$N^{-1/2} \sum_{i=1}^N \left( \sum_{t=1}^T \hat{s}_i^{*2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{t=1}^T \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^*$$

where,

$$\hat{\mathbf{l}}_i = \frac{1}{T} \sum_{s=1}^{k_i} \left( 1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \hat{\mathbf{m}}_{i,t} \hat{\mathbf{m}}_{i,t-s} ,$$

$$\hat{s}_i^2 \equiv \frac{1}{T} \sum_{t=1}^T \hat{\mathbf{m}}_{i,t}^2 ,$$

$$\hat{\mathbf{s}}_i^2 = \hat{s}_i^2 + 2\hat{\mathbf{l}}_i ,$$

$$\tilde{\mathbf{s}}_{N,T}^2 \equiv \frac{1}{N} \sum_{i=1}^N \hat{L}_{11i}^{-2} \hat{\mathbf{s}}_i^2 ,$$

$$\hat{s}_i^{*2} \equiv \frac{1}{T} \sum_{t=1}^T \mathbf{m}_{i,t}^{*2} ,$$

$$\tilde{\mathbf{s}}_{N,T}^{*2} \equiv \frac{1}{N} \sum_{i=1}^N \hat{s}_i^{*2} ,$$

$$\hat{L}_{11i}^2 = \frac{1}{T} \sum_{t=1}^T \mathbf{h}_{i,t}^2 + \frac{2}{T} \sum_{s=1}^{k_i} \left( 1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \mathbf{h}_{i,t} \mathbf{h}_{i,t-s} ,$$

and where the residuals  $\mu$  and  $\eta$  are obtained from the following regressions:

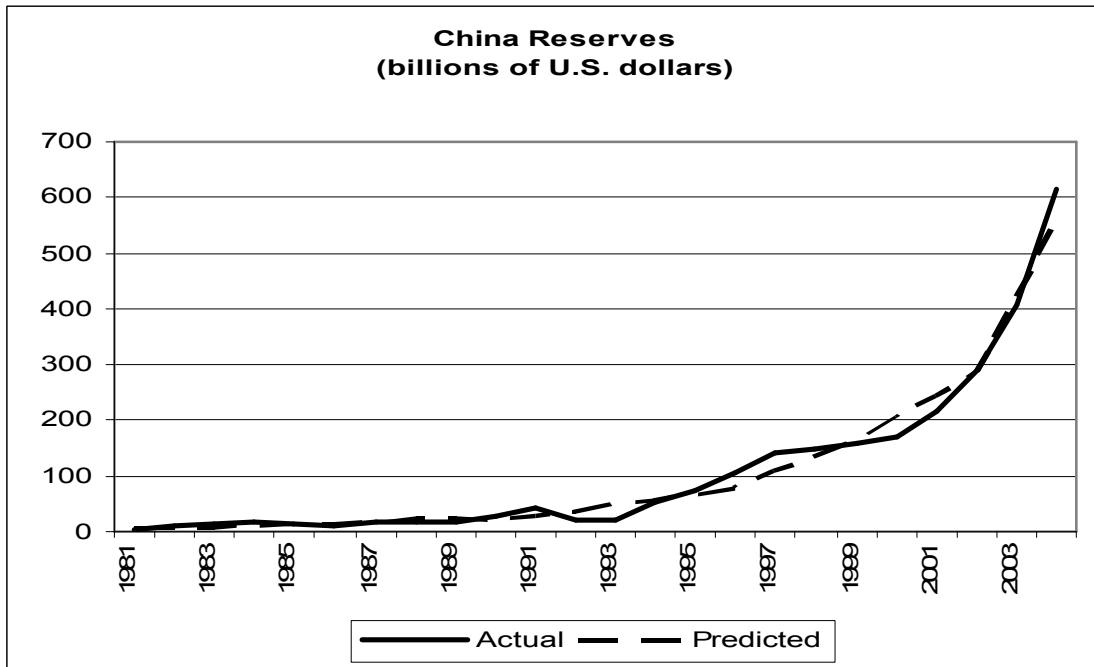
$$\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \hat{\mathbf{m}}_{i,t} ,$$

$$\hat{e}_{i,t} = \hat{\mathbf{g}}_i \hat{e}_{i,t-1} + \sum_{k=1}^{K_i} \hat{\mathbf{g}}_{i,k} \Delta \hat{e}_{i,t-k} + \hat{\mathbf{m}}_{i,t}^* ,$$

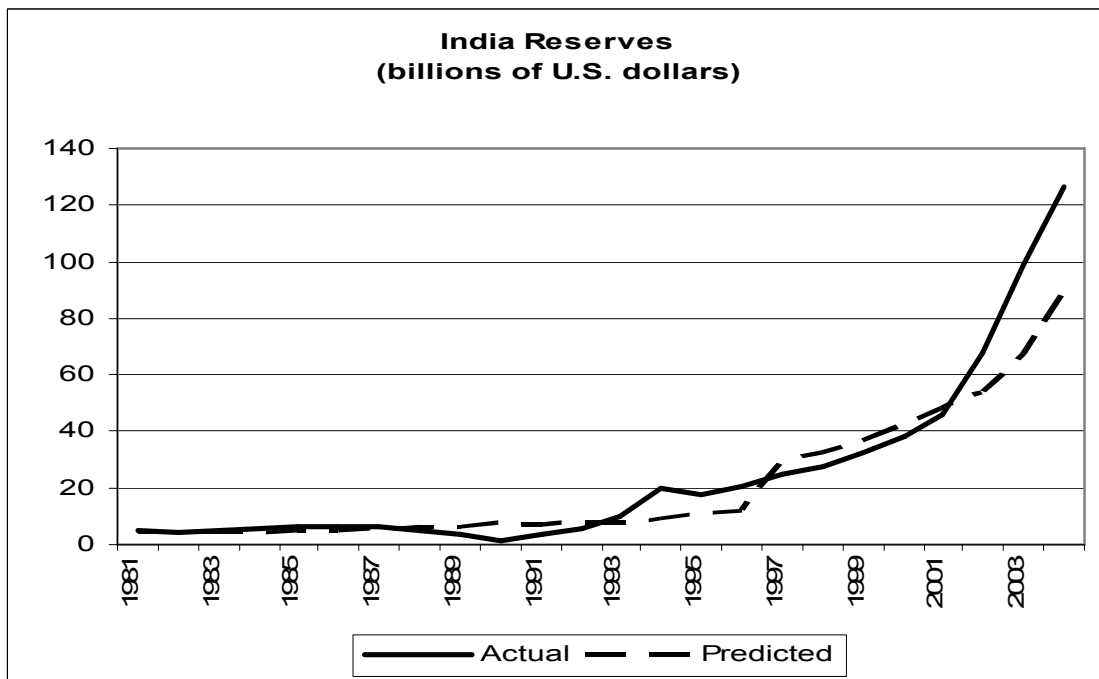
$$\Delta y_{i,t} = \sum_{m=1}^M \hat{b}_m \Delta x_{mi,t} + \mathbf{h}_{i,t} .$$

## Appendix C: Predicted Reserves

### Figure C1: China



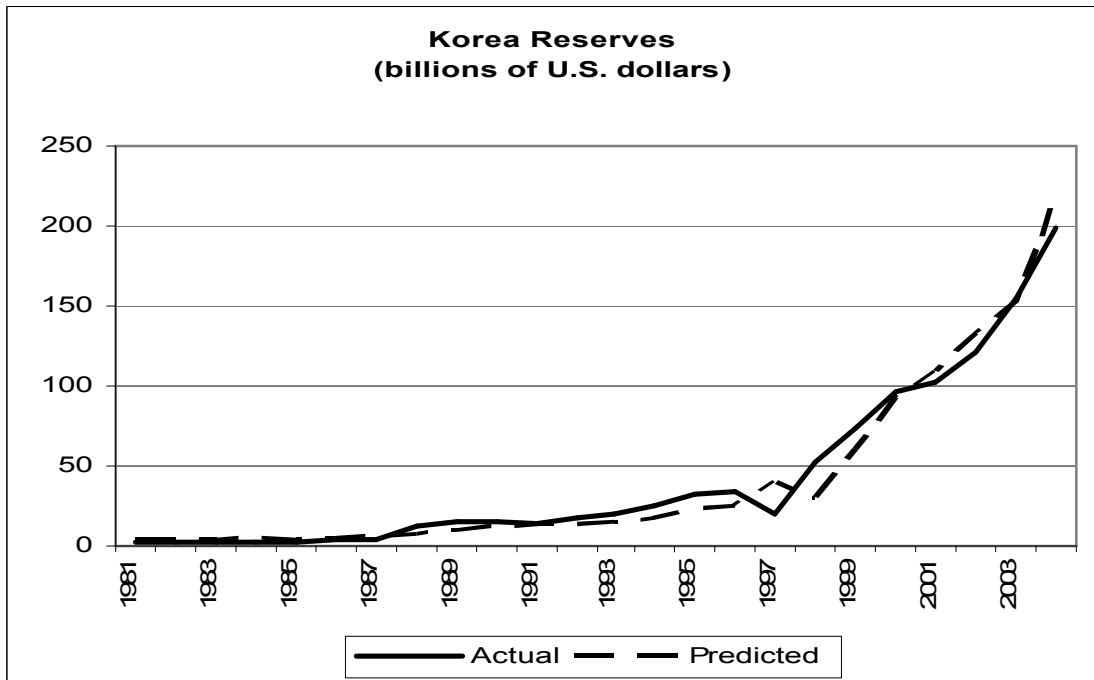
### Figure C2: India



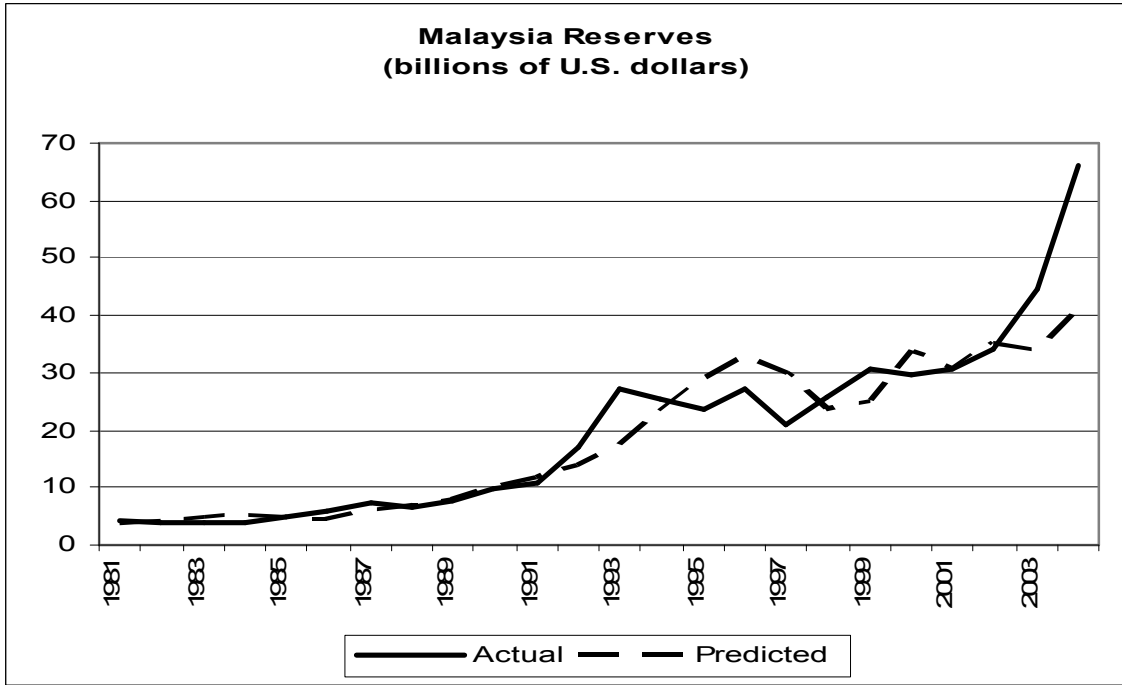
**Figure C3: Indonesia**



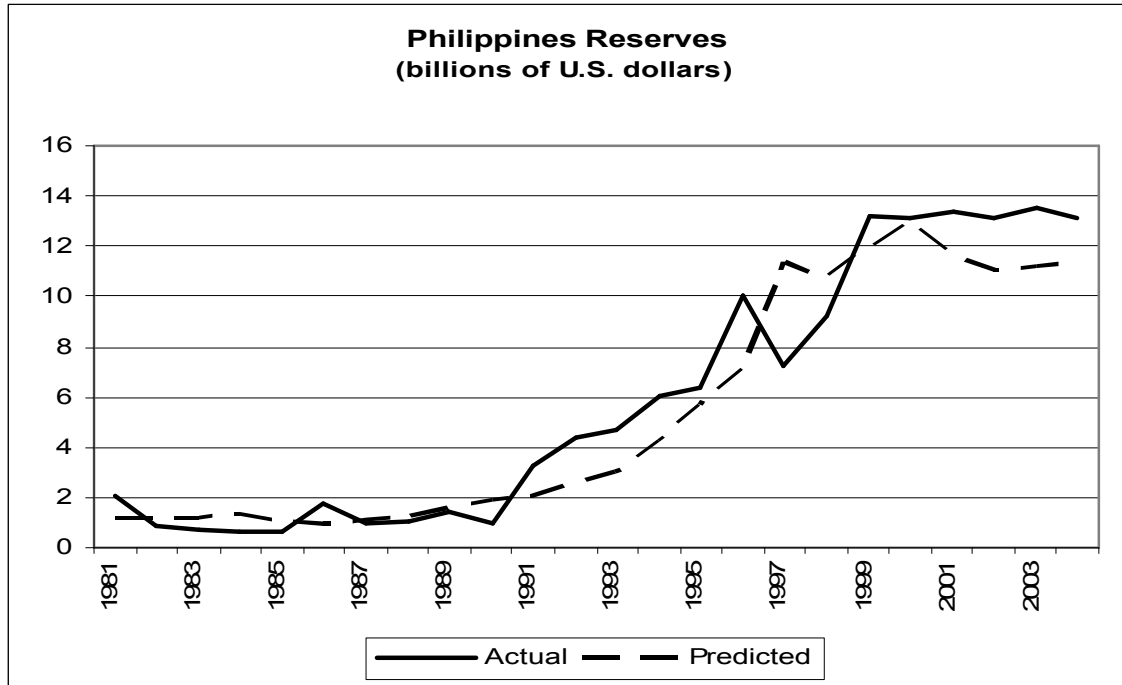
**Figure C4: Korea**



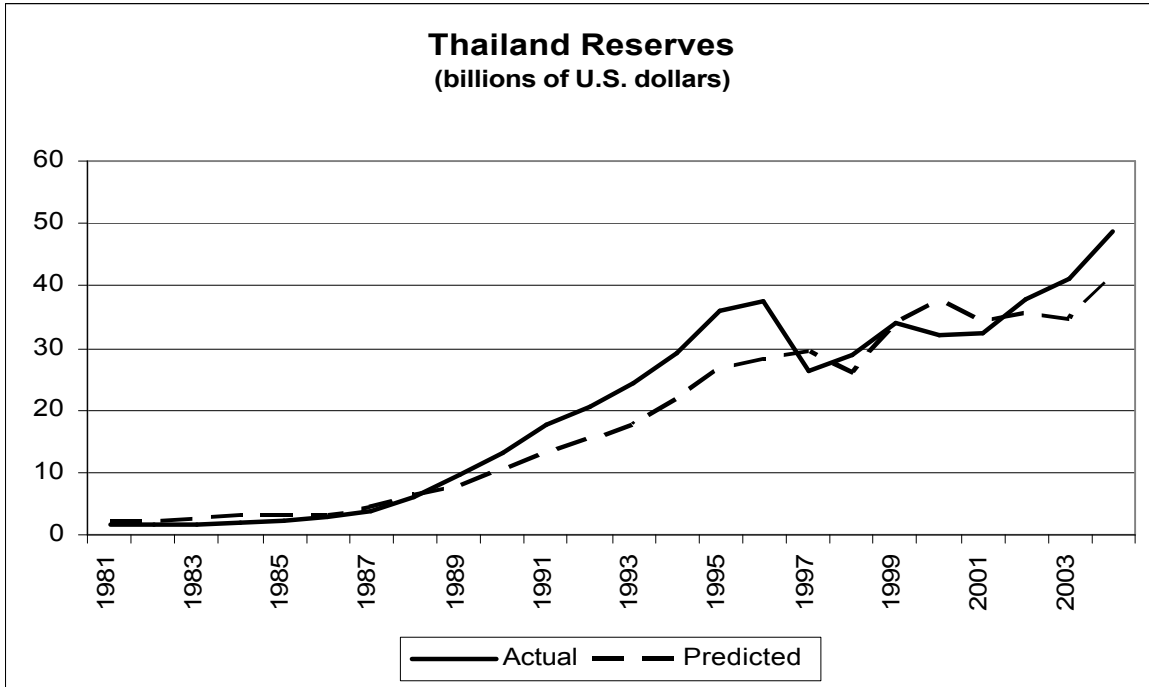
**Figure C5: Malaysia**



**Figure C6: Philippines**



**Figure C7: Thailand**



**Figure C8: Singapore**

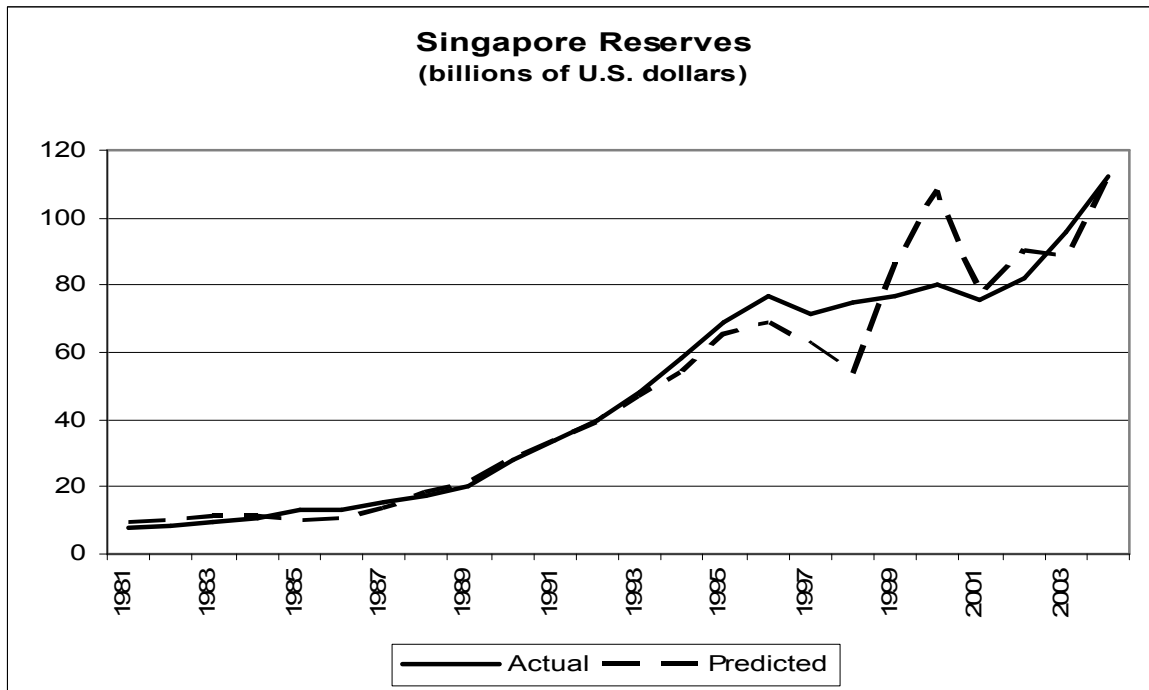
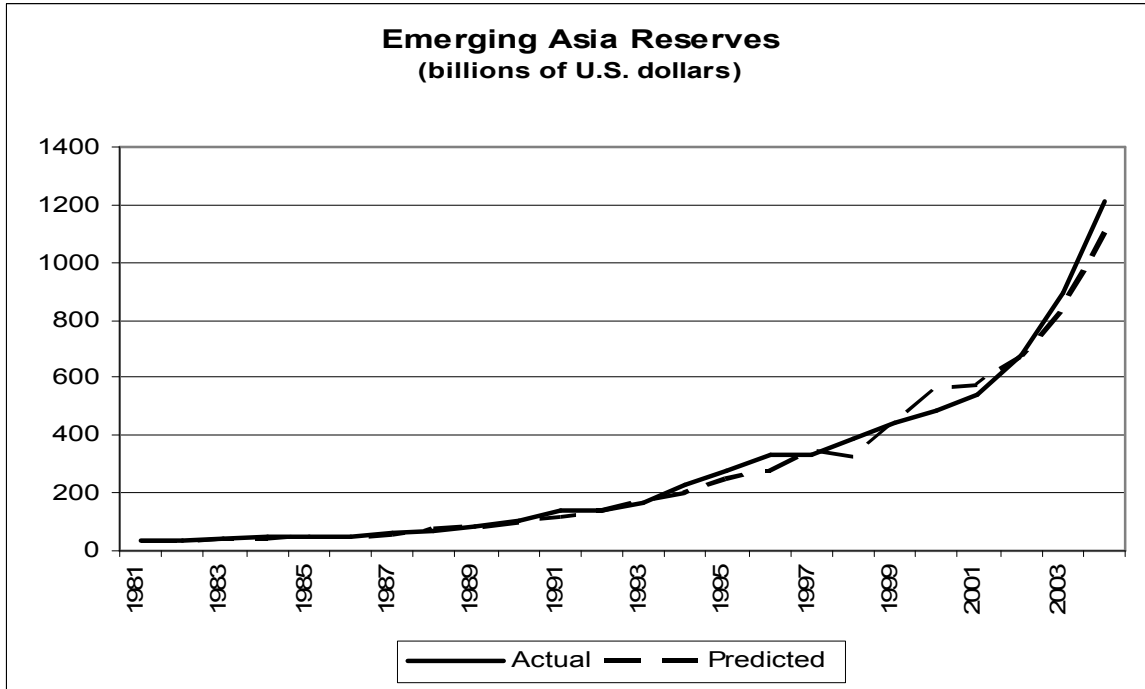




Figure C9: Total



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