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The buffer stock model redux? An analysis of the dynamics of foreign reserve accumulation

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Abstract

Emerging market economies have recently accumulated large stocks of foreign reserves. In this paper we address the question of what are the main factors accounting for reserve holdings in nine developing countries located in Asia and Latin America. Monthly data from January 1985 to May 2006 are used to estimate for each country the long run equilibrium reserve demand, based on the buffer stock model, the short run dynamics governing the process of reserve accumulation (decumulation) and the factors which may influence the speed of adjustment of actual to desired reserves. Cointegration analysis suggests that the buffer stock precautionary model accounts for the optimal reserve demand. The corresponding VECMs are further interpolated, using the permanent and transitory innovations decomposition procedure of Gonzalo and Ng (2001), in order to assess the relative impact of the time series on the convergence to equilibrium after a shock. Finally the (asymmetric) effect on the speed of convergence of positive/negative changes in signal variables - such as the excess reserves of the previous period, relative competitiveness and US monetary stance - is found to be significant, in line with mercantilistic and fear of floating motives for hoarding international reserves.

Keywords: Emerging markets reserves, cointegration, P-T components decomposition, asymmetric adjustment
JEL classification: F32, F34, F36

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The rapid process of reserve accumulation in emerging market economies has become one of the most controversial topics in international economics, coming under investigation in a burgeoning literature. By the end of December 2006 the foreign reserve assets of the emerging countries reporting to the IMF were worth over $1400bn, 1.4 times their value at the end of 2004. Since reserve accumulation is costly it has to be justified by economic benefits. Mercantilist considerations are often mentioned, especially in Asia. Most explanations are, however, still based on a self-insurance rationale; reserves are accumulated in order to avoid the disruptive effects of a rapid capital (short-term foreign investment) outflow in periods of stress. This interpretation is reminiscent of the models discussed in the 1960s and 1970s, when reserves were viewed as a buffer accumulated by the monetary authorities in order to avoid exchange rate crises brought about by a drought of international means of payments due to current account deficits.

This cost benefit interpretation has, in spite of the dramatic changes in both the volume and the nature of international transactions, and despite the exponential growth of international financial markets, maintained its basic validity. The monetary authorities of the emerging market economies accumulate reserves because, in periods of stress, they may be unable to borrow on international markets the foreign funds needed to offset sudden capital outflows. They find themselves in the same predicament as the authorities of the industrialized countries of the previous decades, when the possibility of financing a deficit on international markets was – for institutional reasons – very limited. We therefore maintain that the traditional buffer stock precautionary model of 1970-1980 vintage continues to offer a valid explanation of the optimal reserve demand on the part of the emerging market economies. This does not mean, however, that issues currently discussed in the literature, such as mercantilistic real exchange rate manipulation or fear of floating, are irrelevant to reserves management policies.[1] Indeed, as we shall see, they do play a relevant role in the short run and affect the speed of adjustment of current reserves to their optimal long run level.

In this paper monthly data from January 1985 to May 2006 are used to investigate the national idiosyncratic and international determinants of reserve changes in nine emerging market economies located in Asia and Latin America. Given the broad data span it is possible to estimate, for each country, both the long run equilibrium reserve demand, the short run dynamics governing the process of reserve accumulation (decumulation) and the factors which may influence the speed of adjustment of actual to desired reserves. The non-stationary nature of the data dictates the choice of estimation procedure. The long run demand for reserves is quantified by cointegration relationships and the short run dynamics are modelled with Vector Error Correction parameterizations.

Cointegration reserve demand equations are first independently estimated for each country. The Gonzalo and Ng (2001) procedure is then used to assess the relative impact of the differing time series on the convergence to equilibrium after a shock. Finally, investigation turns to the (asymmetric) impact of positive/negative changes in signal variables, such as the excess reserves of the previous period, and of relative competitiveness and US monetary stance indicators on the speed of convergence. From an economic point of view the cointegration residual time series are assumed to quantify the fraction of the long run reserve holdings that cannot be accounted for by the buffer stock model explanatory variables, viz. domestic factors such as balance of payments variability, and a national short term rate which measures the opportunity cost of holding reserves. International explanatory factors, reflecting the pivotal role

[1] Recent contributions on this topic by Aizenman et al. (2004) and Aizenman and Lee (2005), among others, simply add control variables in static reserve demand relationship estimates and disregard the information provided by the dynamics of reserve accumulation.

played by the US monetary authorities in emerging markets finance and mercantilistic motives, affect the short run disequilibrium dynamics.

The paper improves upon previous research in the following respects:

- the analysis of the adjustment process is consistent with the stringent restrictions of cointegration analysis dynamics and a distinction is drawn between the factors which enter the long run equilibrium demand for reserves and those which appear only in the short run;
- the variable which, in each country, brings about a readjustment after a shock to the equilibrium demand for reserves is carefully identified; in all but two cases the reserves do play the prominent role;
- the impact of exogenous international factors on the speed of adjustment of reserves is then investigated in contexts of both positive and negative overstocking; it sheds light on the role of reserve positive (negative) shifts on the mercantilist policies, and on the reaction of the countries in the sample to US interest rates.

The paper is structured as follows. Section 1 discusses the tenets of the buffer stock model of reserve demand and the recent issues set out in the literature; section 2 estimates reserve demand in the emerging markets of the sample using time series and cointegration analysis procedures, and investigates the dynamics of the adjustment process in each country VECM; section 3 analyses the reaction of excess reserve accumulation to the innovations in past reserve changes and in international competitiveness and financial liquidity indicators. Section 4 concludes the paper.

1 The buffer stock model and the demand for reserves

The buffer stock model posits that the authorities select the stock of reserves which balances the potential macroeconomic costs due to the lack of reserves with the opportunity cost of reserve accumulation. In this section the basic tenets of the model are set forth, from the original version by Heller (1966) to the more sophisticated stochastic one by Frenkel and Jovanovic (1981), along with some recent extensions. The factors that influence the speed of adjustment of reserves to their optimal desired level are then investigated.

1.1 The basic model

An adequate stock of reserves smoothes out payment disequilibria and prevents costly financial readjustments. The Asian crisis has shown that countries holding large reserves were able to weather it better than the others. Rodrik and Velasco (2000) estimated that a country would reduce by ten percent the probability of experiencing a sharp capital outflow if it were to abide by the Guidotti rule and hold reserves equal to its short term debt.

Reserves thus provide a self-insurance service that has to be paid for with the opportunity cost of investing financial resources in potentially sub-optimal foreign currency assets. The first formal discussion of precautionary optimal reserve management, based on the minimization of the total cost of financing and adjusting to external shocks, can be found in Heller (1966); it provides the basic framework for most subsequent research in the field. Frenkel and Jovanovic (1981) extend a previous contribution by Hamada and Ueda (1977) developing a rigorous – possibly definitive - formal setting of the hypotheses set out in Heller.

Reserves follow in their model a Wiener process and, immediately after a restocking, the authorities select the initial level that minimizes total expected costs. The latter have two interrelated stochastic components: (i) the opportunity cost of reserve holdings and (ii) the adjustment cost of reserve restocking whenever the latter have reached a lower limit, set here to zero. The adjustment cost stems from the output reduction brought about by the need to generate the balance of payments surplus which will in turn generate the reserve build up. A higher variability implies that reserves are likely to reach their lower bound more often and require costly restockings. The authorities are faced with a standard cost-benefit choice: the larger the stock of reserves, the lower is the expected cost of adjustment and the higher the expected value of the opportunity cost (and vice versa).

After some algebraic manipulation, Frenkel and Jovanovic obtain the following approximation of optimal initial reserve holdings (R_0)

$$\log R_0 = \log C + 0.5 \log \sigma - 0.25 \log r \tag{1}$$

where *σ* is the balance of payments variability, *r* the opportunity cost and C the fixed cost of accumulating reserves. The additional hypothesis is then made that observable reserves R_t are proportional to optimal (initial) reserves up to an error term that is uncorrelated with the above mentioned right hand variables, i.e., in logarithmic terms, that

$$\log R_t = A + \log R_0 + u_t \tag{1'}$$

Adding as additional scale variable the level of imports M_t, the following testable relationship is then derived

$$\log R_t = b_0 + b_1 \log \sigma_t + b_2 \log M_t + b_3 \log r_t + u_t \tag{2}$$

where it is assumed a priori that $b_1 > 0$, $b_2 > 0$ and $b_3 < 0$.[2]

The explanatory power of the buffer stock model, both in the industrialized and in the developing countries, has been investigated in a large empirical literature spanning more than twenty years, and summarized in Bahmani-Oskooee and Brown (2002).

1.2 Recent issues on optimal reserve accumulation

Two major strands can be identified in the recent literature on the buffer stock model. The first attempts to adapt it to the institutional and financial characteristics of emerging market economies. Aizenman and Marion (2002, 2004) augment a relationship analogous to equation (2) with political uncertainty and corruption proxies and show that they affect reserve holdings in developing countries. They view reserves as a form of precautionary saving for countries with difficult access to global capital markets and insufficient domestic tax collection. Reserves are also seen as output stabilizers. Aizenman et al. (2004) point out that reserve holdings mitigate the probability of a banking crisis and reduce the expected cost of a sudden freeze of international capital inflows. As demonstrated in Aizenman and Lee (2005), a macro-liquidity shock to an emerging market cannot be diversified away and may force liquidation of a first period investment if it exceeds the level of reserves outstanding, reducing second period output.

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[2] Reserve holdings are reduced if the opportunity cost rises and increased whenever the volatility index rises. The coefficient of the value of imports is positive. It reflects the requirements of international trade on the banks' transactions demand for reserves. We are not imposing here the additional restrictions that $b_1 = 0.5$ and $b_3 = -0.25$ even if, as we shall see below, the estimated cointegration coefficients take values that are surprisingly close to the theoretical ones.

The second strand focuses on quantitative analysis of the cost-benefit structure of the model. García and Soto (2004) and Rodrik (2006), among others, adopt an approach originally set out by Ben-Bassat and Gottlieb (1992) to evaluate the impact of reserve accumulation on the probability of default and compare the opportunity cost of excess reserve holdings with the corresponding estimated benefits. The latter are given by the output loss that is avoided thanks to the reduction in the probability of a financial crisis brought about by the accumulation of reserves. It should be noticed that the cost of reserve depletion is now explicitly associated with the output cost of a default.

1.3 Modelling the reserve adjustment process

Actual reserve holdings do not always coincide with optimal reserve demand. Indeed, a reserve accumulation process is onerous and the adjustment may be protracted over time. Kenen and Yudin (1965), Iyoha (1976), Bilson and Frenkel (1979) and Ben-Bassat and Gottlieb (1992) among many others have investigated the dynamics of the reserve accumulation.

They set out a partial adjustment relationship of the form

$$\Delta(\log R_t) = \mu + \gamma(\log R_t^* - \log R_{t-1}) \qquad (3)$$

where R_t^* is the desired (optimal) stock of reserves and γ measures the speed of adjustment. It is usually assumed that $0 < \gamma \leq 1$.

The dynamics of our paper are driven by the error correction parameterization of cointegration analysis. We have therefore used two complementary procedures that are compatible with the Granger Representation Theorem. The first, derived from the Gonzalo and Ng (2001) interpolation of the Wold representation of the model, allows us to assess, in each country, the relative impact of the differing time series on the convergence to equilibrium after a shock.

The second focuses on the VECM parameterization itself, and makes use of a Heaviside function in order to isolate the effects of innovations in a set of variables on the error correction process. The latter can thus be modelled as follows

$$\Delta \log R_t = \varphi + \beta H_t(Z_t) EC_{t-1} + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + e_{1t} \qquad (4)$$

where EC_t is the distance between the actual level of reserves and its equilibrium value provided by equation (2). X_t is the (column) vector of I(1) variables of the cointegration relationship, Γ_j^1 is a (row) vector of distributed lag coefficients and H_t is a Heaviside function of specific exogenous factors (Z_t). It models non linearities in the adjustment process, on which sound evidence is provided in the literature (see Escribano and Granger, 1998, and Escribano and Pfann, 1998, among many others).

2 Assessment of optimal reserve demand

This section describes the econometric strategy used to determine the optimal long run demand for reserves. Our measure of reserve adequacy evolves over time and provides a dynamic benchmark that can be used to assess the magnitude of overstocking. The optimal buffer stock model demand for reserves, discussed in section 1.1, is estimated in a two-step multivariate cointegration approach. The data set spans the January 1985 – May 2006 time period and encompasses some important episodes of distress in both Asia and Latin America.

2.1 Stationarity and volatility analysis

Recent econometric findings summarized in Vogelsang (1999) have shown that additive outliers introduce into the residuals of standard unit root test estimates a moving average component with a negative coefficient which, in turn, inflates the size of the test and causes over-rejection of the null hypothesis. The Latin American and Asiatic crises brought about long lasting changes in Central Bank behavior, and the corresponding outliers in the time series may well be considered additive. We implemented the test procedure of Perron and Rodríguez (2003) and identified several additive outliers.[3] The unit root tests of table A.I, appendix II, are thus performed using the statistic by Ng and Perron (2001), which is robust to size distortions due to negative serial correlation of the residuals. They fail systematically to reject the null of non stationarity.

Additive outliers may also distort inference on cointegration rank in finite samples (Franses and Haldrup, 1994). Following the interpolation strategy suggested by Nielsen (2004), the outlying observations are eliminated and replaced with an average of the respective adjoining data. The smoothed time series will then be used in the cointegration analysis below.

The balance of payments variability index plays a significant role in models of optimal demand for foreign reserves and is to be carefully estimated. Most of the previous empirical studies estimate a multiperiod rolling standard deviation of (detrended) reserve changes. Our sample period includes spells of turbulence, and the reserve changes display volatility clustering i.e. autoregressive conditional heteroskedasticity.[4]

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[3] The value of test statistics significant at the 5 percent level and the corresponding dates are available from the authors upon request.

[4] The presence of ARCH effects is corroborated by the serial correlation of the squared reserve increments, assessed with the help of Ljung Box Q-statistics.

Unbiased volatility estimates are thus obtained using conditional measures, computed as the square root of the (T)GARCH(1,1) variance of monthly reserve changes (for more details see Cifarelli and Paladino, 2005).

2.2 Multivariate cointegration estimates of the demand for reserves

A two-step procedure is used to estimate, for each country, the following VECM

$$\Delta X_t = \Phi + \mathrm{B}\alpha' X_{t-1} + \sum_{j=1}^{k} \Gamma_j \Delta X_{t-j} + e_t \qquad (5)$$

where Φ is a n x 1 vector of constant terms and Γ_j is a matrix of distributed lag coefficients. Β and α are n x m matrices of, respectively, adjustment coefficients and cointegration equation coefficients, n being the number of I(1) time series in X_t and m the number of cointegrating relationships. X_t = $(\log R_t, \log \sigma_t, \log M_t, \log r_t)'$, where σ_t is the fitted value of a preliminary conditional volatility estimate of the reserve change, M_t is the volume of imports and r_t is a domestic short term interest rate.[5] m = 1 in the empirical analyses below since the trace test statistics of Johansen (1991) set out in table A.II of appendix III identify, in each country system, a single cointegration relationship.

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[5] r_t is the opportunity cost of holding reserves, which, in the case of emerging markets, are mostly invested in US assets. It should thus be measured as the spread between a domestic interest rate and a US Treasury bond interest rate. Since the latter is small and tends to vary but little with respect to the domestic rate we quantify the opportunity cost as the emerging market rate. Only short rates are available over the 1985-2006 time span.

Table I presents the cointegration equation estimates. The long run reserve demand is formulated as

$$\log R_t - \rho_0 - \rho_1 t - b_1 \log \sigma_t - b_2 \log M_t - b_3 \log r_t = \alpha' X_t = \varepsilon_t \qquad (6)$$

The estimates are obtained using the asymptotically efficient DOLS approach of Stock and Watson (1993) and – as pointed out by Maddala and Kim (1999) - are less sensitive to the lag specification of the VECM than those computed with the standard Johansen maximum likelihood estimation procedure. We choose the lag/lead order at which the quality of fit becomes stable, even if the latter reacts only marginally to the order of the DOLS specification. Standard errors are computed using the Newey-West heteroskedasticity and correlation consistent procedure.

Table I DOLS COINTEGRATION EQUATION ESTIMATES

The estimates are rather good and, with a few exceptions, in line with the specification of the model. A rise in interest rates is associated with an increase in reserve holdings in Brazil and in Malaysia, possibly reflecting - as suggested by Aportela et al. (2005) – the effect of foreign capital inflow sterilization policies by local monetary authorities. Reserve volatility fits well with the model highlighting the relevance of restocking costs for economies that were most exposed to contagious crises and the corresponding coefficients are always significant and of the correct sign.[6] The same holds true for the coefficient of the volume of imports, the only exceptions being Mexico and Malaysia. The size of these coefficients usually supports the hypothesis of economies of scale in the use of reserves.

[6] It should be noticed that in the case of Chile and Korea the variability coefficient estimate is also close to the theoretical absolute value posited by the Frenkel and Jovanovic model. The same result is obtained for the interest rate coefficient in Indonesia and Venezuela.

With the second step the VECM equation (5) is estimated for each country with standard VAR parameterization, inserting the lagged estimates of $\alpha' X_t$, obtained in the first step of the procedure, as a predetermined variable. The corresponding error correction coefficients, along with standard tests on the system residuals, are set forth in table II.[7]

Table II ERROR CORRECTION COEFFICIENTS

The VECM diagnostics are in all cases satisfactory. In most countries the coefficients of the lagged cointegration residuals that correspond to reserves are significantly different from zero and suggest that the latter play a relevant role in the error correction process. In the case of Singapore and Chile this role seems to be played by the rate of interest, either in isolation or along with the remaining variables of the system. Analysis of the dynamic effects of permanent and transitory shocks is implemented in order to assess these hypotheses.

2.3 Permanent and transitory decomposition of the VECM innovations

Having identified a single cointegration relationship among the four I(1) time series in X_t, we are left with three permanent shocks or common trends and one transitory shock. The approach of Gonzalo and Ng (2001) is applied to decompose X_t into transformed innovations characterized by differing degrees of persistence.

Starting from the following Wold representation

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[7] The VAR lag order is carefully ascertained and is similar to the order used in the trace cointegration tests set out in the appendix. We follow Urbain (1995) and base the choice of the VAR order, in each country, on two criteria: (i) the absence of serial correlation of the residuals (using a multivariate LM test for residual correlation originally set forth by Johansen, 1995,

$$\Delta X_t = \Phi + C(L)e_t \qquad (7)$$

where Φ is a vector of constant terms, e_t is an n x 1 vector of residuals and C(L) is a matrix of distributed lag coefficients, an n x 1 vector of transformed shocks ν_t is computed where $\nu_t = (\nu_{1t}, \nu_{2t}, \nu_{3t}, \nu_{4t})'$. It is assumed that the first three shocks are permanent and the fourth is transitory in the sense of Gonzalo and Granger (1995). They are obtained using the estimated coefficients of the VECM equation (5) above. Assuming that $G = (\mathrm{B}_{\perp}^{'}, \alpha')'$, with $\mathrm{B}_{\perp}^{'}\mathrm{B} = 0$, the transformed residuals become $\nu_t = Ge_t = (\nu_t^P, \nu_t^T)'$, where $\nu_t^P = \mathrm{B}_{\perp}^{'} e_t$ and $\nu_t^T = \alpha' e_t$ define, respectively, the permanent and transitory shocks.

Equation (7) can be rewritten as

$$\Delta X_t = \Phi + C(L)G^{-1}Ge_t = \Phi + \Omega(L)\nu_t \qquad (8)$$

i.e. as a moving average representation of ΔX_t in terms of the vector of permanent and transitory shocks ν_t, where it is assumed that $C(L)G^{-1} = \Omega(L)$. A standard Choleski decomposition of cov(ν_t) ensures mutual independence of the shocks and brings about the permanent-transitory Gonzalo and Ng decomposition.[8] Impulse

page 22) and (ii) a Wald test for the joint significance of all lagged endogenous VAR variables up to the selected lag.

[8] Details on the identification role played by the Choleski decomposition are provided in Gonzalo and Ng (2001, page 1532).

response and variance decomposition analyses can then be performed in terms of permanent and transitory innovations.

In a cointegrated system we expect a time series to have a large weight in the permanent innovations and a small weight in the transitory innovation if the corresponding error correction term coefficient is small. If, on the other hand, the corresponding error correction coefficient is large, it will have a small weight in the permanent innovations and a large one in the transitory innovation. These hypotheses are corroborated by the decomposition of the forecast error variance of the variables in ΔX_t and have relevant economic implications.

Table III VARIANCE DECOMPOSITION

The table gives the fraction of the total variance in the forecast errors of $\Delta \log R_t$, $\Delta \log \sigma_t$, $\Delta \log M_t$ and $\Delta \log r_t$ that is due either to the three permanent shocks combined or to the transitory shock, orthogonal by construction to the permanent ones. The transitory shock accounts for the majority of the variance of the two year ahead forecast error of $\Delta \log r_t$ in Chile and in Singapore, and a significant fraction of the variation $\Delta \log R_t$ in the remaining countries of the sample where permanent shocks account for the majority of the variance of the remaining variables.[9]

The temporary innovation is an important component of the variance of the forecast error of reserves in most countries and the latter are pivotal in the dynamic process. They adjust but slowly to cointegration disequilibria, however, since they react more to permanent than to transitory shocks, even in the countries where they are the variable upon which the transitory shock has the largest impact. The monetary

[9] We could associate, following Lettau and Ludvigson (2004), the temporary shock with interest rate behavior in Chile. The same reasoning links the three permanent shocks with reserves in Singapore. No clearcut association is possible in the remaining countries, where both permanent and temporary shocks account for the variability of all the variables.

authorities seem to react mostly to shifts in the variables co-moving with the reserves that they perceive as permanent. Furthermore, the relevance of the temporary shock in most reserve forecast error variance decompositions suggests that the latter do not adjust rapidly to the permanent changes in the variability index, imports and interest rates.[10]

3 Analysis of the error correction process

Having assessed that – with the exception of Singapore and Chile - reserves play a relevant role in the adjustment mechanism, we further investigate the nature of this process. More precisely, we set out to analyse the impact of "news" on the speed of adjustment of the stock of reserves to its equilibrium value. We refer to "news" as either reserve misalignments, or shifts in drivers of international capital flows, such as the US Federal effective fund rate, or changes in international competitiveness, measured by real effective exchange rate first differences.

3.1 The impact of positive and negative reserve misalignments

We first assess whether the adjustment of reserves to their equilibrium value is affected by the sign of misalignment in the previous period, i.e. whether their speed

[10] These shifts are not followed by a full adjustment of reserves to their trend value thus generating a transitory component and a temporary cointegrating error. This is not the case of Chile and Singapore, where reserve variation is not affected by the temporary shock, and reserves adjust rapidly to permanent innovations in the remaining variables in the cointegration equation. In these two countries the coefficient of the interest rate could be misleading; it indicates the impact of the permanent component only, which accounts for but a tiny fraction of interest rate variation.

of adjustment depends on a previous accumulation (decumulation) phase with respect to the long run attractor.[11] Asymmetric adjustments, according to Granger and Lee (1989), may be analysed partitioning the error correction term to allow for different speed of adjustment on either side of the attractor.

Thus the estimated error correction term of the previous time period is partitioned into positive and negative deviations from the attractor - according as to whether it lies above or below it - using a Heaviside indicator H_t such that

$$H_t = \begin{cases} 1 & \text{if} \quad EC_{t-1} \geq 0 \\ 0 & \text{if} \quad EC_{t-1} < 0 \end{cases}$$

and $EC_t = \varepsilon_t$ is the residual of the cointegration equation (6).[12]

Equation (4) thus becomes

$$\begin{aligned} \Delta \log R_t &= \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t EC_{t-1} + \beta_2 (1 - H_t) EC_{t-1} + e_{1t} \\ &= \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 EC_{t-1}^{+} + \beta_2 EC_{t-1}^{-} + e_{1t} \end{aligned} \qquad (9)$$

where it is posited that $H_t EC_{t-1} = EC_{t-1}^{+}$ and $(1 - H_t) EC_{t-1} = EC_{t-1}^{-}$, in order to simplify the notation.

Table IV ASYMMETRIC VECM ESTIMATES

Table IV presents evidence of asymmetric behaviour driven by the previous period's accumulation phase. The coefficient estimates, which measure the speed of adjustment for positive and negative deviations from equilibrium, differ greatly across

[11] For a previous attempt to test this hypothesis, see the non linear error correction analysis in Bilson and Frenkel (1979).

[12] It should be noticed that we are not performing here the M-TAR threshold cointegration test of Enders and Syklos (2001). We are simply assessing the impact of changes in the z_{t-1} signal variable on the speed of adjustment.

the countries of the sample. In the case of Argentina, Brazil, Mexico, and Indonesia the coefficient of EC^{-}_{t-1} is statistically significant and (in absolute value) larger than both the EC^{+}_{t-1} and the linear EC_{t-1} ones. Hence, despite the fairly small size of the estimates, whenever international reserves fall below their desired level R^{*} the authorities react more rapidly (the half-life of a deviation from long-run equilibrium is of 6.47 months on average), being anxious to reduce exposure to external shocks, which may be extremely harmful in the case of a shortage of reserves. On the contrary, for Venezuela, Malaysia and Korea, the EC^{+}_{t-1} coefficients exceed in absolute value both the EC^{-}_{t-1} and linear EC_{t-1} coefficients. This finding may be due to the authorities' reaction to the cost of sterilized intervention, traditionally large both in terms of fiscal outlays and of the opportunity cost due to the foregone return on public investment and infrastructure.[13] According to Summers (2006), the latter reaches as much as 1.85 percent of the GDP of the ten leading holders of excess reserves. Mohanty and Turner (2006) point out that in the early 1990s large capital inflows brought about an increase in the annual cost of sterilized intervention by 0.25-0.50 percent of GDP in several Latin American countries.

3.2 The role of US monetary stance and external competitiveness on the speed of adjustment

Empirical evidence of asymmetries in the sluggish process of adjustment leaves open the question of the relevance of news to the error correction mechanism. The issue is

[13] The half-life of a deviation from long run equilibrium is, in these countries, somewhat longer and reaches 9.85 months on average

whether there are factors able to influence the process of bridging the gap between actual reserves and their equilibrium level.

During the financial crises of the 1980s and 1990s many emerging market economies had to face sudden capital reversals and costly currency depreciations with inadequate stocks of foreign reserves. It is not surprising that the local authorities should feel the need to self-insure against crises triggered by unexpected capital outflows, especially in the presence of negative discrepancies between actual and desired reserves, and that factors influencing international capital and trade flows should affect the speed of adjustment of reserves to their optimal level. In this section we analyse the impact of news both on international liquidity, using as a proxy changes in the US Federal effective fund interest rate, and on relative competitiveness, measured by changes in the real effective exchange rate ($REER_t$).[14]

3.2.1 The impact of US monetary policy

A number of studies suggest that capital flows are driven by common international factors. As shown by Calvo et al. (1996) and Mody et al. (2001), among others, shifts in US monetary policy influence emerging markets' financial liquidity. A tight US monetary policy makes investment in these countries less attractive, raising debt price. The corresponding increase in the interest differential results in cross border financial outflows and eventually costly currency depreciations that may pass-through to inflation.[15]

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[14] Both the Fed fund and the national $REER_t$ time series are I(1). Their first differences are I(0) and may be interpreted as "news".

[15] See Arora and Cerisola (2001) and Uribe and Yue (2003). It is also believed that US monetary policy plays a significant role in triggering financial and banking crises since a rise in

The extent to which the local monetary authorities react to changes in the US interest rate depends, in principle, both on the nature of the exchange rate arrangements and on the monetary policy framework. Under a pegged exchange rate regime the reaction should be strong, in order to avoid the insurgence of a risk premium. Frankel (1999), however, found that also in free-floating countries (such as Brazil and Mexico) a positive shift in the Fed fund rate brings about a more than proportional increase in the domestic interest rate. The latter is due to the considerable effect of interest rate differentials on capital outflows and to the large premium pricing devaluation and default risks. Under an inflation targeting regime, a depreciation of the national currency may put price stability under pressure. The authorities' fear of floating may thus increase the speed of accumulation of foreign reserves, which may be sterilized in order to prevent an increase in the money supply.

According to the rationale mentioned above, if a US monetary policy tightening has an adverse effect on emerging markets' financial stability, we would expect a positive impact of the first difference of the US Fed fund effective rate on the speed of reserve adjustment.

On the contrary, a positive impact of US expansionary monetary policy on the speed of accumulation is plausibly explained by the fact that a large interest rate differential in favour of the home country may trigger huge capital inflows and hence a rise in reserves that is likely to be considered a threat to price stability -especially in the case of overstocking and under an inflation targeting regime.[16]

The Heaviside indicator function for the US monetary policy is defined as

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industrial country interest rates worsens the conditions for the access of emerging markets to offshore funds.

[16] Moreover, if the authorities intend to sterilize the increase in reserves, they will face higher opportunity costs as the interest rate differential augments. Evidence on this point is provided for most of the developing countries by Mohanty and Scatigna (2005).

$$H_t^{FED} = \begin{cases} 1 & \text{if} \quad \Delta i_{USAt-1} \geq 0 \\ 0 & \text{if} \quad \Delta i_{USAt-1} < 0 \end{cases}$$

where Δi_{USAt-1} is the first difference of the Fed fund effective rate. The impact of the US monetary stance is then measured with the help of a modified version of equation (9), where H_t is replaced by H_t^{FED} . $H_t^{FED} EC_{t-1} = FEC_{t-1}^{+}$ and $(1 - H_t^{FED}) EC_{t-1} = FEC_{t-1}^{-}$ are then the error correction terms corresponding, respectively, to periods of restrictive and expansionary US monetary policy.

Table V ASYMMETRIC ADJUSTMENT AND EXTERNAL FACTORS IN EQ. (9)

Table V shows that for most countries one of the two asymmetric adjustment coefficients is statistically significant and larger - in absolute value - than the corresponding symmetric coefficient estimate, set forth in the first column of table II. The adjustment is more rapid in the case of a restrictive US monetary stance (the half-life time of the adjustment being 9.7 months, on average) than in the case of an expansionary one. As for Singapore and Chile, whose VECM estimates offer clear evidence of the role of interest rates in the adjustment mechanism, US monetary policy news affects interest rate changes and exerts no effect on reserves (see the last two rows in Table V). In both countries monetary policy has a primary objective in price stability while keeping exchange rate volatility under control.[17] Interest rates seem to react rapidly to Federal Reserve restrictive actions exerting an adjustment pressure on reserves via the long run relationship (6).

To investigate the role of the previous period's positive or negative discrepancies between actual and equilibrium (desired) reserves on the impact that news exert on the speed of adjustment, the Heaviside indicator was further modified, multiplying H_t^{FED} and H_t . The following relationship was estimated

[17] For more details on monetary policy and sterilization of reserves see Tee (2005) for Singapore and Cifuentes and Desormeaux (2005) for Chile.

$$\Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t^{FED} EC_{t-1}^{+} + \beta_2 (1 - H_t^{FED}) EC_{t-1}^{-} + \beta_3 (1 - H_t^{FED}) EC_{t-1}^{+} + \beta_4 H_t^{FED} EC_{t-1}^{-} + e_{1t} \quad (10)$$

Table VI INTERACTION BETWEEN EXTERNAL FACTORS AND RESERVE OVER/UNDERSTOCKING (β_i COEFFICIENTS)

The β_i coefficient estimates set forth in table VI are highly informative. The local authorities seem to react more when reserves lie below their equilibrium level, as suggested by the significant increase in the speed of adjustment when the error correction term EC_t is negative. Moreover, for most of the countries in the sample, the absolute value of the coefficients is much larger than in the estimates shown in table V – a finding that may be interpreted as evidence of a precautionary attitude. A qualitative summary of the estimates of β_i in equation (10) is presented in the matrix below.

Countries for which the error correction coefficient in the equation (10) estimates is significantly different from zero at the 5 percent level		
	EC_{t-1}^{+}	EC_{t-1}^{-}
$\Delta FED_{t-1} \geq 0$		Argentina, Mexico, Korea, Malaysia
$\Delta FED_{t-1} < 0$		Brazil, Singapore, Indonesia

3.2.2 The impact of changes in competitiveness

The mercantilistic rationale behind reserve holdings is consistent with the decisions by Central Bankers to purchase foreign currency during a period of upward pressure on the domestic exchange rate and conduct rather limited interventions on the downside. This policy may prevent the deterioration of national competitiveness in accordance with a deep-rooted mercantilist desire to maintain an undervalued exchange rate. This explanation agrees with the suggestion of Dooley et al. (2003) and Genberg et al. (2005) that emerging market countries build up reserves in order to support their exports. A sensible development strategy might then require a distortion in the real exchange rate in order to channel domestic investment towards export industries, and a process of reserve accumulation, which would appear sub-optimal, is in reality an element in an optimal investment strategy.

We thus expect countries with a marked mercantilistic attitude to increase the speed of accumulation in times of real effective appreciation of the national real exchange rate ($REER_t$). This line of reasoning offers only a partial view of the mercantilistic motive, which is better suited to economies exporting manufactured goods. Eichengreen (2004) argues that the export led growth model, while successful in the past, may now have a diminishing explanatory power in the case of most Asian economies. The knowledge spillovers have moved away from traditional traded good to areas such as software development, back office services and financial intermediation. This process requires balanced investments in both traded and non-traded sectors (including human capital), which do not necessarily benefit from a long lasting currency depreciation.

Along these lines, for commodity exporting countries an increase in the terms of trade (corresponding to a real appreciation) is by far more important than other forms of price competition. This could imply that in periods of depreciation the increase in deficit (due to the increasing value of imports in terms of exports) and the pressure on foreign debt will lead the local authorities to change the rate of reserve accumulation, in order to reduce external vulnerability.

The Heaviside indicator is then defined as follows

$$H_t^{REER} = \begin{cases} 1 & \text{if} \quad \Delta REER_{t-1} \geq 0 \\ 0 & \text{if} \quad \Delta REER_{t-1} < 0 \end{cases}$$

where $\Delta REER_{t-1}$ is the $t-1$ year on year percentage first difference of the national real effective exchange rate. In the asymmetric adjustment equation (9), where now $H_t = H_t^{REER}$, let $H_t^{REER} EC_{t-1} = REC_{t-1}^{+}$ be the error correction term corresponding to a real effective appreciation and $(1-H_t^{REER})EC_{t-1} = REC_{t-1}^{-}$ the term corresponding to a phase of real effective depreciation.

Table V shows that for Argentina, Brazil and Korea the coefficient of REC_{t-1}^{+} in equation (9) is statistically significant and indicates that the speed of adjustment increases with appreciation of the real effective exchange rate. The opposite holds true for Chile, Mexico, Venezuela, Indonesia, Malaysia and Singapore, where the coefficient of REC_{t-1}^{-} is statistically different from zero at the 5 percent significance level. This finding is not surprising since the group is made up of economies characterised by a huge share of commodities in exports and/or a high degree of foreign presence in the real and financial sectors. The role of previous period positive or negative reserve discrepancies on the impact of competitiveness related news was analysed using the Heaviside indicator obtained multiplying H_t^{REER} by H_t. Equation (9) then reads as

$$\begin{aligned} \Delta \log R_t = \quad & \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t^{REER} EC_{t-1}^{+} + \beta_2 (1-H_t^{REER}) EC_{t-1}^{-} \\ & + \beta_3 (1-H_t^{REER}) EC_{t-1}^{+} + \beta_4 H_t^{REER} EC_{t-1}^{-} + e_{1t} \end{aligned} \qquad (10')$$

The error correction coefficients estimates are set out in the last two columns of table VI. Here too, for the majority of countries, the dimension of the coefficients is much larger in absolute value than in the case of the unadjusted parameterization of table

V. In the matrix below the estimates are presented qualitatively, according to their significance at the 5 percent level. Most local authorities seem to react more rapidly in the presence of an excess demand of reserves, as suggested by the significant increase in the speed of adjustment when the error correction term EC_t is negative.

Countries for which the error correction coefficient in the equation (10') estimates is significantly different from zero at the 5 percent level		
	EC^{+}_{t-1}	EC^{-}_{t-1}
$\Delta REER_{t-1} \geq 0$		Argentina, Brazil, Korea
$\Delta REER_{t-1} < 0$	Malaysia Singapore	Chile, Mexico, Venezuela, Indonesia

As for Malaysia and Singapore – small open economies with a sizeable foreign presence – they are characterized by a pegged or dirty floating exchange rate regime. Thus, assuming that the $REER_t$ changes are mainly due to changes of the inflation differential, the local authorities may be willing to control the stock of money and medium term inflation getting rid of reserves when in excess (selling foreign assets in exchange for local currency), in order to reduce the monetary base and exert a real appreciation pressure.

4 Conclusion

Foreign exchange reserve accumulation has recently reached dimensions that are, at first sight, difficult to explain on the basis of standard cost opportunity considerations. We find, however, that the traditional buffer stock precautionary model continues to provide a valid explanation of optimal reserve demand on the part of many emerging market economies. The cointegration approach implemented in the paper draws a distinction between factors entering the long run equilibrium demand for reserves and those operating only in the short run. Furthermore, it can be inferred from the variance decomposition analysis of permanent and transitory shocks to the equilibrium relationships that, in all but two cases, the stock of international reserves is the variable which brings about the readjustment over the long term. The other domestic variables seem to influence long run restocking decisions only if their changes are perceived as permanent by the local authorities.

Along with domestic variables, such as the short term interest rate and the variability of the balance of payments, recent strands of the literature point to mercantilistic real exchange rate manipulation and fear of floating as relevant drivers in reserve management policies. Indeed, we established that these factors may affect the speed of adjustment of current reserves to their optimal long run level using estimation techniques that are compatible with cointegration analysis. In particular, news about "international liquidity conditions" and "competitiveness" – represented, respectively, by US monetary stance and domestic real effective exchange rate changes - has been shown to raise the speed of adjustment of reserves to their optimal (desired) level. Asymmetries entering the adjustment mechanism, either as previous over/under stocking of foreign reserves or as expansionary/restrictive US monetary stance or, finally, as over/under valuation of the real effective exchange rate, play a significant role. Indeed, there is strong evidence that international assets are stocked to reduce vulnerability to external shocks and avoid loss of competitiveness even if, in the long run, domestic precautionary motives are still predominant.

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APPENDIX I

Time series are monthly and cover the period between January 1985 and July 2006.

Reserves excluding gold are series l1da quoted in US dollars from the IMF International Financial Statistics Data Base. International reserves do not include gold because of valuation problems and the modest amount of the precious metal in EME reserve stocks.

Interest rates are the money market rates in series 60b for Argentina, Brazil, Indonesia, Korea, Malaysia, Mexico, and Singapore. The deposit rate in series 60l was used for Chile. They are all from the IMF International Financial Statistics Data Base.

Imports are series 71da (cif quoted in US dollars) from the IMF International Financial Statistics Data Base.

Real effective exchange rates are from the IMF International Financial Statistics Data Base.

US short interest rate, the Federal fund effective interest rate is drawn from the Federal Reserve database.

APPENDIX II

Table A.I – UNIT ROOT TESTS

NP*			NP		
Reserves			**Domestic Short Term Interest Rate**		
Argentina C, t, lag 1	-1.62	I(1)	Argentina C, t, lag 7	-1.59	I(1)
Brazil C, lag 1	0.01	I(1)	Brazil C, t, lag 9	-2.07	I(1)
Chile C, lag 13	0.85	I(1)	Chile C, t, lag 10	-2.11	I(1)
Mexico C, t, lag 0	-2.25	I(1)	Mexico C, t, lag 6	-2.26	I(1)
Venezuela C, t, lag 1	-1.51	I(1)	Venezuela C, lag 14	-1.40	I(1)
Indonesia C, t, lag 0	-1.85	I(1)	Indonesia C, lag 1	-2.17	I(1)
Korea C, t, lag 8	-1.96	I(1)	Korea C, t, lag 1,	-1.69	I(1)
Malaysia C, t, lag 2	-1.82	I(1)	Malaysia C, lag 1	-1.27	I(1)
Singapore C, t, lag 8	-0.91	I(1)	Singapore C, lag 2	-1.22	I(1)
Reserve Volatility			**Value of Imports**		
Argentina C, t, lag 14	-2.80	I(1)	Argentina C, t, lag 3	-1.73	I(1)
Brazil C, lag 0	0.15	I(1)	Brazil C, t, lag 13	-2.09	I(1)
Chile C, t, lag 0	-1.84	I(1)	Chile C, t, lag 12	-1.81	I(1)
Mexico C, lag 0	-1.17	I(1)	Mexico C, t, lag 12	-1.79	I(1)
Venezuela C, t, lag 0	-2.02	I(1)	Venezuela C, t, lag 1	-3.43	I(0)
Indonesia C, t, lag 0	1.77	I(1)	Indonesia C, t, lag 5	-1.60	I(1)
Korea C, t, lag 0	-2.04	I(1)	Korea C, t, lag 12	-2.87	I(1)
Malaysia C, t, lag 3	-1.33	I(1)	Malaysia C, t, lag 13	-2.22	I(1)
Singapore C, t, lag 0	-2.12	I(1)	Singapore C, t, lag 15	-2.13	I(1)

Notes. *: Ng Perron (2001) unit root test. The GLS-detrended autoregressive spectral density estimator of the frequency zero spectrum uses the modified AIC to select the number of lags. Critical values with constant, C, and trend, t: -3.42 (1 percent), -2.91(5 percent); with constant without trend: -2.58 (1 percent), -1.98 (5 percent).

APPENDIX III

Table A.II - JOHANSEN COINTEGRATION TESTS

TRACE TEST STATISTICS

List of variables in the VAR: *Log (Reserves) = log* R_t *; Log (Volatility of Reserves)= log* σ_t *; Log (Imports) = log* M_t *;*

Log (Domestic Interest Rate)= log r_t

	Hypothesized No. of Cointegration Relationships	Trace Statitics	0.05 percent Critical Value	Deterministic Trend Assumption	No. of Lags in VAR
Argentina	None	65.50*	63.88	Restricted linear deterministic trend	12
	at most 1	33.48	42.91		
	at most 2	12.91	25.87		
	at most 3	3.90	12.51		
Brazil	None	60.54*	54.08	Restricted constant	4
	at most 1	34.33	35.19		
	at most 2	15.94	20.26		
	at most 3	3.44	9.16		
Chile	None	52.36*	47.86	Linear deterministic trend	6
	at most 1	23.13	29.80		
	at most 2	10.33	15.49		
	at most 3	3.18	3.84		
Mexico	None	86.22*	63.88	Restricted linear deterministic trend	11
	at most 1	40.94	42.91		
	at most 2	14.57	25.87		
	at most 3	4.54	12.51		
Venezuela	None	33.07*	29.80	Linear deterministic trend	12
	at most 1	15.22	15.49		
	at most 2	2.75	3.84		
Indonesia	None	59.55*	54.08	Restricted constant	3
	at most 1	34.13	35.19		
	at most 2	18.07	20.26		
	at most 3	7.04	9.16		
Korea	None	55.98*	47.86	Linear deterministic trend	25
	at most 1	27.66	29.80		
	at most 2	11.18	15.49		
	at most 3	2.27	3.84		
Malaysia	None	83.22*	63.88	Restricted linear deterministic trend	8
	at most 1	41.16	42.91		
	at most 2	15.84	25.87		
	at most 3	3.50	12.51		
Singapore	None	66.51*	47.86	Linear deterministic trend	1
	at most 1	28.11	29.80		
	at most 2	13.42	15.49		
	at most 3	3.47	3.84		

Notes. *: denotes rejection of the null hypothesis at the 5 percent level .

Table I – DOLS COINTEGRATION EQUATION ESTIMATES

$$\log R_t - \rho_0 - \rho_t t - b_1 \log \sigma_t - b_2 \log M_t - b_3 \log r_t = \varepsilon_t \quad (6)$$

	$-\rho_0$	$-\rho_1$	$-b_1$	$-b_2$	$-b_3$	**N. of leads and lags in the DOLS**
Argentina	-2.71 (0.38)	-0.004 (0.001)	-0.26 (0.05)	-0.62 (0.04)	0.07 (0.01)	3
Brazil	1.96 (1.29)	--	-0.21 (0.10)	-1.28 (0.15)	-0.03 (0.04)	3
Chile	0.18 (0.51)	--	-0.39 (0.16)	-1.00 (0.11)	0.04 (0.04)	6
Mexico	-10.38 (1.96)	-0.01 (0.003)	0.01 (0.13)	0.13 (0.28)	0.22 (0.09)	3
Venezuela	-4.71 (0.58)	--	-0.84 (0.09)	--	0.32 (0.07)	3
Indonesia	0.48 (0.73)	--	-0.81 (0.07)	-0.71 (0.09)	0.26 (0.05)	3
Korea	3.85 (0.84)	--	-0.51 (0.10)	-1.32 (0.13)	0.62 (0.07)	3
Malaysia	-5.70 (0.26)	-0.01 (0.00)	-0.75 (0.04)	0.22 (0.07)	-0.04 (0.03)	3
Singapore	-1.54 (0.31)	--	-0.30 (0.03)	-0.82 (0.05)	0.12 (0.02)	5

Table II – ERROR CORRECTION COEFFICIENTS

$$\Delta X_t = \Phi + \mathrm{B}\alpha' X_{t-1} + \sum_{j=1}^{k} \Gamma_j \Delta X_{t-j} + e_t \qquad (5)$$

	Error correction coefficients of the equation of				VAR order VAR Lag Exclusion Wald Test	Cond. Het. Q-stat.+	Joint VAR Serial Corr. LM Test°
	$\Delta \log R_t$	$\Delta \log \sigma_t$	$\Delta \log M_t$	$\Delta \log r_t$			
Argentina	-0.122* (0.046)	0.061 (0.081)	-0.013 (0.046)	-0.654 (0.439)	14 27.58 [0.03]	5.0 (lag 1) [0.02] 17.0(lag 3) [0.00]	19.6 (lag 1) [0.24] 14.7 (lag 3) [0.55]
Brazil	-0.040* (0.017)	0.013 (0.023)	0.079* (0.025)	-0.140 (0.119)	4 36.96 [0.002]	0.1 (lag 1) [0.69] 0.7 (lag 3) [0.87]	16.2 (lag 1) [0.44] 27.5 (lag 3) [0.04]
Chile	-0.0184 (0.012)	-0.006 (0.023)	0.065 (0.034)	-0.182* (0.084)	5 29.15 [0.02]	1.0 (lag 1) [0.31] 1.3 (lag 3) [0.73]	18.2 (lag 1) [0.31] 19.6 (lag 3) [0.24]
Mexico	-0.062* (0.027)	0.058* (0.028)	0.314* (0.018)	-0.054 (0.032)	8 30.08 [0.02]	17.1(lag 1) [0.00] 17.2(lag 3) [0.00]	17.6 (lag 1) [0.35] 24.7 (lag 3) [0.07]
Venezuela	-0.061* (0.026)	0.006 (0.020)		-0.131* (0.050)	12 8.64 [0.47]	1.1 (lag 1) [0.29] 4.4 (lag 3) [0.22]	5.7 (lag 1) [0.77] 13.7 (lag 3) [0.13]
Indonesia	-0.035* (0.012)	0.018 (0.019)	0.026 (0.034)	-0.093* (0.045)	3 28.30 [0.03]	2.8 (lag 1) [0.08] 19.7 (lag 3) [0.00]	12.4 (lag 1) [0.71] 20.9 (lag 3) [0.18]
Korea	-0.023* (0.012)	-0.148 (0.026)	0.021 (0.023)	0.013 (0.022)	20 23.70 [0.10]	1.4 (lag 1) [0.23] 6.5 (lag 3) [0.09]	12.0 (lag 1) [0.74] 15.0 (lag 3) [0.53]
Malaysia	-0.121* (0.056)	0.143* (0.050)	-0.294* (0.094)	0.167 (0.131)	8 26.47 [0.05]	0.5 (lag 1) [0.49] 1.2 (lag 3) [0.75]	28.2 (lag 1) [0.03] 40.8 (lag 3) [0.00]
Singapore	-0.013 (0.012)	0.176* (0.060)	0.130* (0.052)	-0.294* (0.102)	1 74.88 [0.00]	2.5 (lag 1) [0.12] 3.4 (lag 3) [0.34]	31.4 (lag 1) [0.01] 25.7 (lag 3) [0.06]

Notes. *: significantly different from zero at the 5 percent level; + : Ljung Box Q-statistic test for conditional heteroskedasticity; °: Johansen (1995) VECM residual autocorrelation LM test; standard errors are in parentheses and probability values in square brackets.

Table III – VARIANCE DECOMPOSITION

		$(\Delta \log R_{t+h} - E_t \Delta \log R_{t+h})$		$(\Delta \log \sigma_{t+h} - E_t \Delta \log \sigma_{t+h})$		$(\Delta \log M_{t+h} - E_t \Delta \log M_{t+h})$		$(\Delta \log r_{t+h} - E_t \Delta \log r_{t+h})$	
	H	**P**	**T**	**P**	**T**	**P**	**T**	**P**	**T**
Argentina	**1**	28.80	71.20	100.00	0.00	100.00	0.00	100.00	0.00
	2	28.52	71.48	99.49	0.51	99.53	0.47	42.40	57.60
	3	32.21	67.79	99.39	0.61	99.52	0.47	70.69	29.30
	4	33.83	66.17	99.39	0.61	99.37	0.63	72.28	27.72
	24	69.01	30.99	98.88	1.12	97.57	2.43	86.36	13.64
Brazil	**1**	10.06	89.94	100.00	0.00	100.00	0.00	100.00	0.00
	2	9.91	90.09	96.63	3.37	99.32	0.68	93.77	6.23
	3	9.92	90.08	95.96	4.04	98.97	1.03	93.75	6.25
	4	10.26	89.74	95.53	4.47	98.76	1.24	93.64	6.35
	24	10.67	89.33	95.39	4.61	94.67	5.33	93.40	6.60
Chile	**1**	100.00	0.00	100.00	0.00	100.00	0.00	27.56	72.44
	2	99.58	0.42	99.98	0.02	99.97	0.03	33.73	66.27
	3	99.31	0.69	99.93	0.07	98.15	1.85	37.30	62.70
	4	99,27	0.73	99.93	0.07	97.01	2.99	44.40	55.60
	24	98.81	1.19	99.62	0.38	97.81	2.19	46.48	53.52
Mexico	**1**	73.25	26.75	100.00	0.00	100.00	0.00	100.00	0.00
	2	72.80	27.19	99.99	0.01	99.99	0.01	97.87	2.13
	3	73.26	26.74	99.77	0.23	99.99	0.01	97.34	2.66
	4	74.35	25.65	99.77	0.23	99.85	0.15	97.60	2.40
	24	77.04	22.96	98.19	1.81	99.82	0.18	97.23	2.77
Venezuela	**1**	77.98	22.02	100.00	0.00			100.00	0.00
	2	77.43	22.57	92.24	7.76			96.61	3.39
	3	72.22	22.78	92.28	7.72			95.57	4.43
	4	77.29	22.71	92.35	7.65			95.54	4.46
	24	77.52	22.48	92.18	7.82			94.58	5.42
Indonesia	**1**	76.04	23.96	100.00	0.00	100.00	0.00	100.00	0.00
	2	76.00	24.00	98.15	1.84	99.78	0.22	92.94	7.06
	3	75.95	24.04	97.48	2.52	99.69	0.30	93.14	6.86
	4	75.99	24.01	97.47	2.53	99.64	0.36	93.34	6.66
	24	76.35	23.64	93.27	6.73	94.33	5.67	88.08	11.92
Korea	**1**	76.66	23.34	100.00	0.00	100.00	0.00	100.00	0.00
	2	76.91	23.09	98.94	1.06	99.97	0.03	99.71	0.28
	3	76.64	23.36	98.90	1.10	98.73	1.27	98.77	1.23
	4	76.96	23.04	98.78	1.21	98.34	1.66	97.75	2.25
	24	79.24	20.76	97.33	2.66	97.55	2.45	95.49	4.51
Malaysia	**1**	62.03	37.97	100.00	0.00	100.00	0.00	100.00	0.00
	2	64.79	35.21	82.64	17.36	99.95	0.06	99.04	0.95
	3	66.58	33.42	82.78	17.22	99.55	0.45	98.75	1.25
	4	67.46	32.54	82.95	17.05	99.49	0.51	96.94	3.06
	24	70.40	29.60	83.35	16.65	99.18	1.82	93.56	6.42
Singapore	**1**	100.00	0.00	100.00	0.00	100.00	0.00	0.05	99.95
	2	98.99	1.01	11.68	88.32	57.37	43.63	0.58	99.42
	3	98.81	1.19	11.60	88.40	52.61	47.39	0.62	99.38
	4	98.79	1.21	11.61	88.39	51.77	48.23	0.62	99.38
	24	98.78	1.22	11.61	88.39	51.66	48.34	0.62	99.38

Notes. Fraction of the variance in the H-month ahead forecast error due to innovations in the permanent shocks P and transitory shock T.

Table IV – ASYMMETRIC VECM ESTIMATES

$$\Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t EC_{t-1} + \beta_2 (1 - H_t) EC_{t-1} + e_{1t}$$
$$= \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 EC_{t-1}^{+} + \beta_2 EC_{t-1}^{-} + e_{1t} \qquad (9)$$

	β_1	β_2	VAR order VAR Lag Exclusion Wald Test	Cond. Het. Q-stat.+	Joint VAR Serial Corr. LM Test°
Argentina	-0.047 (0.064)	-0.196* (0.063)	14 25.60 [0.06]	6.2 (lag 1) [0.01] 13.2 (lag 1) [0.00]	21.2 (lag 1) [0.17] 18.5 (lag 3) [0.30]
Brazil	-0.024 (0.024)	-0.063* (0.030)	4 38.02 [0.00]	0.2 (lag 1) [0.65] 0.7 (lag 3) [0.87]	22.5 (lag 1) [0.13] 28.1 (lag 3) [0.03]
Chile	-0.023 (0.021)	-0.015 (0.017)	5 29.27 [0.02]	1.0 (lag 1) [0.32] 1.2 (lag 3) [0.76]	21.6 (lag 1) [0.15] 19.1 (lag 3) [0.26]
Mexico	0.004 (0.033)	-0.195* (0.048)	8 28.74 [0.02]	14.4 (lag 1) [0.00] 14.4 (lag 3) [0.00]	16.7 (lag 1) [0.40] 25.4 (lag 3) [0.06]
Venezuela	-0.084* (0.042)	-0.040 (0.039)	12 8.64 [0.47]	1.2 (lag 1) [0.27] 4.5 (lag 3) [0.21]	9.7 (lag 1) [0.37] 19.2 (lag 3) [0.02]
Indonesia	0.006 (0.019)	-0.078* (0.019)	3 29.39 [0.02]	3.2 (lag 1) [0.07] 27.4 (lag 3) [0.00]	10.3 (lag 1) [0.85] 23.7 (lag 3) [0.09]
Korea	-0.039* (0.019)	0.001 (0.024)	20 22.19 [0.14]	2.5 (lag 1) [0.11] 5.9 (lag 3) [0.12]	12.4 (lag 1) [0.72] 18.3 (lag 3) [0.31]
Malaysia	-0.152* (0.077)	-0.099 (0.068)	8 25.16 [0.07]	0.4 (lag 1) [0.54] 0.9 (lag 3) [0.83]	29.4 (lag 1) [0.02] 34.0 (lag 3) [0.01]
Singapore	-0.028 (0.019)	-0.000 (0.018)	1 75.29 [0.00]	2.9 (lag 1) [0.09] 3.7 (lag 3) [0.29]	30.4 (lag 1) [0.02] 25.9 (lag 3) [0.06]

Notes. + : Ljung Box Q-statistic test for conditional heteroskedasticity; °: Johansen (1995) VECM residual autocorrelation LM test; standard errors are in parentheses and probability values in square brackets; *: significantly different from zero at the 5 percent level.

Table V - ASYMMETRIC ADJUSTMENT AND EXTERNAL FACTORS IN EQ. (9)

$$\Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 FEC_{t-1}^{+} + \beta_2 FEC_{t-1}^{-} + e_{1t}, \quad \Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 REC_{t-1}^{+} + \beta_2 REC_{t-1}^{-} + e_{1t}$$

$$\Delta \log r_t = \varphi + \sum_{j=1}^{k} \Gamma_j^4 \Delta X_{t-j} + \beta_1^* FEC_{t-1}^{+} + \beta_2^* FEC_{t-1}^{-} + e_{4t}$$

	Impact of the US monetary stance		Impact of the real effective exchange rate	
	FEC_{t-1}^{+}	FEC_{t-1}^{-}	REC_{t-1}^{+}	REC_{t-1}^{-}
Dependent variable	$\Delta \log R_t$		$\Delta \log R_t$	
Argentina	-0.189* (0.054)	-0.045 (0.057)	-0.145* (0.058)	-0.104 (0.054)
Brazil	-0.017 (0.023)	-0.060* (0.022)	-0.037* (0.020)	-0.027 (0.024)
Chile	-0.015 (0.014)	-0.017 (0.021)	-0.002 (0.014)	-0.047* (0.020)
Mexico	-0.068* (0.033)	-0.055 (0.035)	-0.041 (0.033)	-0.093* (0.038)
Venezuela	-0.022 (0.024)	-0.059* (0.032)	-0.009 (0.024)	-0.095* (0.035)
Indonesia	-0.010 (0.019)	-0.051* (0.015)	-0.010 (0.016)	-0.062* (0.017)
Korea	-0.032* (0.015)	-0.001 (0.021)	-0.039* (0.016)	-0.003 (0.018)
Malaysia	-0.147* (0.067)	-0.094 (0.068)	-0.053 (0.070)	-0.174* (0.064)
Singapore	-0.010 (0.017)	-0.017 (0.018)	0.002 (0.016)	-0.034* (0.019)
Dependent variable	$\Delta \log r_t$			
Singapore	-0.509* (0.131)	-0.033 (0.143)		
Chile	-0.202* (0.098)	-0.128 (0.141)		

Notes. $H_t^{FED} EC_{t-1} = FEC_{t-1}^{+}$, $(1 - H_t^{FED}) EC_{t-1} = FEC_{t-1}^{-}$, $H_t^{REER} EC_{t-1} = REC_{t-1}^{+}$ and $(1 - H_t^{REER}) EC_{t-1} = REC_{t-1}^{-}$, where EC_t is the error correction term, $H_t^{FED} = \begin{cases} 1 & \text{if} \quad \Delta i_{USAt-1} \geq 0 \\ 0 & \text{if} \quad \Delta i_{USAt-1} < 0 \end{cases}$ and $H_t^{REER} = \begin{cases} 1 & \text{if} \quad \Delta REER_{t-1} \geq 0 \\ 0 & \text{if} \quad \Delta REER_{t-1} < 0 \end{cases}$; *: significantly different from zero at the 5 percent level.

Table VI - INTERACTION BETWEEN EXTERNAL FACTORS AND RESERVE OVER/UNDERSTOCKING (β_i COEFFICIENTS)

$$\Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t^{FED} EC_{t-1}^{+} + \beta_2 (1 - H_t^{FED}) EC_{t-1}^{-} + \beta_3 (1 - H_t^{FED}) EC_{t-1}^{+} + \beta_4 H_t^{FED} EC_{t-1}^{-} + e_{1t} \quad (10)$$

$$\Delta \log R_t = \varphi + \sum_{j=1}^{k} \Gamma_j^1 \Delta X_{t-j} + \beta_1 H_t^{REER} EC_{t-1}^{+} + \beta_2 (1 - H_t^{REER}) EC_{t-1}^{-} + \beta_3 (1 - H_t^{REER}) EC_{t-1}^{+} + \beta_4 H_t^{REER} EC_{t-1}^{-} + e_{1t} \quad (10')$$

		EC_{t-1}^{+}	EC_{t-1}^{-}			EC_{t-1}^{+}	EC_{t-1}^{-}
H_t^{FED}	Argentina	-0.134 (0.082)	-0.199* (0.064)	H_t^{REER}	Argentina	0.050 (0.076)	-0.285* (0.081)
	Mexico	-0.025 (0.037)	-0.137* (0.060)		Brazil	-0.022 (0.028)	-0.060* (0.033)
	Korea	-0.019 (0.019)	-0.059* (0.029)		Korea	-0.022 (0.020)	-0.082* (0.040)
	Malaysia	-0.149 (0.092)	-0.152* (0.082)				
$(1 - H_t^{FED})$	Brazil	-0.034 (0.029)	-0.085* (0.032)	$(1 - H_t^{REER})$	Chile	-0.018 (0.027)	-0.082* (0.030)
	Chile	-0.010 (0.029)	-0.024 (0.025)		Mexico	0.043 (0.057)	-0.171* (0.046)
	Venezuela	-0.071 (0.041)	-0.031 (0.051)		Venezuela	-0.012 (0.046)	-0.233* (0.061)
	Indonesia	-0.007 (0.022)	-0.089* (0.021)		Indonesia	-0.007 (0.024)	-0.115* (0.024)
	Singapore	-0.031 (0.028)	-0.052* (0.025)		Malaysia	-0.214* (0.086)	-0.105 (0.082)
					Singapore	-0.040* (0.022)	-0.019 (0.036)

Note. *: significantly different from zero at the 5 percent level.