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The Relationship between Stock Returns and the Foreign Exchange Rate: the ARDL Approach

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Abstract

This study employs the ARDL cointegration approach in order to examine the impact of financial liberalization on the relationships between the exchange rate and share market performance in China. We discovered that cointegration has existed between the Shanghai A Share Index and the exchange rate of the renminbi against the US dollar and Hong Kong dollar since 2005, when the Chinese exchange rate regime became a flexible, managed, floating system. We found that both the exchange rate and money supply influenced stock price, with a positive correlation. We further show that the money supply increase was largely caused by a huge “hot money” inflow from other countries in recent years. After local currency appreciation, hot money, followed by the money supply increase, pushed the market into a high level, based on expectations regarding the local currency’s further appreciation.

Key words: share market index, exchange rate, ARDL cointegration, causality, China

JEL classification: F31, F41, G20, N27

Introduction

Over recent years an increasingly close relationship has emerged between national stock markets and their respective exchange rates as a result of the rising flows of capital between international financial markets. The analysis of the interactions between stock prices and exchange rates is particularly interesting and important from the perspective of emerging markets such as China (one of the largest), for several reasons.

First, in recent years, Chinese markets have displayed qualities similar to other emerging markets at the forefront of financial market liberalization. China abandoned the fixed peg exchange rate regime on 21 July 2005. The Chinese yuan had been pegged at 8.3 to one US dollar for the previous 10 years, and foreign governments had been pressuring China to move towards a more flexible, managed, floating exchange rate. Officially, China shifted to what some called a more “mechanical” version of a currency basket regime (that is, it kept the trade weighted exchange rate within a certain band as a goal in and of itself).¹

Second, Chinese stock markets have recently developed and opened themselves to external investors by introducing qualified foreign institutional investors (QFII) in

¹ Contrary to the public pronouncement of the Chinese authorities that the currency is based on a currency basket, recent empirical studies suggest the de facto regime appears to be a soft peg to the US dollar with the IMF classifies China under “other conventional fixed peg arrangements”. See Shah, Zeileis and Patnaik (2006) and Ogawa and Sakane (2006) for empirical validation.

2003, and opened channels to local investors to allow them to invest in external markets by implementing qualified domestic institutional investors (QDII) in 2005. However, both the liberalization of the foreign exchange market and the opening up of local stock markets became more prone to contagion effects once their integration in the global financial markets intensified. Between July 2005 and December 2009 (while we were writing this paper) the Chinese yuan appreciated from 8.3 to 6.3 (31 percent), while the Shanghai A Share Index rapidly rose from 1,007 to 6,100 in October 2007, representing its highest peak, and then fell to approximately 3,379 in December 2009. Nearly 50 percent of this drop was mainly due to the international stock market correction triggered by the American sub-prime loan crisis. However, overall, the increase in the Shanghai stock index was caused by the appreciation of the Chinese yuan from the point when it had a flexible managed floating rate in 2005.

Two main theories (or approaches) explain the relationship between share price and exchange rate: 1) the goods market theory (also called either the “flow-oriented model” or the “traditional approach”) and 2) the portfolio balance theory (also called the “stock-oriented model”). The goods market theory predicts that the appreciation of a local currency should hurt its exporters; and therefore, the shares of such companies would become less desirable and affect the share market in an export-orientated country. This presumes that causality runs from the exchange rate to the share market in a negative correlation. Conversely, the portfolio balance theory asserts that causality runs from the share market to the exchange rate. However, existing empirical studies report mixed and contrasting findings. If the goods market theory/model prevails, we would expect the recent appreciation of the Chinese yuan to have caused its share market index to fall, assuming China is an export-orientated

country. According to the portfolio approach perspective, since the Chinese exchange rate has been under a managed float system (instead of a flexible system) since 2005, we would expect that stock prices would not affect the exchange rate as much as they would in a flexible exchange rate regime.

Several explanations exist to explain why such a study on China is important. The Chinese case is unique. First, the recent opening of the Chinese stock markets to foreign investors and the liberalization of its currency provides an opportunity to examine the impact of financial liberalization on the relationship between the foreign exchange rate and stock prices. Observations from the four-to-five years after the new foreign exchange regime provide enough data to examine the effects of the local currency on stock prices. Second, no other research exists which examines the impact of appreciation of local currency on stock prices. This study is the first to examine the impact of currency appreciation on stock prices using Chinese dataset. Third, since the opening up of Chinese stock markets, foreign investment in local Chinese markets has rapidly increased over the last few years, thus our results should be of interest to foreign portfolio investors concerned with their currency exposure in China.

Last, apart from the problems presented by their inconsistent results, the extant empirical studies are also riddled with methodological weaknesses (such as omitted variables bias and possible spurious regression results), because they do not ascertain cointegration among the set of variables analyzed. The present research applies the ARDL approach to cointegration and error correction models (ECM) to determine whether any evidence of causality exists between the exchange rate and stock prices

in the long and short run.² In the application of ARDL cointegration, we examine the impact of financial liberalization on the relationships between four main Chinese bilateral exchange rates and their share market performance. We discovered that a long-term equilibrium relationship (measured by cointegration) emerged between the exchange rate of the renminbi (RMB) against the US dollar, as well as the exchange rate against the Hong Kong dollar and the Shanghai A Share Index since the financial liberalization in 2005. We found that the exchange rate actually influences stock prices — and shows a positive correlation, as suggested by the goods market theory for an import-oriented economy. We also found that the money supply strongly positively influences stock prices. This paper establishes that the money supply increase is directly related to the recent external short-term capital inflow (hot money), which in turn is the consequence of the simultaneous appreciation of the RMB. Our research supports the argument that the appreciation of RMB causes the increase in hot money inflow, which then affects the money supply, which in turn affects the stock price.

² The ARDL bounds testing approach has the advantage that the critical values produced by Pesaran et al. (2001) allow for the inclusion of a mix of $I(0)$ and $I(1)$ variables in the cointegrating relationship. Although the Johansen ML approach can also be used with a mixture of $I(0)$ and $I(1)$ variables, Rahbek and Mosconi (1999) suggest that including $I(0)$ series in a VECM can produce nuisance parameters in the asymptotic distribution of the trace for the cointegration rank. See Wickens (1996) for further concerns over the Johansen ML procedure, and Tawadros (2001) for a discussion of the difficulties in interpreting the results from this technique for the monetary model.

The rest of this paper is organized as follows. The next section presents a brief literature review and the main findings in developed and emerging markets. Section 3 presents the data and methodology employed. Section 4 shows the empirical evidence for the interdependencies between stock prices and exchange rates in China. Section 5 concludes the paper.

1 Literature Review

The goods market approach assumes that the exchange rate is determined largely by a country's current account or trade balance performance. This approach posits that changes in exchange rates affect international competitiveness and trade balance, thereby influencing real economic variables such as real income and output (Dornbusch and Fisher, 1980). Stock prices, usually defined as the present value of the future cash flow of companies, should adjust to economic perspectives.³ Depending on these and other factors, an appreciation (depreciation) of the home currency may cause a net increase (net decrease) in the share market index. For example, currency appreciation is expected to stimulate the share market of an import-dominated country (a positive effect) and depress that of an export-dominated economy (a negative effect) (Obben et al., 2007).

In the empirical research, most studies focus on the relationship using only two variables, such as the foreign exchange rate and stock prices; while a few studies explore the effect of multiple real economic variables on stock market returns. Applying a multivariate arbitrage-pricing model (APM), Fama and French (1993) and Carhart (1997)⁴ explored three or four risk factors which affect stock prices with

³ Even firms that are not internationally integrated (that is, they have a low ratio of exports and imports to total sales and a low proportion of foreign currency-denominated assets and liabilities) may be indirectly affected.

⁴ Fama and French (1993) found that three stock-market factors affect stock prices in the US, including an overall market factor and factors related to firm size and book-to-market equity. Using four-factor

cross-sectional firm-level data. Other studies more relevant to the present one examine the effect of multiple macroeconomic variables (including the foreign exchange rate) on stock market returns using market-level data such as stock index data. Chen, Roll, and Ross (1986) investigated the effect of macroeconomic factors (industrial production, the money supply, inflation, the exchange rate, and long- and short-term interest rates) on the stock market returns in the US using an APM model. They found that these variables fundamentally influence either the future cash flow or the risk-adjusted discount rate in a standard stock price valuation model, in which the stock price is broadly interpreted as the present value of the expected future cash flow (Wongbangpo and Sharma, 2002).

Alternatively, the portfolio approach suggests that changes in stock prices may influence movements in exchange rates via portfolio adjustments (inflows/outflows of foreign capital). Under this model, if a persistent upward trend in stock prices occurs, inflow of foreign capital rise. However, a decrease in stock prices would induce a reduction in domestic investor wealth, leading to a fall in demand for money and lower interest rates, causing capital outflows that would result in currency depreciation. Therefore, under the portfolio balance approach, stock prices would influence exchange rates with a negative correlation.

(three factors in Fama and French, 1993, plus a one-year return momentum factor), Carhart (1997) found that these common factors in stock returns and investment expenses almost completely explain persistence in equity mutual funds' mean and risk-adjusted returns.

The relationship between stock prices and the exchange rate has been empirically analyzed over the past three decades. The results are somewhat mixed as to the significance and direction of influences between stock prices and exchange rates. Bahmani-Oskooee and Sohrabian (1992) found that in the short run, a causal relationship exists between US stock prices and the effective exchange rate of the US dollar. Ajayi and Mougoue (1996) found conflicting short-run and long-run causalities for advanced countries. Amihud (1994) and Bartov and Bohnar (1994) found that lagged — but not contemporaneous — changes in US dollar exchange rates explain firms' current stock returns. Nieh and Lee (2001) found a short-run significant relationship of only one day in certain G7 countries; for example, currency depreciation often dragged down stock returns in the German financial market, but it stimulated the Canadian and UK markets on the following day. However, an increase in stock prices often caused currency depreciation the next day in Italy and Japan. Focussing on emerging markets, Abdalla and Murinde (1997) found unidirectional causalities from exchange rates to stock prices in Korea, Pakistan, and India; but that stock prices Granger-cause exchange rates in the Philippines. Wu (2000) found that when the Singapore currency appreciated against the US dollar and Malaysian ringgit, and depreciated against the Japanese yen and Indonesian rupiah, this led to a long-run increase in stock prices in the 1990s; however, they found mixed results for the goods market approach. Mishra (2004) found evidence for a unidirectional causality between exchange rate returns and interest rates, and between exchange rate returns and the demand for money. Wickremasinghe (2006) found one unidirectional relationship from stock prices to the US dollar exchange rate in Sri Lanka, which supported the portfolio approach.

From an examination of the studies on the impact of the Asian financial crisis, Granger et al. (2000) report that the data from South Korea shows that exchange rates lead stock prices with a positive correlation. However, the data from the Philippines was consistent with the portfolio approach that stock prices led exchange rates with a negative correlation. Data from Hong Kong, Malaysia, Singapore, Thailand, and Taiwan indicated strong feedback relationships. Yong and Isa (2000) found that the depreciation of the Malaysian ringgit negatively affected stock market prices and resulted in a sharp fall in stock prices after the Asian crisis, indicating that the exchange rate had a positive effect on the local stock price. However, this result was inconsistent with results for the period prior to the currency crisis. By using daily stock data for Thailand and the four Asian Tigers, Fang (2002) found that currency depreciation adversely affected stock returns and/or increased market volatility over the period of the Asian crisis during 1997–1999. In short, during the financial crisis, the exchange rate had a positive effect on local stock prices in these affected Asian markets, which is inconsistent with what the goods market approach predicts.

By examining the impact of stock market liberalization on the foreign exchange rate and stock markets, Phylaktis and Ravazzollo (2005) found that stock and foreign exchange markets are positively related, and that the US stock market acts as a conduit for these links. Tabak (2006) found that a linear Granger causality exists between stock prices and exchange rates, with a negative correlation in Brazil after its abandonment of the crawling peg exchange rate regime in 1999, and this supports the portfolio approach. Murinde and Poshakwale (2004) found that for the pre-euro period, stock prices unidirectionally Granger-caused exchange rates to shift only in Hungary, while bidirectional causality relations existed in Poland and the Czech Republic. After the

adoption of the euro, exchange rates unidirectionally Granger-caused stock prices to shift in all these countries. Horobet and Ilie (2007) also found a lack of causality before 2004, when the exchange rate was controlled by the Romanian National Bank; while finding that several causality relationships existed after 2004 following the adoption of a more flexible exchange rate. The Granger test results from Horobet and Ilie (2007) indicate that exchange rates are the leading variables for stock prices, and that the stock market adjusts quite dramatically to changes in exchange rates after one month.

Recent empirical studies into the effect of multiple macroeconomic factors on stock prices provide a more robust technique for further examination on the relationship between stock prices and the foreign exchange rate. Numerous studies, including Chen et al. (1986) and Mukherjee and Naka (1995), modeled this relationship in developed countries such as the US and Japan. In their studies, the signs of the long-run elasticities of macroeconomic variables are generally consistent with their a priori hypotheses. For example, Mukherjee and Naka (1995) found that the relationship between the Tokyo Stock Exchange (TSE) and the exchange rate was negative (that is, the TSE increases as the Japanese yen depreciates against the US dollar). This result is consistent with the goods market theory. However, the economic effect of stock markets on emerging markets is less clear.

Researchers have attempted to determine how and whether these emerging markets respond to changes in their fundamental economic variables, including the foreign exchange rate, compared with well-developed markets. Kwon and Shin (1999) explored the relationship between the Korean stock market and basic economic factors using a cointegration and causality analysis. They found a positive relationship between stock prices and the value of a currency, which is inconsistent with goods

market theory. While these results contrast to those found for the US and Japan, their results demonstrate that inflation- and interest rate-related variables are not significant factors for the Korean stock market. Kwon and Shin (1999) also found that the Korean market is more sensitive to foreign exchange rates, the trade balance, the money supply, and the production index.

Based on a study on the ASEAN-5 countries (Indonesia, Malaysia, the Philippines, Singapore, and Thailand), Wongbangpo and Sharma (2002) found that in the long run, the exchange rate variable was negatively related to stock prices in Indonesia, Malaysia and Philippines, but positively related in Singapore and Thailand. They argue that the competition in the world exporting market explains the negative stock price–exchange rate relationship, which is consistent with the goods market theory; yet the positive relationship could be justified via an asset view of the exchange rate. Using only two variable cointegration analyses between stock prices and the foreign exchange rate, they also found that a long-run relationship existed only in the Philippines, which, they argue, was facilitated by the adopted independent floating exchange rate policy there. Moreover, they found that all five national stock indices were positively related to growth in output, and negatively to the aggregate price level, which is consistent with the finding in developed countries. However, the relationship between the stock price index and interest rates were also found to be positive for Indonesia and Malaysia.

In summary, no theoretical consensus exists on either the existence of a relationship between stock prices and the exchange rate or its direction. The goods market approach indicates that currency appreciation is expected to show a positive

correlation between the exchange rate and stock prices in an import-dominated economy; while a negative correlation is expected for an export-dominated economy. Because the definitions of export- or import-dominated countries have been ambiguous for some time, the net effect of the aggregation (stock index) cannot be determined, and therefore the sign is arbitrary. Consequently, the empirical results based on these models are mixed, and contradict each other.

Some results in the extant research found that exchange rates Granger-cause stock prices for some export-dominated East Asian countries with a positive correlation, which is the opposite to what was expected from the goods market approach. Theoretical economists and empirical researchers are far from reaching a consensus regarding the interaction between stock market and foreign exchange markets, and therefore it is advisable that future researchers carry out further tests and analyses of this issue.

Apart from the theoretical gap in the goods market approach, the mixed and contrasting empirical results might also be attributed to the inappropriate statistical methods applied in previous studies. First, most prior research used only two variables (stock prices and the foreign exchange rate), and the regression model based on these two variables may have suffered from the “omitted variable” problem, because other variables should be included in the model.⁵ After our review of the literature on the effect of multiple macroeconomic variables on stock market return, we decided to

⁵ We are indebted to an anonymous reviewer for their constructive suggestions regarding this issue.

include these relevant macroeconomic variables, such as money supply (M1), the CPI, and the industrial production index (IPI), in our regression models for robust tests.

Second, most prior research only applied the DF test for the unit root, which is not robust for structural breaks. The only exception is the work of Granger et al. (2000), which used the Zivot and Andrews (1992) (Z-A) test for the unit root. As argued by Granger et al. (2000), the null hypothesis of a unit root in exchange rates cannot be rejected using the DF test, and is more easily rejected by using more robust tests (such as the Z-A test) for some markets. The prevalence of the lack of a unit root may be because these currencies were pegged to the US dollar before they were floated. This issue is particularly important when considering countries which were transferred from the peg system to a more floating system.

The proposed methodology for cointegration in the present research is the ARDL approach (Pesaran and Shin, 1995; Pesaran et al., 2001). More recent studies indicate that the ARDL approach to cointegration is preferable to other conventional cointegration approaches, such as that of Engle and Granger (1987) and Johansen (1991; 1995). One reason for preferring ARDL is that it is applicable irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. The statistic underlying this procedure is the familiar Wald (or F-statistic) in a generalized Dickey-Fuller type regression, which is used to test the significance of lagged levels of the variables under consideration in a conditional unrestricted equilibrium error correction model (ECM) (Pesaran et al., 2001, pp. 289–290). Another reason for using the ARDL approach is that it is more robust and performs better for small sample sizes (such as in this study) compared to other cointegration techniques.

3. Methodology and Dataset

3.1 Unit Root Test

The conventional ADF and PP unit root tests are biased towards the non-rejection of the unit root null hypothesis in the presence of structural breaks. These tests lack power in the presence of structural breaks in series, and they may fail to show whether a series is first difference stationary (Wilson et al., 2003, p. 445). Structural changes may occur in a time series for different reasons, such as during economic and political crisis, environmental crisis, institutional changes, or policy changes.

Zivot and Andrews (1992) extended Perron's (1989) test to allow for structural change at an unknown break. They considered the following null hypothesis for the series (y_t) in the model:

$$y_t = \eta + y_{t-1} + \varepsilon_t \quad (1)$$

Zivot and Andrews (1992) assert that with this null hypothesis, the dummy variable DT_B is not needed. The alternative hypothesis stipulates that y can be represented by a trend-stationary process with possible structural change occurring at an unknown point in time, as in the following three models:

$$\text{Model (A)} \quad y_t = \hat{\eta}^A + \hat{\theta}^A DU_t(\hat{T}b) + \hat{\beta}^A t + \hat{\alpha}^A y_{t-1} + \sum_{j=1}^k \hat{c}_j^A \Delta y_{t-j} + \hat{e}_t \quad (2)$$

$$\text{Model (B)} \quad y_t = \hat{\eta}^B + \hat{\beta} t + \hat{\gamma}^B DT_t(\hat{T}b) + \hat{\alpha}^B y_{t-1} + \sum_{j=1}^k \hat{c}_j^B \Delta y_{t-j} + \hat{e}_t \quad (3)$$

$$\text{Model (A)} \quad y_t = \hat{\eta}^C + \hat{\theta}^C DU_t(\hat{T}b) + \hat{\beta}^C t + \hat{\gamma}^C DU_t(\hat{T}b) + \hat{\alpha}^C y_{t-1} + \sum_{j=1}^k \hat{c}_j^C \Delta y_{t-j} + \hat{e}_t \quad (4)$$

where $DU_t(\hat{T}b) = 1$ if $t > T_B$, 0 otherwise and $DT_t(\hat{T}b) = t - T_B$ if $t > T_B$.

The k extra regressors in the preceding regressions are added to eliminate possible nuisance-parameter dependencies in the limit distributions of the test statistics caused by temporal dependence in the disturbances.

3.2 ARDL Approach

The ARDL approach involves estimating the conditional error correction version of the ARDL model for variables under estimation. The augmented ARDL

(p, q_1, q_2, \dots, q_k) is given by the following equation (Pesaran and Pesaran, 1997;

Pesaran and Shin, 2001):

$$\alpha(L, p)y_t = \alpha_0 + \sum_{i=1}^k \beta_i(L, q_i)x_{it} + \lambda' w_t + \varepsilon_i \quad \forall_t = 1, 2, \dots, n \quad (5)$$

where

$$\alpha(L, p) = 1 - \alpha_1 L - \alpha_2 L^2 - \dots - \alpha_p L^p$$

$$\beta_i(L, q_i) = \beta_{i0} + \beta_{i1}L + \beta_{i2}L^2 + \dots + \beta_{iq_i}L^{q_i} \quad \forall_t = 1, 2, \dots, k$$

Where y_t is the dependent variable, α_0 is the constant term, L is the lag operator,

such that $Ly_t = y_{t-1}$, w_t is $s \times 1$ vector of deterministic variables such as intercept

term, time trends, or exogenous variables with fixed lags. The long-run elasticities are estimated by:

$$\hat{\phi}_i = \frac{\hat{\beta}_i(1, \hat{q}_i)}{\phi(1, \hat{p})} = \frac{\hat{\beta}_{i0} + \hat{\beta}_{i1} + \dots + \hat{\beta}_{iq}}{1 - \hat{\alpha}_1 - \hat{\alpha}_2 \dots - \hat{\alpha}_{\hat{p}}} \quad \forall i = 1, 2, \dots, k \quad (6)$$

where \hat{p} and $\hat{q}_i, i = 1, 2, \dots, k$ are the selected (estimated) values of \hat{p} and $\hat{p}_i, i = 1, 2, \dots, k$.

The long-run coefficients associated with the deterministic/exogenous variable with fixed lags are estimated by

$$\pi = \frac{\hat{\lambda}(\hat{p}, \hat{q}_1, \hat{q}_2, \dots, \hat{q}_k)}{1 - \hat{\alpha}_1 - \hat{\alpha}_2 - \dots - \hat{\alpha}_{\hat{p}}} \quad (7)$$

where $\hat{\lambda}(\hat{p}, \hat{q}_1, \hat{q}_2, \dots, \hat{q}_k)$ denotes the OLS estimates of λ in equation (1) for the selected ARDL model.

The ARDL approach involves two steps for estimating the long-run relationship (Pesaran et al., 2001). The first step is to investigate the existence of a long-run relationship among all variables in the equation under estimation. The ARDL method estimates $(p + 1)^k$ number of regressions in order to obtain the optimal lag length for each variable, where p is the maximum number of lags to be used and k is the number of variables in the equation. The second step is to estimate the long-run relationship and short-run bidirectional causality between running actors. We ran the second step only if we found a long-run relationship in the first step (Narayan et al., 2005). This study uses a more general formula of the ECM with an unrestricted intercept and unrestricted trends (Pesaran et al., 2001):

$$\Delta y_t = c_0 + c_1 t + \pi_{yy} y_{t-1} + \pi_{yx.x} + \sum_{i=1}^{p-1} \psi_i' \Delta z_{t-1} + w' \Delta X_t + \mu_t \quad (8)$$

Where $c_0 \neq 0$ and $c_1 \neq 0$. The Wald test (F-statistics) for the null hypothesis is

$$H_0^{\pi_{yy}} : \pi_{yy} = 0, H_0^{\pi_{yx.x}} : \pi_{yx.x} = 0', \text{ and alternative hypothesis}$$

$$H_1^{\pi_{yy}} : \pi_{yy} \neq 0, H_1^{\pi_{yx.x}} : \pi_{yx.x} \neq 0'.$$

Hence, the joint null hypothesis of the interest in the

above equation is given by $H_0 = H_0^{\pi_{yy}} \cap H_0^{\pi_{yx.x}}$, and the alternative hypothesis is

correspondingly stated as $H_0 = H_1^{\pi_{yy}} \cap H_1^{\pi_{yx.x}}$.

The asymptotic distributions of the F-statistics are non-standard under the null hypothesis of no cointegration relationships between the examined variables, irrespective of whether the variables are purely $I(0)$ or $I(1)$, or mutually cointegrated. Pesaran and Pesaran (1997) provide two sets of asymptotic critical values. The first set assumes that all variables are $I(0)$; while the second set assumes that all variables are $I(1)$. Therefore, if we find that the computed F-statistic is greater than the upper bound critical value, then we should reject the null hypothesis of no cointegration and conclude that a steady state of equilibrium between the variables exists. If the computed F-statistic is less than the lower bound critical value, then we could not reject the null of no cointegration. If the computed F-statistic should fall within the lower and upper bound critical values, then the result would be inconclusive. The model can be selected using the lag length criteria of Schwartz-Criteria (SBC) or Hannan-Quinn (HQ) information criterion.

Following the discussion of theoretical models and the ARDL technique, we employed the Pesaran et al. (2001) procedure to investigate the existence of a long-run relationship in the form of the unrestricted error correction model for each variable, as follows:

$$\begin{aligned} \Delta \ln(SHA_t) = & \alpha_0 + \alpha_1 t + \sum_{i=1}^p \alpha_2 \Delta \ln(SHA_{t-1}) + \sum_{i=1}^p \alpha_3 \Delta \ln(REER_{t-1}) + \sum_{i=1}^p \alpha_4 \Delta \ln(M1_{t-1}) + \\ & \sum_{i=1}^p \alpha_5 \Delta \ln(CPI_{t-1}) + \sum_{i=1}^p \alpha_6 \Delta \ln(IPI_{t-1}) + \alpha_7 \ln(SHA_{t-1}) + \alpha_8 \ln(REER_{t-1}) + \alpha_9 \Delta \ln(M1_{t-1}) + \\ & \alpha_{10} \Delta \ln(CPI_{t-1}) + \alpha_{11} \Delta \ln(IPI_{t-1}) + \alpha_{12} D(Tb)_t + \alpha_{13} DU_t + \eta_i \end{aligned} \quad (9)$$

Where SHA is the Shanghai stock index, REER is the real exchange rate of the RMB against the other currency, M1 is the money supply, CPI is the consumer price index, and IPI is the industrial production index. We also replaced the money supply (M1) with the hot money proxy (HM) to run the regression. We did not include interest rate-related variables due to their unavailability. The exclusion of these interest rate-related variables should not cause any substantial impact on our results for two reasons. First, previous research on other emerging markets such as Korea (Kwon and Shin, 1999) indicates that these variables are not significant factors for these markets. Second, interest rates in China are regulated and not based on market supply and demand; therefore, it is logical to argue that the interest rates should not be significant factors for both stock prices and the foreign exchange rate in China. We repeatedly estimated the above ADRL model using each real exchange rate, that is, one of RMB against the US dollar (USD), one against the HK dollar (HKD), one against the Japanese yen (yen), and one against the euro dollar (euro).

In the above equation, t is the time trend variable, while η , μ , and ω are error terms in the models. The first part of these equations with $\alpha_2 - \alpha_6$ represents the short-run dynamics of the models; whereas the second part with $\alpha_7 - \alpha_{11}$ represents the long-run phenomenon. The null hypothesis in the equation (8) is $\alpha_7 = \dots = \alpha_{11} = 0$, which indicates the non-existence of a long-run relationship and vice versa.

The third stage includes conducting standard Granger causality tests augmented with a lagged error-correction term. The Granger representation theorem suggests that Granger causality will be present in at least one direction if a cointegration relationship exists among the variables, provided they are integrated in the order of one. An augmented form of the Granger causality test is involved with the error-correction term, and it is formulated in a bivariate p th order vector error-correction model (VECM), as follows:

$$\begin{bmatrix} \Delta \ln(SHA_t) \\ \Delta \ln(REER_t) \end{bmatrix} = \begin{bmatrix} k_1 \\ k_2 \end{bmatrix} + \sum_{i=1}^p \begin{bmatrix} d_{11}(L) & d_{12}(L) \\ d_{21}(L) & d_{22}(L) \end{bmatrix} \begin{bmatrix} \Delta \ln SHA_{t-i} \\ \Delta \ln REER_{t-i} \end{bmatrix} + \begin{bmatrix} \gamma_1 ECT_{t-1} \\ \lambda_1 ECT_{t-1} \end{bmatrix} + \begin{bmatrix} C_1 \\ C_2 \end{bmatrix} + \begin{bmatrix} \eta_1 \\ \eta_2 \end{bmatrix} \quad (10)$$

Where Δ is the difference operator, ECT represents the error-correction term derived from the long-run cointegrating relationship via the ARDL model, $C(i = 1, 2)$ is constant and η ($i = 1, 2$) are serially uncorrelated random disturbance terms with a mean of zero.

Through the ECT, the VECM provides new directions for Granger causality to appear. Long-run causality can be revealed through the significance of the lagged

ECTs by a t test, while the F-statistic (or Wald test) investigates short-run causality through the significance of a joint test with an application of the sum of lags of explanatory variables in the model.

To ascertain the goodness of fit of the ARDL model, we conducted diagnostic and stability tests. The diagnostic test examines the serial correlation, functional form, normality and heteroskedasticity associated with the model. The stability test employs the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of the squares of recursive residuals (SUSUMsq).

We rely on monthly frequency data since weekly or daily data would preclude the use of macroeconomic data such as money and income. The data consists of monthly Shanghai A Share Index expressed in local currency. This index is the most popular index in China for representing share market performance in China. Our dataset also consists of local bilateral spot exchange rates, money supply (M1), consumer price index (CPI), industrial production index (IPI), and hot money proxy (HM) over the period December 1995 to December 2009.

The local bilateral spot exchange rates utilized in the study comprise those major currencies used in constructing China's trade weighted index (TWI) series. These include the bilateral nominal exchange rate of the RMB against the US dollar (USD), the Hong Kong dollar (HKD), the Japanese yen (yen), and euro dollar (euro). The exchange rate convention applied is the price of the domestic currency (that is, the RMB) in units of the foreign currency; hence, an increase implies appreciation of the

RMB, and a decrease implies a depreciation of the RMB.⁶ The real exchange rate (REER) is defined as:

$$\ln(\text{REER}) = \ln \text{ER} + \ln \text{CPI}^{\text{foreign}} - \ln \text{CPI}^{\text{local}}$$

where $\text{CPI}^{\text{local}}$ is the consumer price index for China, ER is the nominal exchange rate and $\text{CPI}^{\text{foreign}}$ is the consumer price index for the respective country.

The hot money proxy (HM) is defined as:

$$\text{HM} = \text{foreign reserve change} - \text{foreign trade surplus} - \text{utilized FDI}.$$

All the data series are expressed in logarithmic form, except the CPI and IPI. All data was collected on a monthly basis and was retrieved from the People's Bank of China, the National Bureau of Statistics of China, the Ministry of Commerce of China, and DataStream.

⁶ However, the Chinese official foreign exchange rate uses the indirect quotation system, in which the price of the foreign currency (for example, the US dollar) is measured in units of the local currency. In this case, the exchange rate of the RMB against the US dollar is expressed as 7 yuan/US dollar. In the text, we use the Chinese official quotation system when we discuss Chinese currency appreciation.

4. Empirical Results

4.1. Main Results

Both the augmented Dickey-Fuller (ADF) tests and Zivot-Andrew tests were carried out on time series in levels and differenced forms. ADF statistics exhibit non-stationary log Shanghai stock index at a significance level of five percent (see Table 1). Table 1 shows that all return (log difference) series are stationary for all periods — whether or not a trend is included in the testing equation. By using traditional tests, Perron (1989) demonstrates that in the presence of a structural break in the time series, many perceived non-stationary series are in fact stationary. Further work by Zivot and Andrews (1992) elaborates and overthrows the presumed exogenous break point, and develops a unit-root test with an endogenous structural break, which we regard as a more suitable test for the order of integration of series (Nieh and Yau, 2004). The results of the Zivot-Andrew tests indicate that the $\ln(\text{SHA})$ series carry unit-roots in the level and reject the null of non-stationarity in the first difference. The results for all exchange rates before 2005–2007 are $I(0)$; while they became the $I(1)$ type afterward.

< Table 1 here >

The ARDL model requires a prior knowledge or estimation of the orders of the extended ARDL. The order of the distributed lag on the dependent variable and the regressors is selected using the Schwartz Bayesian criterion (SBC). Depending on Monte Carlo evidence, Pesaran and Smith (1998) found that the SBC is preferable to the AIC. The maximum lag is set to 12 for our monthly data. The results of the F-

statistics for testing the joint level significance of the lagged level are reported in Table 2.

< Table 2 here >

For the period prior to the financial market liberalization in 2005, the result indicates that calculated F-statistics are smaller than the lower critical value, even at a 10 percent significant level, suggesting that no cointegration relationship exists between these major foreign exchange rates and the Shanghai A Share Index. However, for the period after financial market liberalization, the null hypothesis of no cointegration cannot be accepted. Where the Shanghai stock index is a dependent variable, the result indicates that the calculated F-statistics are greater than the upper critical value at the five percent significance level (for the foreign exchange rate of the RMB against the HKD and USD), and at a 10 percent significance level (for the foreign exchange of RMB against the yen and euro respectively). These results are also supported by robust tests when we control for major macroeconomic variables, including the money supply (M1), the CPI, and the industrial production index (IPI). We dropped out both the hot money proxy (HM) and IPI variables from the equation; otherwise, the F statistics would have been below the lower critical value. This result suggests that both the variables $\ln(\text{USD})$ and $\ln(\text{HKD})$ can be treated as the “long-run forcing” variable for the explanation of $\ln(\text{SHA})$ in the period after financial liberalization, even if we control for the other major macroeconomic variables. Our results indicate that the cointegration between the foreign exchange rate and stock prices emerged since 2005, which is in agreement with some prior studies on countries which experienced foreign exchange market liberalization (Murinde and Poshakwale, 2004; Horobet and Ilie, 2007).

Table 3 reports the long-run coefficient estimates. All the regression equations are based on the ARDL model selected by the SBC. Table 3 shows that the long-run coefficients for the regressor, namely, that $\ln(\text{USD})$, $\ln(\text{HKD})$, CPI, and $\ln(\text{M1})$ are all positive and highly significant at one percent levels in regression equations with $\ln(\text{USD})$ and $\ln(\text{HKD})$ as one of the regressors respectively; while the regressor $\ln(\text{yen})$ and $\ln(\text{euro})$ are both negative and insignificant in the regression equation with $\ln(\text{yen})$ and $\ln(\text{euro})$ as one of their regressors respectively. This result suggests that a long-run relationship exists between $\ln(\text{SHA})$ and $\ln(\text{USD})$, and between $\ln(\text{SHA})$ and $\ln(\text{HKD})$ after the financial liberalization in China; however, no cointegration relationship emerged between $\ln(\text{SHA})$ and $\ln(\text{yen})$, and between $\ln(\text{SHA})$ and $\ln(\text{euro})$. This result is consistent with intuitive knowledge that both US dollars and HK dollars are the most frequently used currency in China's foreign exchange transactions; therefore, they became the key driving forces for stock prices once they had been integrated with the stock market.

These results also suggest that an increase in the exchange rate of the RMB against the US dollar by one percent will have a significant long-run impact on the Shanghai stock price index by 32 percent. This compares to an increase in the exchange rate of the RMB against the Hong Kong dollar by one percent, which had a significant long-run impact on the stock price index by 38 percent. Thus, our study found a clear and positive effect of the foreign exchange rate on the stock price index.

< Table 3 here >

Table 4 presents the estimated error correction model (ECM) of the selected ARDL (1, 4) for regression equations with $\ln(\text{USD})$ and $\ln(\text{HKD})$ as one of the regressors

respectively. The results show that the coefficient of the ECM is negative, as expected, and highly significant at the one percent level for all regression equations. The ECM represents the speed of adjustment of the $d\ln(\text{SHA})$ to its long-run equilibrium following a shock. Moreover, the significance of the ECM confirms the existence of a stable long-run relationship, and points to a long-run cointegration relationship between the significant regressor and the dependent variable. The ECMs suggest that following a shock, an adjustment of 9.8 percent and 7.7 percent back to long-run equilibrium is completed after one month in $\ln(\text{USD})$ and $\ln(\text{HKD})$ regression equations respectively. In addition to their long-run cointegration relationship, this result also suggests that both the money supply and CPI over the previous month had Granger-caused the Shanghai stock index, because we cannot accept the null hypothesis of non-causality running from $\ln(\text{M1})$ to $\ln(\text{SHA})$ as well as from CPI to $\ln(\text{SHA})$.

We also included the results of the short-run diagnostic tests for the ECM model, which indicates that our results are robust. The stability test is further conducted by employing the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of the squares of recursive residuals (SUSUMsq), which are presented in Figures 1–4. These results further confirm the robustness of our results achieved in the diagnostic tests in Table 4.

< Table 4 here >

< Figure 1 here >

< Figure 2 here >

< Figure 3 here >

< Figure 4 here >

4.2 Further Discussions

Although we found that cointegration emerged between the foreign exchange rate and stock prices after the financial liberalization in 2005, the positive correlation between these two markets in China is, however, different from the negative effect of the foreign exchange rate on stock prices found in some countries (such as Japan, in Mukherjee and Naka, 1995).

The evidence of the positive correlation between foreign exchange and stock price was also found in the existing literature., for example, as in most South-East Asian countries during the financial crisis (see Granger et al., 2000; Yong and Isa, 2000; Fang, 2002), in Singapore and Thailand during the period 1986–1996 (Wongbangpo and Sharma, 2002) and in Korean market (Kwon and Shin, 1999). Wu (2000) also found a reversal from a positive correlation to a negative correlation in the longer term, in his examination of the impact of currency depreciation on stock prices in Asia. He observes that when the local currency depreciated, foreign funds withdrew from the local share market and pushed the market down, based on the expectation of the local currency's further depreciation. In the longer term, the expectation of a local currency fall will diminish; and while a low foreign exchange pushes the interest rate down, reducing the cost of capital and promoting firms' export activities, this pushes the local market price up. From here, the share price starts to rise, thereby reversing the positive correlation into a negative correlation between the exchange rate and stock prices.

The reasons behind the positive effect of currency appreciation on the Chinese stock market are twofold. First, although recent Chinese economic growth have been driven mainly by export-led economic sectors, Shanghai A Share index (a value weighted index) is dominated mainly by banks and real estate stocks while listed exported-orientated manufacturing firms take smaller role of the index capitalisation.

Therefore the positive correlation between foreign exchange and stock price we found in this paper is not inconsistent with goods market theory. Second, the exchange rate of the Chinese yuan, built on a large foreign exchange surplus from export-led economic growth, had remained fairly stable before 2005 (at approximately 8.33 to 8.22 yuan per US dollar under the peg system). Since then, its currency has appreciated in value owing to pressure from the US and other western European countries. The appreciation took place gradually, so as to minimize the impact on firms. Such a policy was considered suitable to alleviate the adverse effects on domestic manufacturers, but served as an open invitation to hot money, which flowed into China's financial market. In order to control the induced inflation related to the increase in hot money inflow, the Central Bank of China increased the money supply. Figure 5 shows a phenomenal increase of money supply, which is accompanied by a drastic increase of the hot money (proxy) inflow and outflow we estimated during the same period.

The hot money proxy is calculated as the difference between the change in the foreign currency reserve and utilized FDI and the foreign trade balance. The hot money outflow during recent years should be overestimated, because recently, the value of the foreign assets held by the central bank should be reduced markedly due to the weakening of the US dollar. Therefore, the large negative value of these hot money

proxy figures we estimated in recent years may be simply due to the reduction of the value of foreign assets, rather than due to the capital outflow. This may also explain why we cannot find the cointegration between our hot money proxy and the foreign exchange rate, or the hot money proxy and stock prices. Although the actual hot money inflow data in China is never available for public access, the consensus from media coverage from both inside China and from Western countries indicate that hot money inflow into China increased dramatically during recent year. It was estimated that the flow of hot money into the market increased from \$US7.1 billion each year for the period 2002–2004, to \$US113.3 billion each year for the period 2005–2007, and this figure reached \$US119.5 in the first five months of 2008(China Galaxy Securities Company, 2008). Consequently, the inflow of hot money as a percentage of the increased amount of total foreign currency reserve in China rose dramatically, from approximately three percent in 2002 to 44.5 percent (China Galaxy Securities Company, 2008).

Our results indicate that money supply has very strong positive influence on the stock market. Therefore, we further argue that the overflow of this hot money inevitably led to an induced huge increase in money supply in China and then resulted in an overheated stock and real estate market, and overvalued prices. In addition, the overflow of the hot money noticeably increased the demand for the Chinese yuan, which in turn exerted pressure on its currency to appreciate. The increased capital inflows strengthened the local currency, hence its stock prices, as is anticipated under a more flexible exchange rate. However, both the liberalization of the foreign exchange market and an opening upward movement of local stock markets became more prone to contagion effects once their integration in the global financial markets

intensified. This can be observed from the dramatic fall in Chinese stock markets over the first half of 2008, due to the international stock market correction triggered by the American sub-prime loan crises. This area in particular warrants further research.

< Figure 5 here >

The efficiency and market integration of asset markets are two basic issues in financial economics that have been examined using cointegration methods for some time. The empirical literature on cointegration and market integration is based on the assumption that if two markets are (economically) integrated and their respective prices are $I(1)$, then these prices must be cointegrated (see also Granger, 1986). This assumption produces an odd conclusion: that the integrated financial markets cannot be efficient markets. However, Dwyer and Wallace (1992) demonstrate that when the random walk condition is replaced with a no-arbitrage condition, a link between efficiency and cointegration disappears. Lence and Falk (2005) further clarify the relationships among market efficiency, market integration, and statistical cointegration more clearly from theoretical perspectives. They found that four possible combinations exist (concerning market efficiency and integration) to demonstrate that the concept of cointegration is unrelated to market integration or to market efficiency, and that equity prices may be either cointegrated — or not — in all four possible combinations. We found the emergence of cointegration between the foreign exchange rate and stock prices in Chinese markets after 2005 when foreign exchange rates started to rely on more on market supply and demand instead of controlled by a pegged system. Logically, the foreign exchange market has become more efficient since then; therefore, the emergence of the cointegration should not suggest these markets became less efficient than they were before 2005. We also may

be not able to argue sufficiently that Chinese foreign and stock markets are now efficient markets, because both markets are cointegrated.

Conclusions

After the Chinese financial market was liberalized (including the opening of the stock markets to foreign investors and the change of the exchange rate system from a currency pegged to the US dollar to a more flexible managed floating system), we found that Chinese stock prices became cointegrated with the foreign exchange rate of the RMB against the US dollar and HK dollar. We found that a one percent change in the exchange rate of the RMB against the US dollar causes a 32 percent change in the Shanghai stock market index in the long run; while a one percent change in the exchange rate of RMB against the Hong Kong dollar results in a 38 percent change in the stock index.

Our results basically support the goods market theory but not the portfolio balance theory. That the portfolio theory was not supported is understandable, due to the Chinese still having a managed floating exchange rate. Therefore, we did not expect that stock prices would affect the exchange rate as much as expected in a floating exchange rate regime. We found that the exchange rates (both the exchange rate of the RMB against the US dollar and Hong Kong dollar) influences stock prices with a positive correlation. We argue that the stock price of the Shanghai A Share index, a value weighted index mainly driven by banks and real estate stocks, is affected positively by the rise of local currency . Therefore, this positive correlation is not inconsistent with the goods market theory.

We also found that the money supply strongly influenced stock price returns in the Chinese market. We further documented that the money supply was largely caused by

recent hot money inflow from other countries. After local currency appreciation, hot money flowing from foreign funds into the local markets, followed by local investors' speculations on markets, pushed the market to a high level based on the expectation of the local currency's further appreciation.

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Table 1 Unit Root Tests

Prior to Liberalization 1995M12 to 2005M06					
Market	Augmented Dickey-Fuller		Zivot-Andrews		Structural break
	Levels	First differences	Levels	First differences	
ln(SHA)	-1.66	-4.74**	-3.62	-5.75**	1996M6
ln(USD)	-9.79**	-12.96**	-4.85**	-6.34**	1998M8
ln(HKD)	-2.63	-5.53**	-5.75**	-7.59**	1999M2
ln(yen)	-2.35	-3.77**	-5.51**	-8.29**	1998M8
ln(euro)	2.54	4.32**	5.87**	7.34**	1998M4

Post Liberalization 2005M07–2009M12					
Market	Augmented Dickey-Fuller		Zivot-Andrews		Structural break
	Levels	First differences	Levels	First differences	
ln(SHA)	-1.34	-4.92**	-1.49	-5.55**	2006M12
ln(USD)	2.35	-5.42**	-1.31	-4.97**	2007M10
ln(HKD)	-2.60	-4.53**	-2.75	-5.59**	2006M04
ln(yen)	-2.35	-4.87**	-2.51	-5.29**	2007M08
ln(euro)	2.54	4.67**	2.43	5.43**	2006M12

Notes:

“GLS-Dickey-Fuller” refers to the generalized least squares (GLS) demeaned Dickey-Fuller test (Elliott, Rothenberg and Stock, 1996).

The 5% critical value for all these test statistics is -2.86.

The number of lags was selected optimally using the Schwarz criterion.

The Zivot and Andrews (1992) test assumes that the series undergoes a deterministic shift in both the mean and trend.

The timing of the break is endogenous.

For our sample size, the 5% critical value of the test is -4.8.

** denotes a significance at a 5% level.

Table 2F-Statistics for Testing the Existence of a Long-Run Relationship

Prior to liberalization period			
Relationship(effect/cause)	Calculated F-statistics	Relationship(effect/cause) with macro-variables	Calculated F-statistics
F(SHA/USD)	1.40	F(SHA/USD, M1, CPI)	1.53
F(USD/SHA)	2.45	F(USD/SHA, M1, CPI)	2.56
F(SHA/HKD)	1.34	F(SHA/HKD, M1, CPI)	1.73
F(HKD/SHA)	2.05	F(HKD/SHA, M1, CPI)	2.33
F(SHA/yen)	2.52	F(SHA/yen, M1, CPI)	2.93
F(yen/SHA)	2.34	F(yen/SHA, M1, CPI)	2.45
Post liberalization period			
Relationship(effect/cause)	Calculated F-statistics	Relationship(effect/cause) with macro-variables	Calculated F-statistics
F(SHA/USD)	4.14**	F(SHA/USD, M1, CPI)	4.13**
F(USD/SHA)	1.22	F(USD/SHA, M1, CPI)	1.56
F(SHA/HKD)	4.86**	F(SHA/HKD, M1, CPI)	4.34**
F(HKD/SHA)	1.30	F(HKD/SHA, M1, CPI)	1.9
F(SHA/yen)	4.19*	F(SHA/yen, M1, CPI)	3.84*
F(yen/SHA)	0.44	F(yen/SHA, M1, CPI)	0.46
F(SHA/euro)	4.22*	F(SHA/euro, M1, CPI)	3.93*
F(euro/SHA)	1.23	F(euro/SHA, M1, CPI)	1.23

Notes:

The relevant critical value bounds are obtained from Pesaran and Shin (2001).

We use Table CI(v) Case V (with unrestricted intercept and unrestricted trend), where the critical values in the case of two regressors are between 3.182 and 4.126 at the 10% significant level and between 3.793 and 4.833 at the 5% significant level.

The critical values in base of four regressors are between 2.43 and 3.574 at the 10% significant level and between 2.85 and 4.049 at the 5% significant level.

* denotes that the F-statistic falls above the 90% upper bound.

** denote that the F-statistic falls above the 95% upper bound.

Table 3 Long-Run Coefficient Estimated Based on ARDL Model Selected Based on SBC Dependent Variable: Ln(SHA)

Regressors	ARDL(1,0,0,0)		ARDL(1,0,0,0)		ARDL(1,0,0,0)		ARDL(1,0,0,0)	
Intecept	-329.12***	(-3.80)	-310.82***	(-2.96)	-39.22***	(-2.8)	-3.22	(-0.61)
Ln(USD)	32.32***	(3.74)						
Ln(HKD)			38.64***	(2.89)				
Ln(yen)					-2.29	(-1.4)		
Ln(euro)							-11.99	(-0.48)
CPI	0.077***	(3.38)	0.09***	(2.76)	-0.218***	(-3.0)	-0.69	(-0.68)
Ln(M1)	9.68***	(4.18)	11.18***	(3.19)	6.45***	(3.7)	16.98	(0.77)
T	0.036**	(4.56)	0.049**	(4.90)	0.063**	(5.28)	0.0031*	(2.21)
DU	-0.004**	(-3.43)	-0.005**	(-3.93)	-0.029**	(2.83)	-0.078**	(2.99)

Notes:

*, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

**Table 4 Error Correction Model (ECM) Results for the Selected ARDL Selected Based on SBC
Dependent Variable: $\ln(\text{SHA})$**

Regressors	ARDL(1,0,0,0,0)		ARDL(1,0,0,0,0)	
$\ln(\text{USD})$	3.26***	(5.27)		
$\ln(\text{HKD})$			2.98**	(5.28)
$\ln(\text{M1})$	0.966***	(5.53)	0.86***	(5.36)
$d(\text{CPI})$	0.0082**	(2.21)	0.007***	(2.81)
$d\text{INPT}$	32.39***	(5.51)	23.99***	(5.3)
T	0.0032**	(4.42)	0.0013**	(3.45)
dDU	-0.0034	(-0.265)	-0.0046	(-0.311)
Ecm(-1)	-0.098***	(-3.24)	0.077***	(2.65)
Short-Run Diagnostic Tests				
Serial Correlation LM Test	0.891	(0.589)	0.766	(0.452)
ARCH Test	0.623	(0.878)	0.432	(0.756)

Notes:

*, ** and *** denote significance at the 10%, 5% and 1% levels respectively.

Figure 1 SUSUM and SUSUMsq for $\ln(\text{USD})$ and $\ln(\text{HKD})$ equations — Plot of Cumulative Sum of Recursive Residuals, RMB/USD

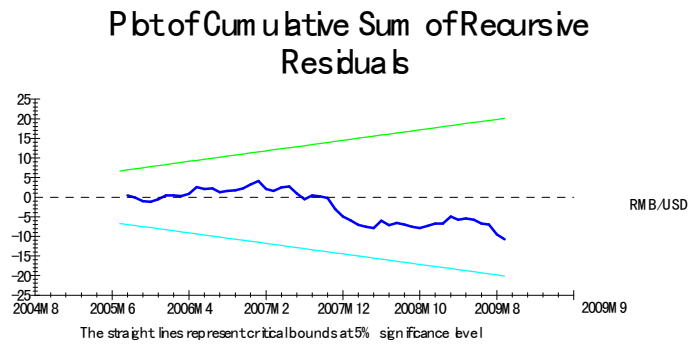


Figure 2 SUSUM and SUSUMsq for $\ln(\text{USD})$ and $\ln(\text{HKD})$ equations — Plot of Cumulative Sum of Squares of Recursive Residuals, RMB/USD

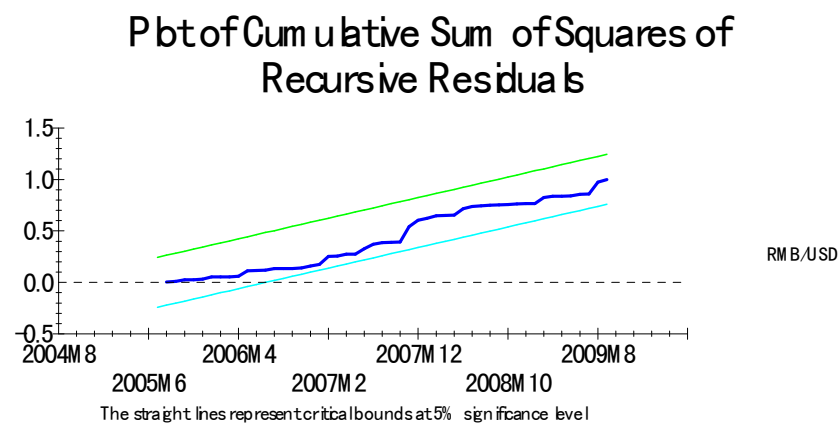


Figure 3 SUSUM and SUSUMsq for $\ln(\text{USD})$ and $\ln(\text{HKD})$ equations — Plot of Cumulative Sum of Recursive Residuals, RMB/HKD

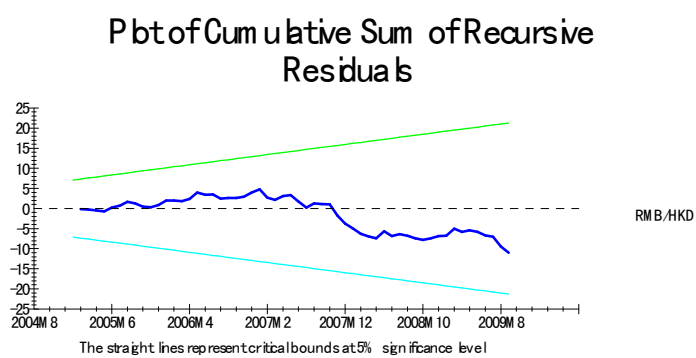
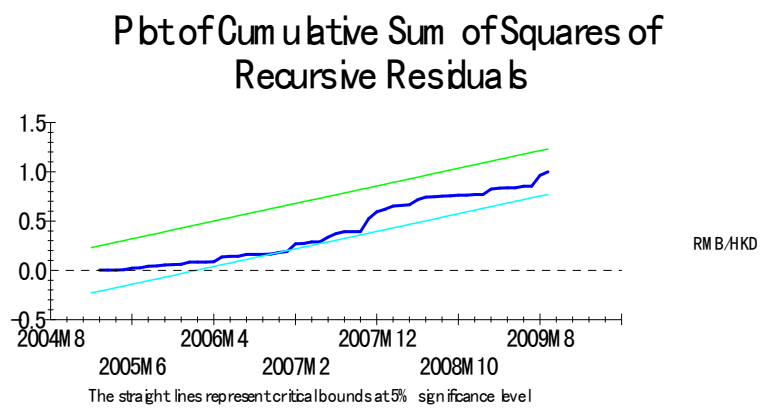


Figure 4 SUSUM and SUSUMsq for $\ln(\text{USD})$ and $\ln(\text{HKD})$ equations — Plot of Cumulative Sum of Squares of Recursive Residuals, RMB/HKD



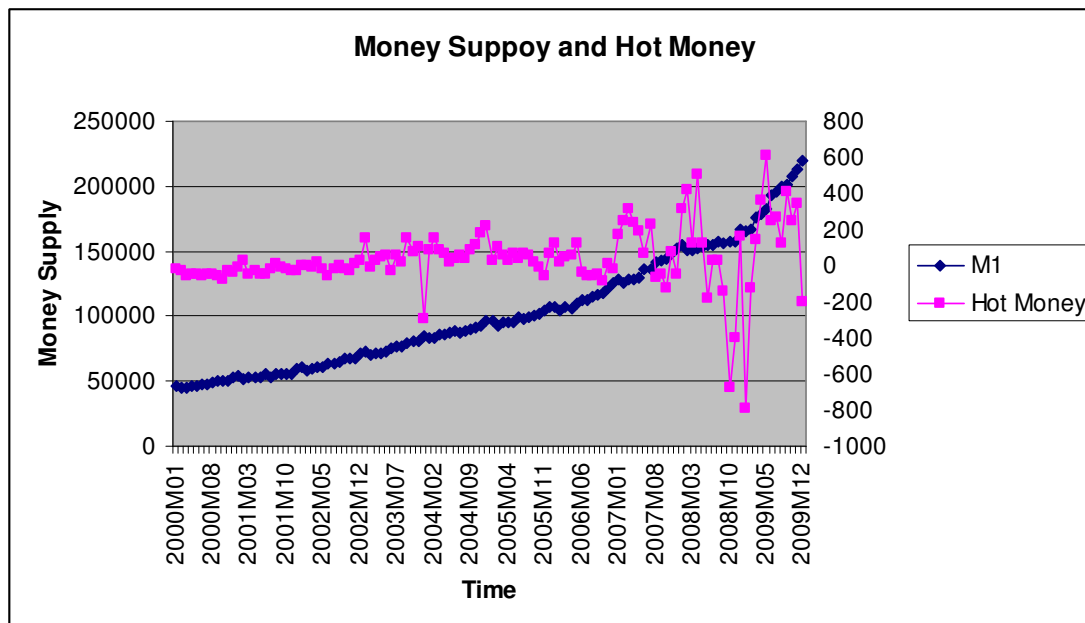


Figure 5 Monthly Hot Money Inflow (Approximation) and Money Supply