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**Working Paper**

## Global liquidity and exchange rates

Staff Report, No. 361

**Provided in Cooperation with:**  
Federal Reserve Bank of New York

*Suggested Citation:* Adrian, Tobias; Etula, Erkki; Shin, Hyun Song (2009) : Global liquidity and exchange rates, Staff Report, No. 361, Federal Reserve Bank of New York, New York, NY

This Version is available at:

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Federal Reserve Bank of New York  
Staff Reports

Risk Appetite and Exchange Rates

Tobias Adrian  
Erkko Etula  
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Staff Report no. 361  
January 2009  
*Revised December 2009*

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## **Risk Appetite and Exchange Rates**

Tobias Adrian, Erkki Etula, and Hyun Song Shin

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January 2009; revised December 2009

JEL classification: F30, F31, G12, G24

### **Abstract**

We present evidence that the funding liquidity aggregates of U.S. financial intermediaries forecast exchange rate growth—at weekly, monthly, and quarterly horizons, both in-sample and out-of-sample, and for a large set of currencies. We estimate prices of risk using a cross-sectional asset pricing approach and show that U.S. dollar funding liquidity forecasts exchange rates because of its association with time-varying risk premia. We provide a theoretical foundation for a funding liquidity channel in an intertemporal equilibrium pricing model where the “risk appetite” of dollar-funded intermediaries fluctuates with the tightness of their balance sheet constraints. Our empirical evidence shows that this channel is separate from the more familiar “carry trade” channel.

Key words: asset pricing, financial intermediaries, exchange rates

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## 1. Introduction

In market-based financial systems, the risk-bearing capacity of financial intermediaries is tightly linked to the pricing of risky assets. At the margin, all financial intermediaries borrow to fund positions in risky assets. Short-term credit aggregates such as outstanding repurchase agreements (repos) or commercial paper allow financial intermediaries to expand and contract balance sheets (see Adrian and Shin, 2007). Repos and commercial paper can thus be expected to provide a window on funding liquidity. To the extent that such credit aggregates reflect the risk appetite of financial intermediaries via the associated leverage constraints they face, we may expect a close relationship between intermediary credit aggregates and the marginal project that receives funding. Thus, we may expect financial intermediary funding conditions to convey information on market-wide risk premia.

In this paper, we uncover a link between financial intermediary funding conditions and risk premia in the foreign exchange market. We show that short-term U.S. dollar credit aggregates—the U.S. financial intermediary repos and financial commercial paper outstanding—forecast movements in the U.S. dollar cross-rates against a wide cross-section of currencies, both for developed countries as well as for emerging countries. The forecastability holds at as short as weekly forecast horizons, both in sample and out-of-sample.

Our favored explanation for the empirical findings is in terms of the risk-bearing capacity of financial intermediaries funded primarily in U.S. dollars. As the funding constraints faced by financial intermediaries loosen, their balance sheets expand and leverage rises. To an outside observer, it would be as if the preferences of the intermediaries were changing toward greater willingness to take on risk. In this way, fluctuations in intermediary credit aggregates will be associated with changes in risk appetite. When the U.S. dollar funding liquidity is high, the risk appetite of dollar-funded intermediaries is high and their required com-

pensation for holding risky assets is low. In particular, their risk premia on risky holdings of foreign currency are low, which in equilibrium implies a depreciation of such risky currencies (i.e. a dollar *appreciation* against such risky currencies). In short, we would expect expansions in dollar funding to be followed by subsequent appreciations of the dollar. This is exactly what we find in our forecasting exercises. We find further support to our risk-based explanation by estimating a cross-sectional asset pricing model, which shows that U.S. dollar credit aggregates forecast exchange rates because of their association with the price of foreign exchange risk.

It is important to distinguish our funding liquidity channel from the more familiar “carry trade” mechanism that rests on interest rate differences across currencies.<sup>1</sup> We find that the same qualitative results on the funding liquidity channel hold for US dollar cross rates against currencies as diverse as the Australian dollar, Japanese yen and the New Zealand dollar. For our sample period, the Yen is well known as a funding currency in a carry trade, while the Australian and New Zealand dollars are favored destination currencies in the carry trade. Nevertheless, expanding short-term US dollar funding forecast dollar appreciations against all three currencies. This suggests that the mechanism underlying our funding liquidity channel is distinct from the carry trade channel. In addition, controlling for interest rate differentials and for the absolute level of U.S. short-term interest rates do not change the forecasting power of the credit aggregates for the dollar cross-rates.

To the extent that our focus is on risk premia, our findings are in the broad spirit of the asset pricing approaches of Fama (1984), Hodrick (1989) and Dumas and Solnick (1995) who explain foreign exchange movements in terms of compen-

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<sup>1</sup>Empirical studies of carry trades are examined by Brunnermeier, Nagel and Pedersen (2008), Gagnon and Chaboud (2008) and Burnside, Eichenbaum, Kleshchelski and Rebelo (2007), among others. Hattori and Shin (2008) examine the role of the interoffice accounts of foreign banks in Japan for the yen carry trade.

sation for risk. Our new twist is that changes in U.S. dollar funding conditions impact the risk appetite of dollar-funded market participants, and hence market-wide risk premia. Balance sheet funding conditions in the U.S. dollar money markets impact the risk appetite of market participants in the foreign exchange market, leading to predictable fluctuations in exchange rates. A similar logic is shown to hold for commodities by Etula (2009), who shows that the risk-bearing capacity of U.S. securities brokers and dealers is a strong determinant of risk premia in commodity markets; and for options markets by Adrian and Shin (2007), who show that repo funding conditions forecast innovations to the VIX.

The pivotal role of the U.S. dollar in international capital markets gives it a special status in our investigations. However, the logic underlying our mechanism should hold more generally provided that short-term funding in a particular currency plays an important cross-border role in a particular region or sphere of influence. The increasing importance of the euro as a funding currency is a case in point. As a cross check, we conduct a supplementary empirical exercise using short-term liability aggregates denominated in euros and yen. In our panel studies, we find that just as expansions in dollar-funded balance sheets forecast dollar appreciations, expansions in euro (yen) funded balance sheets forecast appreciations in the euro (yen). However, the effects are weaker than for the U.S. dollar.

While our approach is notable in that it uses only U.S. variables to forecast the movements of the dollar against a wide cross-section of currencies, our data source also has its limitations. Chief among them is that many foreign intermediaries that use U.S. dollar funding markets are not captured in our data.<sup>2</sup> If such foreign intermediaries operate with large dollar liabilities, there may be fluctuations in dollar funding liquidity that are not fully represented in our data. The severe financial crisis and the accompanying dollar appreciation in the second half of 2008

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<sup>2</sup>Our data on repos and financial commercial paper includes only U.S. financial intermediaries plus foreign intermediaries with U.S. subsidiaries.

following the Lehman Brothers collapse had such a flavor as foreign intermediaries were widely reported as scrambling to roll over their dollar liabilities, resulting in a sharp appreciation of the US dollar. Indeed, we will see later in our paper that the crisis period of 2008-9 shows a decisive break in the empirical properties of one of our series. Modeling of the crisis period would therefore benefit from a more comprehensive database of dollar funding.

The outline of our paper is as follows. We first set the stage with our empirical analysis. We demonstrate the role of liquidity variables in explaining exchange rate movements, in both in-sample and out-of-sample forecasting exercises, for a sample of 23 currencies. We relate our results to the large literature on the forecasting of exchange rates, beginning with Meese and Rogoff's (1983) initial contribution. Our forecast exercises reveal that liquidity variables perform surprisingly well when considering the much-discussed difficulties in forecasting exchange rates out of sample. We also discuss how our results relate to the empirical literature on the carry trade, and how the funding liquidity channel explored in our paper differs from the standard carry trade logic. Having established the forecasting power of funding liquidity variables, we then focus on providing a possible rationalization for the role of dollar funding liquidity in terms of balance sheet constraints and the level of risk appetite. Based on these insights, we formulate an otherwise standard asset pricing model, but where the balance sheet constraints appear in the pricing kernel, which is modeled as being exponentially affine in a set of state variables. We go on to decompose the foreign exchange risk into systematic and idiosyncratic components to obtain prices of the risk factors. These results represent the first step in reconciling the strong empirical findings with a coherent theoretical framework.

## 2. Forecasting Exchange Rates

Despite numerous studies and a wide variety of approaches, forecasting nominal exchange rates at short horizons has remained an elusive goal. Meese and Rogoff's (1983) milestone paper finds that a random walk model of exchange rates fares no worse in forecasting exercises than macroeconomic models, and often does much better.

Evans and Lyons (2002, 2005) show that private order flow information helps forecast exchange rates, but forecasting exchange rates using public information alone has seen less success. Froot and Ramadorai (2005) show that institutional investor order flow helps explain transitory discount rate news of exchange rates, but not longer term cash flow news. Rogoff and Stavrakeva (2008) argue that even the most recent attempts that employ panel forecasting techniques and new structural models are inconclusive once their performance is evaluated over different time windows or with alternative metrics: Engel, Mark and West (2007) implement a monetary model in a panel framework to find limited forecastability at quarterly horizons for 5 out of 18 countries but their model's performance deteriorates after the 1980s. Molodtsova and Papell (2008) introduce a Taylor rule as a structural fundamental and exhibit evidence that their single equation framework outperforms driftless random walk for 10 out of 12 countries at monthly forecast horizons. However, their results are not robust to alternative test statistics, which Rogoff and Stavrakeva attribute to a severe forecast bias. Finally, Gourinchas and Rey (2007) develop a new external balance model, which takes into account capital gains and losses on the net foreign asset position. Their model forecasts changes in trade-weighted and FDI-weighted U.S. dollar exchange rate one quarter ahead and performs best over the second half of the 1990s and early 2000s.

Engel and West (2005) have provided a rationalization for the relative success of the random walk model by showing how an asset pricing approach to exchange rates leads to the predictions of the random walk model under plausible assump-



tions on the underlying stochastic processes and discount rates. In particular, when the discount factor is close to one and the fundamentals can be written as a sum of a random walk and a stationary process, the asset pricing formula puts weight on realizations of the fundamentals far in the distant future - the expectations of which are dominated by the random walk component of the sum. For plausible parameter values, they show that the random walk model is a good approximation of the outcomes implied by the theory.

In this paper, we part company with earlier approaches by focusing on U.S. dollar funding liquidity. We show that short-term liability aggregates of U.S. financial intermediaries have robust forecasting power for the bilateral movements of the U.S. dollar against a large number of currencies, both in-sample and out-of-sample. Some of our results are surprisingly strong; changes in many individual exchange rates are forecastable at as short as *weekly* horizons.

## 2.1. Data

The empirical analysis that follows uses weekly, monthly, and quarterly data on the nominal exchange rates against the US dollar of 23 countries. Our initial investigation covers the period 1993-2007. We examine the longer sample that includes the crisis period of 2008-9 in a later section. The countries include nine advanced countries (Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, UK) and fourteen emerging countries (Chile, Colombia, Czech Republic, Hungary, India, Indonesia, Korea, Philippines, Poland, Singapore, South Africa, Taiwan, Thailand, Turkey). We have excluded countries with fixed or highly controlled exchange rate regimes over most of the sample period. The exchange rate data is provided by Global Financial Data.

Our main forecasting variables are constructed from the outstanding stocks of U.S. dollar financial commercial paper and repurchase agreements of the Federal Reserve's primary dealers. The primary dealers are a group of designated banks

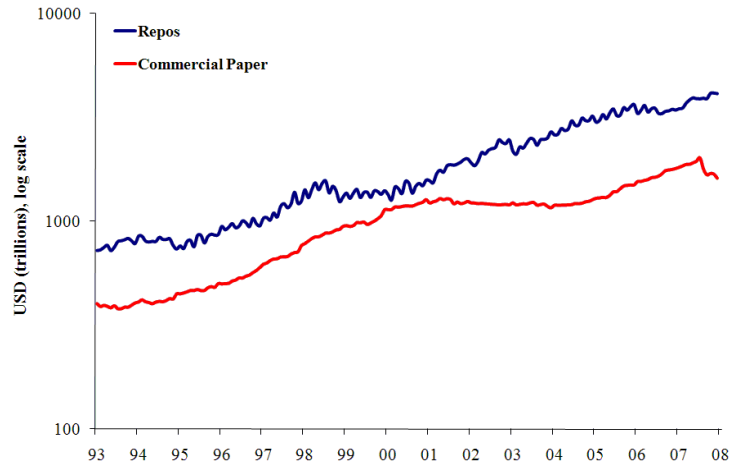


Figure 2.1: Primary dealer repos and financial commercial paper outstanding, 1/1993-12/2007

who have a daily trading relationship with the Federal Reserve Bank of New York, and which are required to report data on a weekly basis as a condition of their designation. This allows us to consider one-period-ahead forecastability at as short as weekly horizons. A plot of the logs of repos and commercial paper issuance is provided in Figure 2.1, which shows that even though both variables have exhibited strong growth over the sample period, they have hardly moved in lockstep. The apparent substitution between repos and commercial paper is better illustrated in Figure 2.2, which plots the detrended series of the logs of these variables. The detrending (with respect to a linear time trend) is performed *out of sample* in order to avoid look-ahead bias. The monthly correlation between the annual growth rates of repos and commercial paper is  $-0.43$  over 1/1993 - 12/2007.

In supplementary regressions, we also use data on the stocks of aggregate repos from Europe and Japan. The euro-denominated repos are obtained from Eurostat,

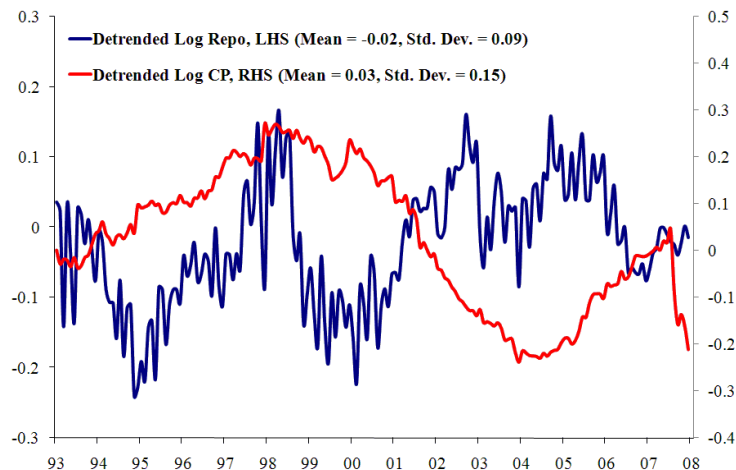


Figure 2.2: Out-of-sample detrended series of US primary dealer repos and financial commercial paper outstanding, 1/1993-12/2007

which reports the series monthly since September 1997. The yen-denominated repos are from the Bank of Japan and are reported monthly since April 2000. We were unable to find a reliable time-series for the outstanding stocks of euro or yen financial commercial paper.

In cross-sectional pricing exercises and robustness checks, we also employ country-level data on short-term interest rates and aggregate equity returns. The interest rates are 30-day money market rates, which are often most accessible to foreign investors. The equity data correspond to the returns on the country's main stock-market index. These variables are obtained from the Economist Intelligence Unit country database.

## 2.2. In-Sample Forecasting Regressions

We begin by considering a panel regression of monthly nominal exchange rate growth on lagged forecasting variables and country fixed effects. The nominal

exchange rates are defined as the units of foreign currency that can be purchased with the U.S. dollar. Hence, an *increase* in a country's exchange rate corresponds to an appreciation of the dollar against that currency. We will focus on two forecasting variables, the detrended series of U.S. dollar repos and financial commercial paper outstanding. The time period under consideration is 1/1993-12/2007. We also include controlling variables, such as the level of U.S. short-term interest rate and the interest rate differential between a particular currency and the U.S. dollar.

The regression results are displayed in Tables 1A (for whole sample of countries) and 1B (for the advanced countries only). We also provide the results at a weekly and quarterly frequency in Table 1C. We see that the credit aggregates have explanatory power for future exchange rate growth. High U.S. dollar liquidity this month tends to be followed by U.S. dollar appreciation next month. The baseline monthly panel specification is displayed in columns (i)-(ii) of Tables 1A-1B, which demonstrate that both lagged liquidity variables are highly significant forecasters of monthly exchange rate growth at 1% level. Columns (iii)-(xi) show that both the statistical significance and the magnitude of the regressions coefficients of repo growth and commercial paper growth are preserved as one includes lags of common controls, including the VIX implied volatility index, interest rate differential, and the stock market return differential. For the group of advanced countries, the TED spread seems to convey liquidity information that is similar to that incorporated by the outstanding financial commercial paper. Economically, a one standard deviation increase in detrended repos forecasts a roughly 0.2% increase in the rate of U.S. dollar appreciation; similarly, a one standard deviation increase in detrended commercial paper forecasts a 0.5% increase in the rate of dollar appreciation over the following month.

While the monthly time-series explanatory power of our panel regressions is rather modest, we emphasize that the power of our regressors stems from their

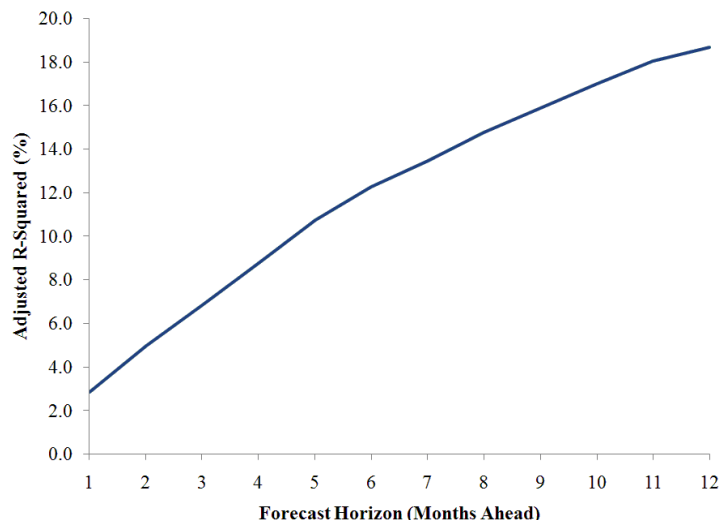


Figure 2.3: Forecasting exchange rate growth several months ahead. Time-series explanatory power in the panel of 9 developed countries, 1/1993-12/2007.

ability to predict equilibrium returns and it increases at longer forecast horizons. This result is illustrated in Figure 2.3, which plots the time-series adjusted R-squared for month-ahead to year-ahead forecasts horizons within the sample of developed countries. We see that the time-series explanatory power of the regression increases from 3% to 7% for quarter-ahead forecasts and to 12% for two-quarter-ahead forecasts. At one-year forecast horizons the balance-sheet variables are able to forecast nearly 19% of the time-series variation in exchange rate growth.

The panel regressions reveal the role of the usual carry trade channel in influencing exchange rates. In both Table 1A and Table 1B, we see that a higher U.S. short-term interest rate forecasts a future *appreciation* of the U.S. dollar. The interest rate differential is defined as the difference between the foreign (non-U.S.) short-term interest rate against the U.S. short-term interest rate. For the sample of all countries (Table 1A), the U.S. dollar tends to appreciate when the interest

differential is high (i.e. when U.S. dollar interest rate is low relative to the foreign interest rate). This result is at variance with the usual carry trade mechanism that rests on high interest rate differentials. Instead, it is consistent with dollar funding liquidity being a window on risk premia on dollar-funded risky positions across the world.

However, when the sample is restricted to the set of 9 advanced countries, the sign on the interest differential term turns negative, and highly significant. The negative sign is consistent with the carry trade channel of exchange rate movements. Thus, for the group of advanced countries, the carry trade channel appears to be a strong factor in determining exchange rate movements, independently of the risk appetite channel. We regard the negative coefficient on the interest rate differential term for the sample of 9 advanced countries as being more credible, due to greater scope of market prices to adjust to the external environment for these countries in the absence of explicit policies to peg the exchange rate, or more implicit policies of currency management.

Finally, we conduct a simple OLS regression for each country. The results are reported in Table 1D. The results indicate that at least one of the two balance sheet variables is statistically significant at 10% level for 22 out of 23 countries. In all of these cases, the significant liquidity variable enters the regression with a positive sign, implying that an increase in the U.S. dollar funding liquidity this month forecasts a U.S. dollar appreciation over the next month.

### **2.3. Contemporaneous Responses**

We motivated our forecasting regressions by arguing that the short-term liability aggregates of financial intermediaries proxy for investor risk appetite. As the funding constraints faced by financial intermediaries loosen, their balance sheets expand via higher leverage. To an outside observer, it would be as if the preferences of market participants were changing toward greater willingness to take on

risk. In this way, fluctuations in intermediary credit aggregates will be associated with changes in investor risk appetite. When short-term dollar credit is plentiful, the risk appetite of dollar-funded investors is high and their required compensation for holding risky assets is low. In particular, their risk premia on risky holdings of foreign currency are low, which in equilibrium implies a depreciation of such risky currencies (i.e. a dollar *appreciation* against such risky currencies). In short, we would expect expansions in short-term dollar credit to be followed by subsequent dollar appreciations. This is what we observed in Tables 1A-1C.

The proposed liquidity channel also has implications for the contemporaneous relationship between credit aggregates and exchange rates. An increase in dollar liquidity accompanied by an increase in risk appetite should drive up risky asset prices today.<sup>3</sup> Thus, the contemporaneous relationship between the credit aggregates and exchange rate growth should be the opposite of the lagged relationship.

To investigate the contemporaneous responses in exchange rates, we first construct series of fitted innovations for repos and commercial paper, conditioning on both variables. We then run a panel regression of exchange rate growth on lagged repo and commercial paper plus their contemporaneous innovations. These regressions are displayed in Table 2. Column (ii) shows that the contemporaneous innovations are statistically insignificant for the sample that includes all countries. Column (iv) runs the same regression for the group of developed countries. Now, the contemporaneous innovation in repos is negative and significant while the lagged balance sheet variables remain positive and significant. This finding lends some support to the contemporaneous *negative* relationship between innovations to U.S. intermediary risk appetite and the dollar.

Although the evidence on contemporaneous exchange rate responses is consistent with our intuition, we also recognize the limitations of any study of contemporaneous returns when the data frequency is so low. The instantaneous reactions

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<sup>3</sup>The intuition originates in Campbell and Shiller (1988).

in the foreign exchange market may not be captured by our low frequency data — some large movements being intra-day, for instance. Nevertheless, we offer the evidence on contemporaneous movements as further corroboration of our hypothesis.

#### 2.4. Out-of-Sample Forecasting Regressions

As is well known, the high in-sample forecasting power of a regressor does not guarantee robust out-of-sample performance, which is more sensitive to mis-specification problems. To show the extent to which the above in-sample results survive this tougher test, we turn to investigate the forecastability of exchange rate changes out of sample.

The out-of-sample performance of the monthly forecast regressions is displayed in Table 3. In order to exploit both time and cross-sectional variation in the data, the coefficient estimates for each country are generated using the fixed-effect panel specification of Table 1A. The recursive regression uses the first 4 years (1/1993-12/1996) of the sample as a training period and begins the out-of-sample estimation of betas in 1/1997.

We compare the predictive power of the our liquidity model against two benchmarks (restricted models) that are standard in the literature of out-of-sample forecasting: (1) random walk and (2) first-order autoregression.<sup>4</sup> These benchmarks are nested in the “unrestricted” specifications, which allows one to evaluate their performance using the Clark-West (2006) adjusted difference in mean squared errors:  $MSE_r - (MSE_u - adj.)$ . The Clark-West test accounts for the small-sample forecast bias ( $adj.$ ), which works in favor of the simpler restricted models and is present in the (unadjusted) Diebold-Mariano/West tests. As Rogoff and Stavrakeva (2008) show, a significant Clark-West adjusted statistic implies that there exists an optimal combination between the unrestricted model and the

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<sup>4</sup>The results are also robust to tests against other common benchmarks such as random walk with a drift.



restricted model, which will produce a combined forecast that outperforms the restricted model in terms of mean squared forecast error; i.e. the forecast will have a Diebold-Mariano/West statistic that is significantly greater than zero. The results in Table 3 indicate that the liquidity model outperforms both benchmarks at 10% significance level for 14 out of 23 countries.

Among the sample of advanced countries, we obtain out-of-sample forecastability for the exchange rates of Australia, Canada, Japan, New Zealand and Sweden. This list is notable for the fact that it includes both the typical funding currency for the carry trade (the Japanese yen) as well as two high-yielding destination currencies (Australian and New Zealand dollars). The fact that our liquidity variables enter with the same sign in all three cases suggests that the forecasting power of the liquidity variables derive from a source different from the more familiar carry trade incentives. Among the emerging countries, the liquidity variables help forecast the exchange rates of Chile, Colombia, Czech Republic, Hungary, India, Poland, South Africa, Taiwan and Turkey.

## **2.5. Supplementary Evidence from Foreign Funding Markets**

To complement our main empirical analysis, which employs only U.S. dollar liability aggregates, we also investigate the extent of exchange rate forecastability using similar variables from other funding markets. That is, if increases dollar funding liquidity forecast dollar appreciations, then one would expect increases in (say) euro funding liquidity to forecast euro appreciations.

Table 4 displays the results from simple fixed-effects panel regressions using short-term credit aggregates from the euro and yen repo markets and the exchange rates of 9 developed countries (same as above). Due to the short time-series available, we use the annual growth rates of repos instead of attempting to detrend the series out-of-sample. The first column shows that an increase in euro-denominated repos forecasts an appreciation of the euro against a panel of euro-based bilateral

exchange rates. Similarly, the second column demonstrates that an increase in yen-denominated repos forecasts an appreciation of the yen against a panel of yen-based bilateral exchange rates. Taken together, these results lend additional support to our risk-based explanation for the link between exchange rates and short-term credit aggregates.

## **2.6. Events of 2008-09**

Before we leave our empirical results section, it would be important to qualify our results in the light of the significant deterioration in financial market liquidity in the global financial crisis of 2008-09. The baseline regressions were based on data up to the end of 2007 to emphasize that our results are not driven by a few large events of the recent crisis period.

The conjunction of sharp U.S. dollar appreciation and contracting U.S. credit aggregates, which followed the bankruptcy of Lehman Brothers in the second half of 2008 could be attributed in part to contemporaneous shifts in risk appetite due to a series of shocks from the unfolding crisis, as explored above. But we find it more plausible to appeal to the fact that non-U.S. financial intermediaries (especially in emerging Europe, Latin America and Asia) were funding their operations with short-term U.S. dollar obligations. The second half of 2008 was associated with sharp depreciations of such emerging market currencies as their financial intermediaries scrambled to roll over their dollar funding.

We examine the statistical significance of our U.S.-based forecasting variables in Figure 2.4. We implement the panel regression specification of Table 1B, column (ii), recursively for 1/1993-11/2009 and plot the t-statistics of lagged repo and lagged financial commercial paper from these regressions. The figure confirms our result that both repo and commercial paper are highly significant forecasters of U.S. dollar exchange rate growth over the baseline period. However, following the Lehman bankruptcy, the statistical significance of lagged repos deteriorates

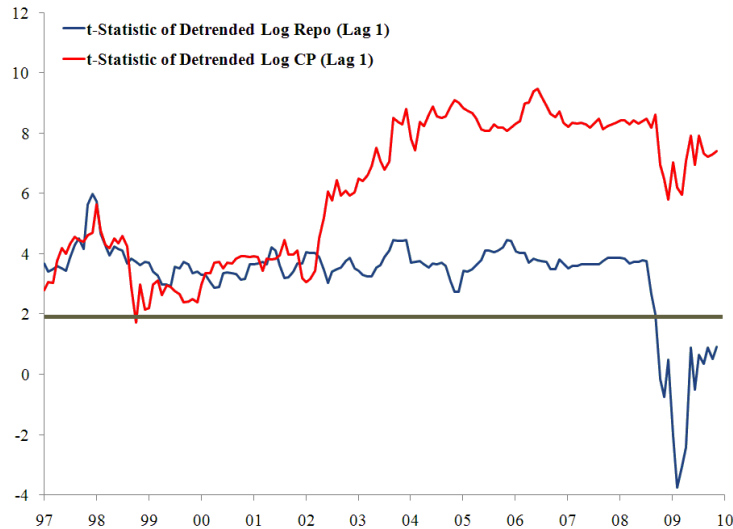


Figure 2.4: Statistical significance of lagged U.S. credit aggregates as predictors of the U.S. dollar exchange rate growth. The t-statistics are obtained from recursive fixed-effects panel regressions of exchange rate growth on lagged repo, lagged commercial paper and lagged exchange rate growth over 1/1993-11/2009 (see column (ii) of Table 1B). The critical value 1.96 corresponds to significance at 5% level.

substantially. Lagged commercial paper, on the other hand, remains a highly significant predictor of dollar exchange rate growth throughout the crisis.

Taken together, the lesson of the post-Lehman liquidity crisis is that the movements of a major funding currency such as the U.S. dollar during an acute crisis stage may not be easily captured by U.S. financial variables alone. Thus, we urge caution in interpreting our results when drawing lessons for the ongoing financial crisis.

### **3. Toward a Theoretical Framework**

Having established our benchmark empirical findings, we now turn our attention to how these results can be given firmer theoretical foundations. It is illuminating to begin by taking the cue from our empirical results, which show that the forecasting power of our funding liquidity variables is separate from the usual “carry trades” explanation for exchange rates, which emphasizes the relative attractiveness of currencies of high interest rate countries. In particular, we show that expansions in U.S. dollar funding aggregates forecast appreciations of the dollar against both high and low-yielding currencies. Thus, our approach is based on a very different rationale from the carry trades literature.

Funding liquidity conditions provide a possible explanation for why the U.S. dollar may strengthen even when the U.S. interest rate decreases. It is when short-term interest rates are low that funding conditions are favorable, and financial institutions are able to build up the size of their balance sheets through greater short-term debt (see Adrian and Shin, 2008b). Thus, more favorable funding conditions seem to increase the appetite of financial intermediaries to take on risk. To the extent that foreign currencies are regarded as risky assets by dollar-funded investors, high dollar funding liquidity should be associated with low equilibrium expected returns on these assets. That is, high dollar funding liquidity should forecast appreciations of the dollar.

In order to investigate the funding liquidity hypothesis more systematically, we now proceed to work out a simple equilibrium asset pricing framework, which illustrates how balance sheet constraints lead to fluctuations in risk appetite.

#### **3.1. Balance Sheet Constraints and Asset Prices**

Consider a leveraged financial institution such as a security broker-dealer that funds itself in the U.S. dollar and invests in foreign assets. Denote by  $Y_i$  the number of assets from country  $i$  in the dealer’s portfolio. The price of the foreign

asset in foreign currency units is  $P_i$ , and the exchange rate of the foreign currency relative to the U.S. dollar is  $\epsilon_i$ . In this section,  $\epsilon_i$  denotes the dollars that can be bought with foreign currency, and is the reciprocal of the definition of exchange rate used so far. The purpose of this switch in units is to enable us to write the balance sheet in terms of US dollars, as we will see below.

The U.S. dollar value of the foreign portfolio is thus  $\Sigma_i \epsilon_i P_i Y_i$ . Funding in the U.S. dollar market comes from two sources: capital  $w$ , and U.S. dollar debt with price  $P_{US}$  and quantity  $Y_{US}$ . The dealer's balance sheet can then be depicted as:

| Assets                        | Liabilities            |
|-------------------------------|------------------------|
| $\Sigma_i \epsilon_i P_i Y_i$ | $P_{US} Y_{US}$<br>$w$ |

with the balance sheet identity:

$$\Sigma_i \epsilon_i P_i Y_i = P_{US} Y_{US} + w. \quad (3.1)$$

For simplicity, we assume that the foreign portfolio is invested in riskless debt, and that U.S. dollar funding is riskless. But note that the analytical framework below can accommodate risky funding at the cost of some added complexity. It follows that foreign and domestic bond returns evolve according to:

$$\begin{aligned} dP_i &= P_i r_i dt \\ dP_{US} &= P_{US} r_{US} dt \end{aligned}$$

We can take the derivative of the balance sheet identity to obtain the self-financing dynamic budget constraint:

$$\begin{aligned} dw &= \Sigma_i d\epsilon_i P_i Y_i + \Sigma_i \epsilon_i r_i P_i Y_i dt - Y_{US} P_{US} r_{US} dt \\ &= \Sigma_i P_i Y_i (d\epsilon_i + \epsilon_i r_i dt) - Y_{US} P_{US} r_{US} dt \end{aligned}$$

such that, using the balance sheet identity (3.1),

$$\begin{aligned} \frac{dw}{w} &= \Sigma_i \frac{\epsilon_i P_i Y_i}{w} \left( \frac{d\epsilon_i}{\epsilon_i} + (r_i - r_{US}) dt \right) + r_{US} dt \\ &= \Sigma_i y_i dR_i + r_{US} dt, \end{aligned}$$

where

$$y_i = \frac{\epsilon_i P_i Y_i}{w} \quad \text{and} \quad dR_i = \frac{\epsilon_i P_i Y_i}{w} \left( \frac{d\epsilon_i}{\epsilon_i} + (r_i - r_{US}) dt \right). \quad (3.2)$$

Suppose that dealers are risk neutral and maximize expected portfolio returns subject to a balance sheet constraint related to their Value-at-Risk (VaR), in the manner examined in another context by Danielsson, Shin and Zigrand (2008).<sup>5</sup> The investment problem is:

$$\begin{aligned} J(t, w, x) &= \max_{\{y\}_i} E_t [e^{-\delta T} w(T)] \\ \text{subject to} &: \\ (1) &: \quad \kappa \langle dw \rangle^{\frac{1}{2}} \leq w \\ (2) &: \quad \frac{dw}{w} = \sum_i y_i dR_i + r_{US} dt \end{aligned}$$

The quadratic variation of the wealth is  $\langle dw \rangle$ . The first constraint is interpreted as a restriction on the VaR, where VaR is a constant  $\kappa$  times the forward-looking standard deviation of returns on equity. Due to risk neutrality, the VaR constraint binds with equality. We assume that returns evolve according to:

$$dR_i = \mu_i(x) dt + \sigma_i(x) dZ_i \quad (3.3)$$

$$dx = \mu_x(x) dt + \sigma_x(x) dZ_x \quad (3.4)$$

where  $\mu_i(x)$  is the conditional mean of asset returns, and  $\sigma_i(x)$  is the conditional volatility.  $Z_i$  and  $Z_x$  are Brownian Motions, with correlations  $\rho^{ij} = \langle dZ_i, dZ_j \rangle$  and  $\rho^{ix} = \langle dZ_i, dZ_x \rangle$ . Both depend on the economy's state variables. It follows that the Hamilton-Jacobi-Bellman (HJB) equation is:

$$0 = \max_{\{y\}_i} \frac{E_t [dJ]}{dt} - \Phi \left( \left\langle \frac{dw}{w} \right\rangle^{\frac{1}{2}} - \frac{1}{\kappa} \right) \quad (3.5)$$

---

<sup>5</sup> Adrian and Shin (2008a) provide a microeconomic foundation for the Value-at-Risk constraint.

where  $\Phi$  is the Lagrange multiplier on the transformed risk management constraint. We make the following guess for the value function (see Merton, 1973):

$$\begin{aligned} J(t, x, w) &= e^{f(t,x)} w \\ f(T, x) &= -\delta T, \end{aligned}$$

which implies

$$\frac{E_t[dJ]}{Jdt} = f_t + f_x \frac{E_t[dx]}{dt} + E_t \left[ \frac{dw}{wdt} \right] + \left\langle \frac{dwdx}{wdt} \right\rangle f_x + \frac{1}{2} (f_{xx} + (f_x)^2) \frac{\langle dx \rangle^2}{dt},$$

where partial derivatives are denoted by subscripts. The stacked first order conditions for portfolio choice are:

$$E_t[dR] + \langle dRdx \rangle f_x = \frac{\Phi}{J} \left\langle \frac{dw}{w} \right\rangle^{-\frac{1}{2}} \sigma \sigma' y.$$

Invoking the binding VaR constraint  $\left\langle \frac{dw}{w} \right\rangle^{\frac{1}{2}} = \frac{1}{\kappa}$  and defining  $\tilde{\Phi} = \kappa \Phi / J$ , one obtains:

$$E_t[dR] + \langle dRdx \rangle f_x = \tilde{\Phi} \sigma \sigma' y,$$

so that the portfolio choice is:

$$y = \frac{1}{\tilde{\Phi}} (\sigma \sigma')^{-1} (\mu + \sigma \sigma'_x f_x). \quad (3.6)$$

By the VaR constraint,

$$\langle dw \rangle^{\frac{1}{2}} = w \sqrt{y' (\sigma \sigma') y} = \frac{w}{\tilde{\Phi}} \sqrt{(\mu + f_x \sigma_x \sigma')' (\sigma \sigma')^{-1} (\mu + f_x \sigma_x \sigma')} = \frac{w}{\kappa},$$

which implies that the scaled Lagrange multiplier is given by:

$$\tilde{\Phi}_t = \kappa \sqrt{(\mu + f'_x \sigma_x \sigma')' (\sigma \sigma')^{-1} (\mu + f'_x \sigma_x \sigma')}.$$

From (3.6), we see that the asset demands of the intermediaries are identical to the standard ICAPM choices, but where the risk-aversion parameter  $\tilde{\Phi}_t$  is the

scaled Lagrange multiplier associated with the balance sheet constraint. Even though the intermediary is risk-neutral, it behaves as if it were risk-averse, but where the risk-aversion fluctuates with funding conditions. In other words, the intermediary’s risk appetite fluctuates with shifts in  $\tilde{\Phi}_t$ . As the balance sheet constraint binds harder, leverage must be reduced.<sup>6</sup>

Since “as if” preferences are changing with funding conditions, we would expect market prices to be affected by such changes. Our measures of aggregate short-term credit — primary dealer repos and financial commercial paper outstanding — reflect these funding conditions. Thus, our approach delivers an otherwise standard asset pricing model, but where the pricing kernel incorporates explicitly such funding liquidity effects.

Replacing the optimal portfolio choice (3.6) back into the HJB (3.5), and using the expression for the (scaled) Lagrange multiplier, one obtains:

$$0 = f_t + f_x \mu_x + r_{US} + \left( \frac{1}{\tilde{\Phi}} (\sigma \sigma')^{-1} (\mu + \sigma \sigma'_x f_x) \right) (\mu + \sigma \sigma'_x f_x) + \sigma \sigma'_x f_x + \frac{1}{2} (f_{xx} + (f_x)^2) \sigma_x \sigma'_x,$$

which simplifies to:

$$0 = f_t + f_x \mu_x + r_{US} + \tilde{\Phi} / \kappa^2 + \sigma \sigma'_x f_x + \frac{1}{2} (f_{xx} + (f_x)^2) \sigma_x \sigma'_x. \quad (3.7)$$

In the Appendix, we give a simple example where this partial differential equation can be solved explicitly.

### 3.2. Equilibrium Pricing

We assume that there is a second, passive ( $P$ ) group of investors that has constant relative risk aversion  $\Gamma$  and myopic demand. Their demand is:

$$y^P = \frac{1}{\Gamma} (\sigma \sigma')^{-1} \mu. \quad (3.8)$$

---

<sup>6</sup>Danielsson, Shin and Zigrand (2008) solve for the rational expectations equilibrium of a continuous time dynamic model along these lines.



Market clearing implies:

$$y^P w^P + yw = S, \quad (3.9)$$

where  $S$  is an aggregate supply vector of foreign positions. Plugging the two asset demands (3.6) and (3.8) in the market clearing condition gives:

$$\left( \frac{w^P}{\Gamma} + \frac{w}{\tilde{\Phi}} \right) (\sigma\sigma')^{-1} \mu + (\sigma\sigma')^{-1} \sigma\sigma'_x \frac{w}{\tilde{\Phi}} f_x = S,$$

or

$$\mu = \frac{\sigma\sigma' S}{w^P/\Gamma + w/\tilde{\Phi}} - \sigma\sigma'_x \frac{w/\tilde{\Phi}}{w^P/\Gamma + w/\tilde{\Phi}} f_x. \quad (3.10)$$

Denote the equilibrium weights of each asset by

$$s = \frac{S}{w^P + w}, \quad (3.11)$$

and the covariance matrix of individual currency returns with the portfolio by

$$\sigma\sigma'_W = (\sigma\sigma') s. \quad (3.12)$$

Furthermore, denote the wealth-weighted risk aversion by

$$\Lambda = \frac{w^P + w}{w^P/\Gamma + w/\tilde{\Phi}}, \quad (3.13)$$

and the risk-aversion weighted hedging demand terms by

$$F_x = \frac{w/\tilde{\Phi}}{w^P/\Gamma + w/\tilde{\Phi}} f_x. \quad (3.14)$$

Note that  $\Lambda$  is time-varying because  $\tilde{\Phi}$  is time-varying.

Using the notation in (3.11)–(3.14), the expected returns (3.10) can be written in the usual ICAPM form:

$$\begin{aligned} \mu &= (\sigma\sigma'_W) \Lambda - \sigma_x \sigma' F_x \\ &= Cov_t(dR, dR_W) \Lambda - Cov_t(dR, dx) F_x. \end{aligned} \quad (3.15)$$

Thus, the expected return on each asset is proportional to the scaled Lagrange multiplier. The state variables of the pricing kernel  $X$ , and prices of risk  $\lambda$ , are:

$$X = \{dR_W, dx\}' \quad (3.16)$$

$$\lambda = \{\Lambda, F_x\}' \quad (3.17)$$

So, the pricing kernel is

$$\frac{dM}{M} = -r_{US}dt - \Lambda (dR_W - E_t [dR_W]) - F_x (dx - E_t [dx]), \quad (3.18)$$

and the asset pricing implications from the model are described by:

$$\begin{aligned} dR_i &= \frac{\epsilon_i P_i Y_i}{w} \left( \frac{d\epsilon_i}{\epsilon_i} + (r_i - r_{US}) dt \right) \\ &= (Cov_t(dR_i, dR_W) \Lambda - Cov_t(dR_i, dx) F_x) dt + \sigma_i(x) dZ_i, \end{aligned} \quad (3.19)$$

where we have used (3.2), (3.3) and (3.15).

#### 4. Estimation of Foreign Exchange Risk Premia

In order to test the pricing prediction in (3.19), we write it in discrete time as:

$$\underbrace{\frac{\epsilon_{t+1}^i}{\epsilon_t^i}}_{\substack{\text{Exchange Rate} \\ \text{Appreciation}}} - \underbrace{\frac{1 + r_t^{US}}{1 + r_t^i}}_{\substack{\text{Interest Rate} \\ \text{Carry}}} = \underbrace{Cov_t \left[ X_{t+1}, \frac{\epsilon_{t+1}^i}{\epsilon_t^i} \right]}_{\substack{\text{FX Risk} \\ \text{Premium}}} \lambda_t + \underbrace{\sigma_t^i Z_{t+1}^i}_{\substack{\text{FX} \\ \text{Risk}}} \quad (4.1)$$

We will proceed by estimating the prices of risk  $\lambda_t$  that correspond to (3.19).

Since we cannot observe the prices of risk  $\lambda = \{\Lambda, F_x\}'$  directly, we make the assumption that they are affine functions of observable variables  $X$ :

$$\lambda_t = (\sigma_t^x \sigma_t^{x'})^{-1} (\lambda_0 + \lambda_1 X_t), \quad (4.2)$$

where  $\sigma_t^x \sigma_t^{x'}$  is the conditional variance-covariance matrix of  $X_{t+1}$  (see (3.4)).

Recall that  $X = \{dR_W, dx\}'$ . We proxy the return to the FX market portfolio  $dR_W$  by the first principal component of carry returns across all countries in our sample. As additional state variables  $x$ , we include the detrended log repo and the detrended log commercial paper. Thus, the vector of state variables is:

$$X_t = \begin{pmatrix} \text{FX Market Excess Return} \\ \text{Detrended Log Repo} \\ \text{Detrended Log CP} \end{pmatrix}. \quad (4.3)$$

Denoting the discrete time shocks to  $X_t$  by  $v_{t+1}$ , and using (4.2), Adrian and Moench (2008) show that equation (4.1) can be written as

$$\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i} - \frac{1 + r_t^{US}}{1 + r_t^i} = \beta^{i'} (\lambda_0 + \lambda_1 X_t + v_{t+1}) + e_{t+1}^i, \quad (4.4)$$

where  $\beta_t^{i'} = Cov_t \left[ X_{t+1}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i} \right] (\sigma_t^x \sigma_t^{x'})^{-1}$ , and  $e_{t+1}^i$  is idiosyncratic FX risk. The cross-sectional model (4.4) is estimated by way of three-step OLS regressions applied to the cross-section of 23 currencies (see Adrian and Moench (2008) for details of the estimation methodology). For simplicity, we estimate the model with constant betas for each currency  $i$ .

Table 5 displays the prices of risk for our three state variables. The first row shows that the price of FX market risk is significant and it has significant negative loadings on lagged repos and commercial paper outstanding. This result confirms our earlier intuition that funding liquidity conditions matter for the pricing of foreign exchange returns through their association with market-wide risk premia. The second and third rows indicate that any risk that stems from the innovations in repos and commercial paper can be diversified away in the cross section.

The variation in the price of FX risk over time is illustrated in Figure 4.1. The plot highlights three run-ups in market-wide risk premia that correspond to the escalation of the Enron scandal in late 2001, the Sarbanes-Oxley Act in 2002 and the subprime mortgage meltdown in late 2007.

We also investigate the significance of currency-specific factor loadings. Column (i) of Table 6 tests the joint significance of betas for each currency. The boot-

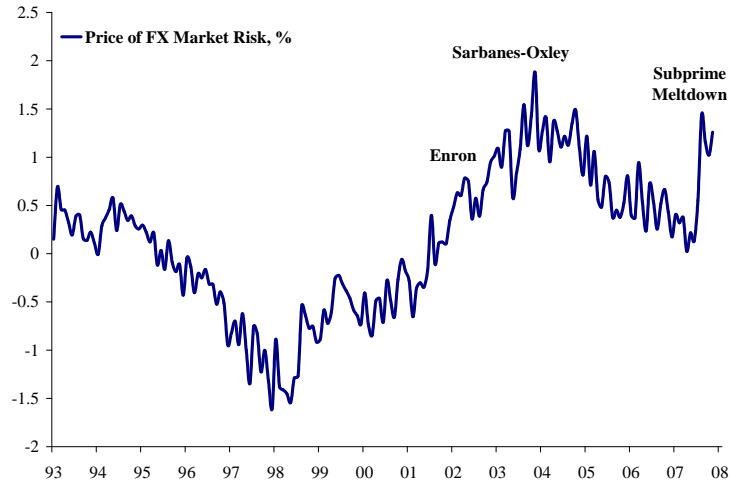


Figure 4.1: Time-variation in the price of FX risk

strapped p-values in brackets indicate that all currencies have significant loadings on the innovations of state variables. Column (ii) conducts similar tests for the FX risk premia, which correspond to the currency-specific betas multiplied by the prices of risk. The FX risk premium is significant at the 5% level for 16 out of 23 currencies.

Finally, column (iii) assesses the quality of the pricing model by testing the predictability of forecast residuals by lagged state variables. The tests of excess forecastability are significant only for New Zealand, Norway, UK, Hungary and India, which suggests that our model does a good job in pricing the rest of the cross section. That is, the observed predictability of exchange rates is largely explained by market-wide risks, which cannot be diversified away in the cross-section of currencies.

We regard the cross-sectional results as further confirmation for our favored rationalization of the channel through which the liquidity variables operate. As suggested in the sketch of our theoretical model, balance sheet constraints and

the associated Lagrange multipliers have the effect of varying the apparent risk preferences of market participants. Times of ample dollar liquidity correspond to times when constraints on dollar-funded balance sheets are relatively loose, enabling dollar-funded market participants to expand their balance sheets on the back of permissive funding conditions. In contrast, market stringency is associated with tighter balance sheet constraints and higher values of associated Lagrange multipliers. The fact that the observed predictability is explained by market-wide risks, and cannot be diversified away in the cross-section of currencies is additional evidence for liquidity variables operating through fluctuations in risk appetite.

In sum, the cross-sectional evidence supports our view that the forecastability of exchange rate growth uncovered in Tables 1-3 is in fact a reflection of systematic changes in risk premia. Higher dollar funding liquidity compresses the equilibrium returns on all risky dollar-funded positions, including those denominated in foreign currencies. This puts appreciation pressure on the dollar going forward.

## **5. Conclusion**

The random walk model has been an important benchmark in explanations of exchange rate movements. Since Meese and Rogoff's (1983) milestone paper, finding a convincing alternative to the random walk benchmark has been an elusive goal. In this paper, we have presented two related contributions that shed light on how exchange rate movements can be understood in the context of broader financial conditions.

First, building on the random walk model of exchange rates, we have demonstrated strong evidence that the short-term credit aggregates of financial intermediaries have a role in explaining future exchange rate movements. Expansions in U.S. dollar components of financial intermediary liabilities forecast appreciations of the U.S. dollar, both in sample and out of sample. The results hold over horizons as short as one week and for a wide range of cross rates. We have shown how

this result goes beyond the usual “carry trade” story, in favor of funding liquidity conditions as expressed in balance sheet fluctuations.

Second, motivated by our new empirical evidence on forecastability, we have constructed an asset pricing framework that could potentially accommodate liquidity variables in an otherwise standard asset pricing framework. Our hypothesis that funding liquidity conditions are important in the foreign exchange market is further bolstered by evidence from euro- and yen-based funding markets.

Taken together, our two contributions are first steps toward a more general framework for thinking about exchange rate movements and how the funding liquidity of investors matters for such movements. Our findings open up the possibility of understanding exchange rate movements and external adjustments in terms of the long swings associated with financial cycles and the leverage adjustments of financial intermediaries that accompany them. Much more research beckons in exploring this hypothesis further.

## Appendix

In order to obtain a closed-form solution for the PDE in (3.7), we assume that volatilities are determined as the following functions of state variables:

$$\begin{aligned}\sigma^i(x) &= \sigma^i \sqrt{x}, \\ \sigma_x &= \sqrt{x}, \\ [\sigma_x \sigma]_i &= \sigma_i \rho_i x, \\ \sigma_i \sigma_j &= \sigma_i \sigma_j \rho_{ij} x.\end{aligned}$$

We also make a somewhat restrictive assumption about the Lagrange multiplier:

$$\tilde{\Phi} = g_0 + g_1 x.$$

With these ingredients, we make the following guess for the value function:

$$f(t, x) = A(T - t) + B(T - t)x,$$

which implies:

$$\begin{aligned} f_x &= B(T - t), \\ f_t &= -A' - B'x. \end{aligned}$$

It follows that the HJB in (3.5) simplifies to:

$$A' + B'x = f'_x \gamma (\bar{x} - x) + r_{US} + \frac{g_0}{\kappa^2} + \frac{g_1}{\kappa^2} x + \sigma \rho_x x f'_x + \frac{1}{2} B^2 x,$$

with boundary conditions  $A(0) = \gamma$  and  $B(0) = 0$ . Thus, the problem can be expressed as a system of two equations:

$$A' = B\gamma\bar{x} + r_{US} + \frac{g_0}{\kappa^2}, \quad (5.1)$$

$$B' = -B\gamma + \frac{g_1}{\kappa^2} + \sigma \rho_x B + \frac{1}{2} B^2. \quad (5.2)$$

Equation (5.2) is a Ricatti differential equation, which can be solved in closed form.

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**Table 1A: Forecasting Monthly Exchange Rate Growth (All Countries)**

This table uses panel regressions with currency fixed effects to forecast exchange rate growth. The dependent variable is the monthly growth of the U.S. dollar bilateral exchange rate against 23 foreign currencies. Forecasting variables are the one-month lags of detrended log repo and detrended log financial commercial paper outstanding. Control variables (each lagged by one month) are: the interest rate differential (“carry”), the annual stock market return differential, the U.S. interest rate, the annual growth of the VIX implied volatility index and the interaction of this variable with the interest rate differential, the annual growth of the TED spread (difference between Libor and U.S. treasury bill rate) and the interaction of this variable with the interest rate differential. A lag of the dependent variable is included in (ii)-(viii). The table reports point estimates with t-statistics clustered by currency in parentheses; \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ . The sample period is 1/1993- 12/2007.

|                                    | Dependent Variable: Exchange Rate Growth (%) |                     |                      |                       |                       |                       |                       |                       |
|------------------------------------|--|---------------------|----------------------|-----------------------|-----------------------|-----------------------|-----------------------|-----------------------|
|                                    | (i)  | (ii)                | (iii)                | (iv)                  | (v)                   | (vi)                  | (vii)                 | (viii)                |
| Detrended Log Repo (Lag1)          | 1.778***<br>(3.684)                          | 1.549***<br>(3.230) | 1.512***<br>(3.510)  | 1.358***<br>(2.954)   | 1.856***<br>(3.350)   | 1.846***<br>(3.439)   | 1.927***<br>(3.550)   |                       |
| Detrended Log CP (Lag 1)           | 4.436***<br>(7.804)                          | 4.010***<br>(9.011) | 3.557***<br>(12.462) | 3.561***<br>(9.163)   | 2.719***<br>(5.690)   | 2.613***<br>(5.044)   | 1.990***<br>(3.283)   |                       |
| Exch. Rate Growth (Lag 1)          |  | 0.095***<br>(4.801) | 0.078***<br>(5.391)  | 0.062***<br>(3.590)   | 0.061***<br>(3.292)   | 0.060***<br>(3.187)   | 0.059***<br>(3.106)   | 0.060***<br>(3.161)   |
| Interest Rate Differential (Lag 1) |  |                     | 0.054***<br>(23.226) | 0.049***<br>(17.582)  | 0.048***<br>(14.160)  | 0.050***<br>(16.049)  | 0.049***<br>(14.101)  | 0.049***<br>(14.951)  |
| Stock Mkt. Ret. Dif. Ann. (Lag 1)  |  |                     |                      | -0.001<br>(-0.203)    | -0.001<br>(-0.262)    | -0.001<br>(-0.258)    | -0.001<br>(-0.248)    | -0.001<br>(-0.248)    |
| U.S. Interest Rate                 |  |                     |                      |                       | 0.105***<br>(2.616)   | 0.107***<br>(2.836)   | 0.150***<br>(3.365)   | 0.150***<br>(3.365)   |
| VIX Growth Annual (Lag 1)          |  |                     |                      |                       | -0.000<br>(-0.228)    | -0.000<br>(-0.228)    | 0.002<br>(1.008)      | 0.002<br>(1.008)      |
| Signed VIX Growth Ann. (Lag 1)     |  |                     |                      |                       | 0.003***<br>(2.838)   | 0.003***<br>(2.838)   | 0.002<br>(1.261)      | 0.002<br>(1.261)      |
| TED Growth Annual (Lag 1)          |  |                     |                      |                       | -0.002***<br>(-3.128) | -0.002***<br>(-3.128) | 0.001**<br>(2.197)    | 0.001**<br>(2.197)    |
| Signed TED Growth Ann. (Lag 1)     |  |                     |                      |                       | 0.188***<br>(9.194)   | 0.161***<br>(7.122)   | -0.052***<br>(-3.774) | -0.052***<br>(-3.774) |
| Constant                           |  |                     |                      | -0.082***<br>(-3.885) | -0.087**<br>(-2.295)  | -0.509***<br>(-3.102) | -0.516***<br>(-3.247) | -0.652***<br>(-3.670) |
| # Countries                        | 23   | 23                  | 23                   | 23                    | 23                    | 23                    | 23                    | 23                    |
| # Observations                     | 4,140  | 4,117               | 4,117                | 4,117                 | 3,761                 | 3,761                 | 3,757                 | 3,757                 |
| Adjusted $R^2$                     | 2.4%   | 3.2%                | 3.7%                 | 5.2%                  | 5.2%                  | 5.3%                  | 5.3%                  | 5.3%                  |

**Table 1B: Forecasting Monthly Exchange Rate Growth (Developed Countries)**

This table uses panel regressions with currency fixed effects to forecast exchange rate growth. The dependent variable is the monthly growth of the U.S. dollar bilateral exchange rate against 9 developed-country currencies. Forecasting variables are the one-month lags of detrended log repo and detrended log financial commercial paper outstanding. Control variables (each lagged by one month) are: the interest rate differential (“carry”), the annual stock market return differential, the U.S. interest rate, the annual growth of the VIX implied volatility index and the interaction of this variable with the interest rate differential, the annual growth of the TED spread (difference between Libor and U.S. treasury bill rate) and the interaction of this variable with the interest rate differential. A lag of the dependent variable is included in (ii)-(viii). The table reports point estimates with *t*-statistics clustered by currency in parentheses; \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ . The sample period is 1/1993–12/2007.

|                                    | Dependent Variable: Exchange Rate Growth (%) |                        |                        |                       |                       |                       |                       |                     |
|------------------------------------|--|------------------------|------------------------|-----------------------|-----------------------|-----------------------|-----------------------|---------------------|
|                                    | (i)  | (ii)                   | (iii)                  | (iv)                  | (v)                   | (vi)                  | (vii)                 | (viii)              |
| Detrended Log Repo (Lag 1)         | 2.211***<br>(3.425)                          | 2.135***<br>(3.091)    | 2.133***<br>(2.853)    | 2.113***<br>(2.544)   | 2.423***<br>(2.129)   | 2.514***<br>(2.174)   | 2.697***<br>(2.437)   |                     |
| Detrended Log CP (Lag 1)           | 3.457***<br>(8.429)                          | 3.405***<br>(8.084)    | 2.869***<br>(6.830)    | 2.269***<br>(5.295)   | 1.844***<br>(3.236)   | 1.861***<br>(3.149)   | 1.095<br>(1.171)      |                     |
| Exch. Rate Growth (Lag 1)          |  | 0.015<br>(0.941)       | 0.012<br>(1.621)       | 0.000<br>(0.742)      | 0.000<br>(0.014)      | 0.000<br>(0.031)      | -0.002<br>(-0.138)    | -0.002<br>(-0.149)  |
| Interest Rate Differential (Lag 1) |  | -0.190***<br>(-6.297)  | -0.095***<br>(-4.232)  | -0.183***<br>(-5.421) | -0.150***<br>(-3.369) | -0.162***<br>(-3.576) | -0.172***<br>(-4.017) |                     |
| Stock Mkt. Ret. Dif. Ann. (Lag 1)  |  |                        |                        | -0.006**<br>(-1.986)  | -0.007**<br>(-2.009)  | -0.007**<br>(-2.033)  | -0.007**<br>(-1.978)  |                     |
| U.S. Interest Rate                 |  |                        |                        | 0.074<br>(0.896)      | 0.074<br>(0.896)      | 0.069<br>(0.795)      | 0.122<br>(1.267)      |                     |
| VIX Growth Annual (Lag 1)          |  |                        |                        |                       |                       | -0.001<br>(-0.373)    | 0.002<br>(0.957)      |                     |
| Signed VIX Growth Ann. (Lag 1)     |  |                        |                        |                       |                       | 0.002**<br>(2.547)    | 0.000<br>(0.122)      |                     |
| TED Growth Annual (Lag 1)          |  |                        |                        |                       |                       |                       | -0.003***<br>(-2.688) |                     |
| Signed TED Growth Ann. (Lag 1)     |  |                        |                        |                       |                       |                       | 0.002***<br>(3.446)   |                     |
| Constant                           | -0.142***<br>(-11.201)                       | -0.150***<br>(-10.540) | -0.110***<br>(-74.012) | -0.136***<br>(-7.396) | -0.167***<br>(-9.091) | -0.458<br>(-1.463)    | -0.429<br>(-1.312)    | -0.587*<br>(-1.685) |
| # Countries                        | 9  | 9                      | 9                      | 9                     | 9                     | 9                     | 9                     | 9                   |
| # Observations                     | 1,620  | 1,611                  | 1,611                  | 1,611                 | 1,440                 | 1,440                 | 1,439                 | 1,439               |
| Adjusted $R^2$                     | 2.9%   | 2.9%                   | 1.5%                   | 3.1%                  | 3.6%                  | 3.6%                  | 3.5%                  | 4.0%                |

Table 1C: Forecasting Quarterly and Weekly Exchange Rate Growth

This table uses panel regressions with currency fixed effects to forecast exchange rate growth. The dependent variable is the growth of the U.S. dollar bilateral exchange rate against 23 currencies. Forecasting variables are the one-period lags of detrended log repo and detrended log financial commercial paper outstanding. A lag of the dependent variable is included as a control in columns (ii) and (iv). The table reports point estimates with t-statistics clustered by currency in parentheses; \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ . The sample period is 1993-2007.

|                           | Quarterly Exch. Rate Growth |                      | Weekly Exch. Rate Growth |                      |
|---------------------------|-----------------------------|----------------------|--------------------------|----------------------|
|                           | (i)                         | (ii)                 | (iii)                    | (iv)                 |
| Exch. Rate Growth (Lag 1) |                             | 0.142<br>(1.591)     |                          | 0.001<br>(0.034)     |
| Detrended Log Repo (Lag1) | 3.985<br>(1.611)            | 3.906*<br>(1.777)    | 0.518***<br>(4.624)      | 0.511***<br>(4.949)  |
| Detrended Log CP (Lag 1)  | 12.576***<br>(6.741)        | 10.948***<br>(9.305) | 1.051***<br>(7.502)      | 1.050***<br>(7.739)  |
| Constant                  | 0.522***<br>(6.770)         | 0.406***<br>(3.793)  | 0.047***<br>(12.077)     | 0.047***<br>(11.758) |
| # Countries               | 23                          | 23                   | 23                       | 23                   |
| # Observations            | 1,357                       | 1,334                | 20,264                   | 20,241               |
| Adjusted $R^2$            | 6.0%                        | 8.0%                 | 0.7%                     | 0.7%                 |

Table 1D: Forecasting Exchange Rate Growth Currency by Currency

This table uses OLS regressions to forecast exchange rate growth. The dependent variable is the monthly growth of the U.S. dollar bilateral exchange rate against 23 currencies (in rows). Forecasting variables (in columns) are the one-month lags of detrended log repo and detrended log financial commercial paper outstanding. The table reports point estimates with heteroskedasticity-robust t-statistics in parentheses; \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1. The sample period is 1/1993- 12/2007.

| Dep. Variable  | Independent Variables |               |            |          |          | Adj. $R^2$ |
|----------------|-----------------------|---------------|------------|----------|----------|------------|
|                | Exchange Rate Growth  | Detrended Log |            | Constant |          |            |
|                |                       | Repo (Lag 1)  | CP (Lag 1) |          |          |            |
| Australia      | 2.837                 | (1.103)       | 4.435***   | (2.909)  | -0.164   | 3.5%       |
| Canada         | 1.229                 | (0.713)       | 2.569**    | (2.516)  | -0.155   | 2.4%       |
| Germany        | 0.344                 | (0.148)       | 3.297**    | (2.384)  | -0.161   | 2.5%       |
| Japan          | 5.011*                | (1.716)       | 2.760      | (1.595)  | 0.009    | 1.0%       |
| New Zealand    | 5.752**               | (2.067)       | 6.142***   | (3.722)  | -0.212   | 6.4%       |
| Norway         | 0.706                 | (0.277)       | 3.110**    | (2.060)  | -0.152   | 1.5%       |
| Sweden         | 1.669                 | (0.620)       | 3.638**    | (2.281)  | -0.093   | 1.8%       |
| Switzerland    | 1.348                 | (0.503)       | 3.265**    | (2.056)  | -0.187   | 1.3%       |
| UK             | 1.006                 | (0.508)       | 1.895      | (1.616)  | -0.163   | 0.4%       |
| Chile          | 3.481*                | (1.661)       | 4.462***   | (3.591)  | 0.094    | 5.7%       |
| Colombia       | 1.363                 | (0.552)       | 5.602***   | (3.826)  | 0.447**  | 7.3%       |
| Czech Republic | 1.844                 | (0.635)       | 4.698***   | (2.730)  | -0.306   | 3.1%       |
| Hungary        | -2.357                | (-0.884)      | 4.133***   | (2.615)  | 0.280    | 5.1%       |
| India          | 2.174                 | (1.179)       | 2.375**    | (2.172)  | 0.233    | 1.5%       |
| Indonesia      | 4.757                 | (0.441)       | 12.776**   | (1.997)  | 1.069    | 1.2%       |
| Korea          | 5.337                 | (1.234)       | 4.501*     | (1.755)  | 0.196    | 0.8%       |
| Philippines    | 3.385                 | (1.283)       | 4.202***   | (2.686)  | 0.269    | 2.8%       |
| Poland         | 0.603                 | (0.221)       | 4.018**    | (2.486)  | 0.189    | 2.7%       |
| Singapore      | 2.420*                | (1.650)       | 2.582***   | (2.969)  | -0.080   | 3.8%       |
| South Africa   | -0.855                | (-0.223)      | 4.183*     | (1.844)  | 0.436    | 1.5%       |
| Taiwan         | 2.286                 | (1.640)       | 1.984**    | (2.400)  | 0.143    | 2.3%       |
| Thailand       | 0.293                 | (0.084)       | 3.434*     | (1.667)  | 0.122    | 0.7%       |
| Turkey         | -3.743                | (-0.636)      | 11.959***  | (3.427)  | 2.509*** | 7.8%       |

Table 2: Contemporaneous Credit Innovations and Exchange Rate Growth

This table uses panel regressions with currency fixed effects to investigate the relationship between contemporaneous innovations in short-term U.S. credit aggregates and exchange rate growth. The dependent variable is the growth of the U.S. dollar bilateral exchange rate against 23 currencies. Independent variables are the one-period lags of detrended log repo and detrended log financial commercial paper outstanding and the contemporaneous innovations of these variables computed from first-order vector auto regressions. A lag of the dependent variable is included as a control in all specifications. The table reports point estimates with t-statistics clustered by currency in parentheses; \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1. The sample period is 1/1993-12/2007.

|                            | Dependent Variable: Exchange Rate Growth (%) |                     |                        |                        |
|----------------------------|--|---------------------|------------------------|------------------------|
|                            | All Countries                                |                     | Developed Countries    |                        |
|                            | (i)  | (ii)                | (iii)                  | (iv)                   |
| Exch. Rate Growth. (Lag 1) | 0.095***<br>(4.801)                          | 0.094***<br>(4.781) | 0.015<br>(0.941)       | 0.016<br>(1.080)       |
| Detrended Log Repo (Lag1)  | 1.549***<br>(3.230)                          | 1.547***<br>(3.227) | 2.135***<br>(3.091)    | 2.134***<br>(3.092)    |
| Detrended Log CP (Lag 1)   | 4.010***<br>(9.011)                          | 4.015***<br>(8.953) | 3.405***<br>(8.084)    | 3.404***<br>(8.131)    |
| Repo Innovation            |  | 0.853<br>(0.347)    |                        | -2.429***<br>(-4.964)  |
| CP Innovation              |  | 2.903<br>(0.990)    |                        | 0.665<br>(0.286)       |
| Constant                   | 0.161***<br>(7.122)                          | 0.161***<br>(7.112) | -0.150***<br>(-10.540) | -0.150***<br>(-10.670) |
| Number of observations     | 4,117  | 4,117               | 1,611                  | 1,611                  |
| Adjusted $R^2$             | 3.2%   | 3.2%                | 2.9%                   | 3.0%                   |

Note: \*\*\* p<0.01, \*\* p<0.05, \* p<0.1; standard errors clustered by country, t-stats in parentheses

Table 3: Forecasting Exchange Rate Growth Out of Sample

This table investigates the out-of-sample forecastability of the monthly growth of U.S. dollar bilateral exchange rate relative to 23 foreign currencies. We compare the performance of our funding liquidity model against two benchmarks: (1) random walk and (2) first-order autoregression. In (1), the forecasting variables are the one-month lags of detrended log repo and detrended log financial commercial paper outstanding. In (2), we also include a lag of the dependent variable as an additional regressor. The table reports the Diebold-Mariano/West difference in mean-squared errors and the Clark-West adjusted difference in mean-squared errors. The p-values associated with the Clark-West statistic are displayed; \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ . The out-of-sample period is 1/1997-12/2007.

|                      | Random Walk Benchmark |                     |         | AR(1) Benchmark |                     |         |
|----------------------|-----------------------|---------------------|---------|-----------------|---------------------|---------|
|                      | $\Delta MSE$          | $\Delta MSE - Adj.$ | p-value | $\Delta MSE$    | $\Delta MSE - Adj.$ | p-value |
| Australia            | 0.360                 | 0.988***            | 0.010   | 0.417           | 0.910***            | 0.005   |
| Canada               | -0.094                | 0.579**             | 0.036   | -0.041          | 0.454**             | 0.050   |
| Germany              | -0.205                | 0.462*              | 0.085   | -0.108          | 0.390               | 0.101   |
| Japan                | -0.045                | 0.618*              | 0.093   | 0.071           | 0.569*              | 0.079   |
| New Zealand          | 0.476                 | 1.119***            | 0.008   | 0.595           | 1.084***            | 0.004   |
| Norway               | -0.205                | 0.465               | 0.132   | -0.093          | 0.404               | 0.148   |
| Sweden               | -0.071                | 0.570*              | 0.059   | 0.027           | 0.525*              | 0.060   |
| Switzerland          | -0.283                | 0.389               | 0.161   | -0.204          | 0.295               | 0.195   |
| UK                   | -0.349                | 0.296               | 0.164   | -0.228          | 0.275               | 0.149   |
| Chile                | 0.407                 | 0.940***            | 0.000   | 0.380           | 0.867***            | 0.001   |
| Colombia             | 0.817                 | 1.963***            | 0.000   | 0.624           | 1.104***            | 0.000   |
| Czech Republic       | 0.056                 | 0.818*              | 0.078   | 0.161           | 0.663*              | 0.088   |
| Hungary              | -0.545                | 1.205**             | 0.014   | 0.208           | 0.703**             | 0.035   |
| India                | -0.061                | 0.743***            | 0.007   | 0.040           | 0.536***            | 0.005   |
| Indonesia            | 0.271                 | 4.512               | 0.225   | 2.088           | 2.601               | 0.107   |
| Korea                | -0.041                | 0.861               | 0.288   | 0.583           | 1.096*              | 0.097   |
| Philippines          | 0.117                 | 0.856               | 0.143   | 0.155           | 0.643*              | 0.100   |
| Poland               | -0.500                | 0.934*              | 0.070   | 0.231           | 0.728*              | 0.041   |
| Singapore            | -0.265                | 0.312               | 0.203   | -0.108          | 0.389               | 0.133   |
| South Africa         | 0.447                 | 1.481*              | 0.030   | 0.342           | 0.834*              | 0.084   |
| Taiwan               | -0.097                | 0.537*              | 0.075   | -0.035          | 0.457*              | 0.068   |
| Thailand             | -0.723                | 0.122               | 0.467   | -0.126          | 0.379               | 0.307   |
| Turkey               | 1.425                 | 22.138***           | 0.000   | 1.423           | 1.888***            | 0.000   |
| # In-Sample Obs.     | 50                    | 50                  | 50      | 50              | 50                  | 50      |
| # Out-of-Sample Obs. | 130                   | 130                 | 130     | 130             | 130                 | 130     |



Table 4: Evidence from Euro and Yen Repo Markets

This table uses panel regressions with currency fixed effects to forecast exchange rate growth. The dependent variable in the first (second) column is the monthly growth of the euro (yen) bilateral exchange rate against 9 developed-country currencies. The forecasting variable is the one-month lag of the annual growth rate of euro (yen) repo outstanding. A lag of the dependent variable is included as a control. The table reports point estimates with t-statistics clustered by currency in parentheses; \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ . The sample periods are 9/1997-12/2007 (euro) and 4/2000-12/2007 (yen).

|                                  | Exch. Rate Growth   |                     |
|----------------------------------|---------------------|---------------------|
|                                  | Euro-Based          | Yen-Based           |
| Exch. Rate Growth (Lag 1)        | -0.010<br>(-0.23)   | 0.145<br>(4.68)     |
| Euro Repos (Annual Growth, Lag1) | 0.021***<br>(3.69)  |                     |
| Yen Repos (Annual Growth, Lag1)  |                     | 0.008***<br>(5.12)  |
| Constant                         | -0.001**<br>(-2.82) | 0.853***<br>(27.53) |
| # Countries                      | 9                   | 9                   |
| # Observations                   | 1179                | 972                 |
| Adjusted $R^2$                   | 1.1%                | 3.1%                |

**Table 5: Cross-Sectional Prices of Risk**

This table reports the results from the estimation of a cross-sectional arbitrage-free asset pricing model for U.S. dollar funded investments in 23 foreign currencies. The first four columns display the point estimates for the loadings of each risk factor (in rows) on a constant and the lagged state variables. The last column test the hypothesis that the price of risk of a factor is zero. Bootstrapped t-statistics based on 1000 iterations are displayed in parentheses, p-values are in brackets; \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1. The sample period is 1/1993- 12/2007.

| Residual                | $\lambda_0$         | $\lambda_1^{FX}$    | $\lambda_1^{Repo}$    | $\lambda_1^{CP}$       | $\lambda_0 = \dots = \lambda_1^{CP} = 0$ |
|-------------------------|---------------------|---------------------|-----------------------|------------------------|--|
| FX Market Excess Return | 0.209***<br>(6.180) | 0.062***<br>(3.450) | -2.273***<br>(-4.560) | -4.988***<br>(-14.240) | [0.000]***                               |
| Detrended Log Repo      | 0.020<br>(1.620)    | -0.014*<br>(-1.790) | -0.067<br>(-0.340)    | -0.039<br>(-0.290)     | [0.202]                                  |
| Detrended Log CP        | 0.007<br>(1.340)    | -0.004<br>(-1.310)  | 0.009<br>(0.150)      | 0.012<br>(0.260)       | [0.290]                                  |

Note: bootstrapped p-values in brackets, \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1

Table 6: Significance of  $\beta$ 's,  $\beta'\lambda$ 's and Excess Predictability

The first column tests the joint significance of individual currencies' loadings on the risk factors and the second column tests the joint-significance of individual currencies' loadings on the lagged state variables. The third column tests the hypothesis that the forecast residuals are not predictable by lagged state variables. Bootstrapped p-values based on 1000 iterations are reported in brackets; \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1. The sample period is 1/1993- 12/2007.

| Test Asset     | $\beta^{FX} = \dots = \beta^{CP} = 0$ | $\beta'\lambda_0 = \dots = \beta'\lambda_1^{CP} = 0$ | Predictability of<br>Forecast Residuals |
|----------------|---------------------------------------|--|---|
| Australia      | [0.000]***                            | [0.000]***   | [0.505]                                 |
| Canada         | [0.000]***                            | [0.059]**  | [0.392]                                 |
| Germany        | [0.000]***                            | [0.000]***   | [0.665]                                 |
| Japan          | [0.000]***                            | [0.000]***   | [0.128]                                 |
| New Zealand    | [0.000]***                            | [0.000]***   | [0.092]*                                |
| Norway         | [0.000]***                            | [0.000]***   | [0.010]**                               |
| Sweden         | [0.000]***                            | [0.000]***   | [0.923]                                 |
| Switzerland    | [0.000]***                            | [0.000]***   | [0.376]                                 |
| UK             | [0.000]***                            | [0.012]**  | [0.011]**                               |
| Chile          | [0.001]***                            | [0.458]  | [0.309]                                 |
| Colombia       | [0.023]**                             | [0.988]  | [0.412]                                 |
| Czech Republic | [0.000]***                            | [0.000]***   | [0.300]                                 |
| Hungary        | [0.000]***                            | [0.000]***   | [0.009]***                              |
| India          | [0.076]*                              | [0.780]  | [0.000]***                              |
| Indonesia      | [0.003]***                            | [0.565]  | [0.996]                                 |
| Korea          | [0.009]***                            | [0.508]  | [0.222]                                 |
| Philippines    | [0.001]***                            | [0.165]  | [0.641]                                 |
| Poland         | [0.000]***                            | [0.000]***   | [0.352]                                 |
| Singapore      | [0.000]***                            | [0.000]***   | [0.535]                                 |
| South Africa   | [0.000]***                            | [0.038]**  | [0.133]                                 |
| Taiwan         | [0.000]***                            | [0.009]***   | [0.226]                                 |
| Thailand       | [0.000]***                            | [0.034]**  | [0.836]                                 |
| Turkey         | [0.007]***                            | [0.243]  | [0.408]                                 |

Note: Bootstrapped p-values in brackets, \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.1