Federal Reserve Bank of New York
Staff Reports

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Staff Report no. 361
January 2009
Revised May 2010

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JEL classification: F30, F31, G12, G24

Abstract

We present evidence that the funding liquidity aggregates of U.S. financial intermediaries forecast U.S. dollar exchange rate growth—at weekly, monthly, and quarterly horizons, both in-sample and out-of-sample, and against a large set of foreign currencies. We provide a theoretical foundation for a funding liquidity channel in a simple asset pricing model where the effective risk aversion of dollar-funded intermediaries fluctuates with the tightness of their risk constraints. 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. 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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This paper was previously distributed under the title "Global Liquidity and Exchange Rates." We thank John Campbell, Jan Groen, Matti Keloharju, Lars Ljungqvist, Ken Rogoff, Andrei Shleifer, Jeremy Stein, Matti Suominen, John C. Williams, and seminar participants at Harvard University, the International Monetary Fund, the Bank of Korea, Aalto University, Georgetown University, and the Federal Reserve Bank of Dallas for comments. The views expressed in this paper are those of the authors and do not necessarily reflect the position of the Federal Reserve Bank of New York or the Federal Reserve System.

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 instruments such as repurchase agreements (repos) or commercial paper allow financial intermediaries to rapidly expand and contract balance sheets (see Adrian and Shin, 2007). Weekly reported figures of primary dealer repos and financial commercial paper outstanding can thus be expected to provide a high-frequency 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 would expect a close relationship between intermediary credit aggregates and the riskiness of 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—primary dealer 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 some emerging countries. The forecastability holds at as short as weekly forecast horizons, both in sample and out of sample.

Our 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 effective risk aversion, or "risk appetite." When the U.S. dollar funding liquidity is high, the risk appetite of dollar-funded intermediaries 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 dollar funding to be followed by subsequent appreciations of the dollar. This is exactly what we find in our forecasting exercises. We rationalize the mechanism within a simple asset pricing framework, which illustrates how fluctuations in financial intermediary risk constraints may lead to time-variation in effective risk aversion. We estimate the model in the cross-section of exchange rates and confirm that our short-term U.S. dollar credit aggregates forecast exchange rates because of their association with systematic risk premia.

It is important to distinguish our funding liquidity channel from the more familiar "carry trade" mechanism that rests on interest rate differences across currencies.[1] Specifically, we find that expansions in short-term U.S. dollar funding forecast dollar appreciations against both high and low interest rate currencies, suggesting that the mechanism underlying our funding liquidity channel is distinct from the carry trade channel.[2] 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 short-term 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

[1]Empirical studies of carry trades include Lustig, Roussanov and Verdelhan (2010), Brunnermeier, Nagel and Pedersen (2008), Gagnon and Chaboud (2008) and Burnside, Eichenbaum, Kleshchelski and Rebelo (2007), among others. Jylha and Suominen (2009) investigate the role of hedge fund capital in carry trades. Hattori and Shin (2008) examine the role of the interoffice accounts of foreign banks in Japan for the yen carry trade.

[2]For our sample period, the Yen is well known as a funding currency in the carry trade, while the Australian and New Zealand dollars are favored destination currencies in the carry trade. Nevertheless, expansions in short-term US dollar funding forecasts dollar appreciations against all three currencies.

and Solnik (1995) who explain foreign exchange movements in terms of compensation for risk. Lyons (1997) notes that since up to 80% of foreign exchange volume consists of trades between dealers, financial intermediary trading activity is expected to have a substantial impact on the information content of exchange rates. Our paper builds on these studies to empirically demonstrate that the funding conditions of financial intermediaries determine risk premia in foreign exchange markets. Measurable time-variation in foreign exchange risk premia in turn translates to predictability of future 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 (over 90% of volume in commodity derivatives is conducted over the counter); and for options markets by Adrian and Shin (2007), who show that 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 other 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.[3] 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 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 break in the empirical properties of one of our forecasting variables. 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 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 our funding liquidity variables, we then focus on providing a possible rationalization for the role of dollar funding liquidity in terms of balance sheet risk constraints and the associated level of risk appetite. Based on these insights, we formulate a simple asset pricing model where the economy's effective risk aversion varies over time with the tightness of leveraged intermediaries' risk constraints. We express the effective risk aversion as a function of aggregate balance sheet components of financial institutions and estimate the model in the

[3] Our data on repos and financial commercial paper includes only U.S. financial intermediaries plus foreign intermediaries with U.S. subsidiaries.

data. Our formulation represents the first step in reconciling the strong empirical 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 assumptions 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 nomimal exchange rates of 23 countries against the US dollar. Our initial investigation covers the period 1/1993-12/2007. 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 Datastream.

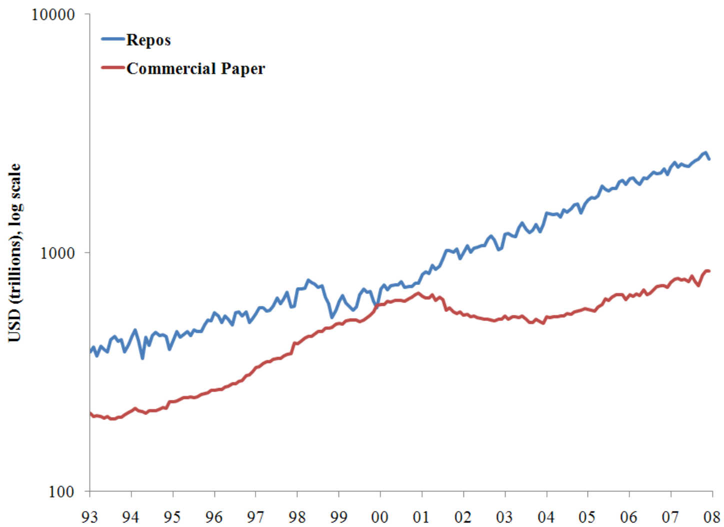

Figure 2.1: Primary dealer overnight repos and financial commercial paper outstanding, 1/1993-12/2007

Our main forecasting variables are constructed from the outstanding stocks of U.S. dollar financial commercial paper (hencefort, commercial paper) and overnight repurchase agreements of the Federal Reserve's primary dealers (henceforth, repos).[4] These data are published weekly by the Federal Reserve Board and the Federal Reserve Bank of New York, respectively. A plot of the logs of repos and commercial paper outstanding 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 detrended series of log repos and log commercial paper is

[4] The primary dealers are a group of designated banks and securities broker-dealers who have a trading relationship with the Federal Reserve Bank of New York.

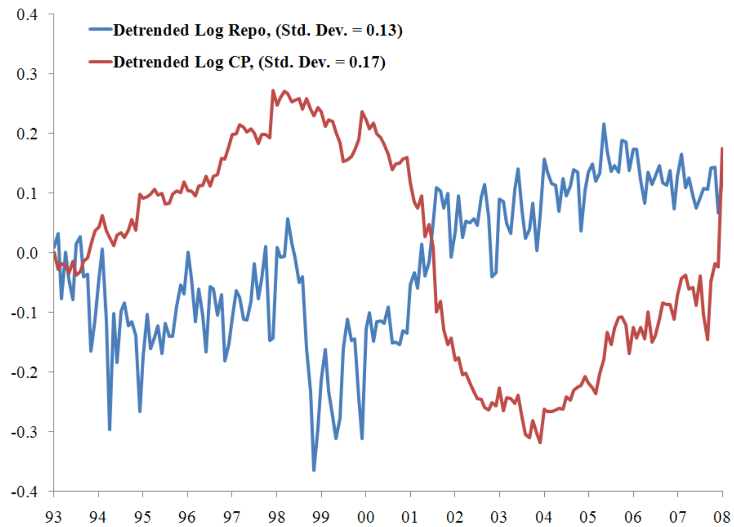

Figure 2.2: Out-of-sample detrended series of primary dealer overnight repos and financial commercial paper outstanding, 1/1993-12/2007

-0.46 between 1993 and 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, 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 (or equivalent), 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 and Bloomberg.

2.2. In-Sample Forecasting Regressions

Our in-sample analysis entails a set of regressions of nominal exchange rate growth on lagged forecasting variables. 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. We also include control variables, such as the U.S. short-term interest rate and the interest rate differential between a particular currency and the U.S. dollar. The time period under consideration is 1993-2007.

2.2.1. OLS Regressions

As a preliminary exercise, we considering simple OLS regressions of monthly nominal exchange rate growth on one-month lags of our two credit aggregates. The results (see Table 1A) indicate that at least one of the two short-term credit aggregates is statistically significant for 9 out of 9 advanced countries and 11 out of 14 emerging countries. In all of these cases, the significant variable enters the regression with a positive sign, implying that an increase in U.S. dollar funding liquidity this month forecasts a U.S. dollar appreciation over the next month. Since our sample of cross rates includes both high and low interest rate countries, this suggests that the forecasting power of the liquidity variables derive from a source different from the more familiar carry trade incentives. For some countries, the economic power of the forecasts is substantial: for example, the lagged credit aggregates forecast 8.3% of the monthly variation in the New Zealand dollar exchange rate growth.

2.2.2. Panel Regressions

Since our short-term credit aggregates forecast U.S. dollar appreciations against all currencies, we may conduct our investigation in the context of a panel regression. Given the nature of our panel, it is possible that the prediction errors are correlated both among different dollar cross rates in the same time period and different time periods within the same cross rate. Hence, we calculate standard errors which allow for two dimensions (currency and time) of within-cluster correlation (see Cameron, Gelbach and Miller, 2006; Thompson, 2006; and Petersen, 2008).

The results from our monthly panel regressions are displayed in Table 1B (for the sample of advanced countries) and Table 1C (for the whole sample of countries). We also provide the results at a weekly and quarterly frequency in Table 1D.[5] The panel specifications echo the same message as our country-by-country OLS regressions: High U.S. dollar liquidity today tends to be followed by U.S. dollar appreciation in the future. For advanced countries, columns (i)-(ii) of Table 1B demonstrate that both credit aggregates are highly statistically significant forecasters of monthly exchange rate growth, controlling for lagged exchange rate growth. Columns (iv)-(viii) show that the statistical significance of the regression coefficients of repo and commercial paper is preserved as one includes lags of common controls, including the interest rate differential (or "carry," defined as the difference between the foreign short-term interest rate and the U.S. short-term interest rate), the VIX implied volatility index, and the stock market return differential (difference between the annual return on the foreign stock market and the annual return on the U.S. stock market). We also control for the interaction of the VIX with the carry and the interaction of the TED spread (difference between Libor and U.S. Treasury bill rate) with the carry, following the finding of Brunnermeier, Nagel and Pedersen (2008) that these variables forecast exchange

[5] Since the weekly and quarterly results are qualitatively similar to the results obtained from our monthly regressions, we save space by focusing our discussion on the monthly results.

rate movements related to unwinding of carry trades.[6]

The magnitudes of the regression coefficients of repo and commercial paper are also preserved across all specifications. Economically, a one standard deviation (0.13) increase in detrended repo forecasts a roughly 0.4 percentage point increase in the rate of U.S. dollar appreciation; similarly, a one standard deviation (0.17) increase in detrended commercial paper forecasts a 0.7 percentage point increase in the rate of dollar appreciation over the following month. It is also notable that the addition of controls has only a limited impact on the explanatory power of the regression: the adjusted R-squared statistic increases from 3.7% to 4.6% as one accounts for the full set of controls.

We emphasize that the power of our regressors, U.S. dollar repos and commercial paper, stems from their 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 of adjusted R-squared for month-ahead to year-ahead forecast horizons. We see that the time-series explanatory power of the regression increases from 3.7% to 8.8% for quarter-ahead forecasts and to 15.9% for six-months-ahead forecasts. The highest explanatory power is obtained at the ten-month horizon where our two credit aggregates are able to forecast nearly 22.5% of the time-series variation in exchange rate growth.

Table 1C displays the panel regression results for the sample of both advanced and emerging countries. We see that lagged commercial paper continues to be a robust forecaster of exchange rate growth across all specifications (i)-(viii) while lagged repo becomes significant only when one includes the full set of controls in column (viii). This finding is consistent with the OLS regressions of Table 1A, which suggest that the predictive ability of repos is strongest for the advanced

[6] The results are also robust to the inclusion of lagged measures of money supply, which are not statistically significant in the regressions (these additional results can be obtained from the authors). This suggests that our results are not driven by the quantity of U.S. dollars but by the type of the investor holding these dollars. Section 3 formalizes this intuition within a simple theoretical framework.

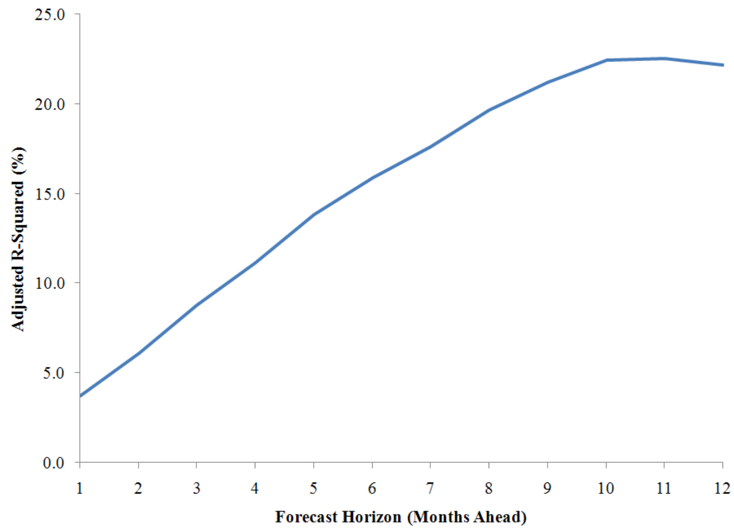

Figure 2.3: Forecasting exchange rate growth several months ahead. Time-series explanatory power in the panel of 9 advanced countries, 1/1993-12/2007.

countries. Accordingly, the combined explanatory power of our credit aggregates is lower for the whole sample of countries, where trends and interest rate differentials tend to play a greater role (see columns (ii)-(iii)).

2.2.3. Funding Liquidity Channel and Carry Trade Channel

In addition to uncovering a new funding liquidity channel of exchange rate determination, our panel regressions confirm the role of the more familiar carry trade channel for the sample of advanced countries (Table 1B). Specifically, the effect of the interest rate differential on the U.S. dollar cross rates is negative and highly significant. That is, the U.S. dollar tends to depreciate when the U.S. dollar interest rate is low relative to the foreign interest rate. This finding is consistent with the usual carry trade mechanism that rests on flows of speculative capital from low to high interest rate countries (see e.g. Jylha and Suominen, 2009). But while the carry trade channel appears to be a strong factor in determining exchange rate movements, it is notably separate from the funding liquidity channel that is the focus of our paper.

The unpredictable nature of the carry trade channel outside of advanced countries is exemplified in our panel regression for the whole sample of 23 countries, where the sign of the interest differential term is surprisingly *positive* and significant. Note that although this finding is at variance with the usual carry trade mechanism, it is nevertheless consistent with U.S. dollar funding liquidity being a window on risk premia on dollar-funded risky positions across the world. All told, we regard the negative coefficient of the interest rate differential 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.

2.3. 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 2. In order to exploit both time and cross-sectional variation in the data, the coefficient estimates for each country are generated using the panel specification of Table 1B. 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 our liquidity model against two benchmarks (restricted models) that are standard in the literature on out-of-sample forecasting: (1) random walk and (2) first-order autoregression.[7] 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 Diebold-Mariano/West tests that employ the unadjusted statistic $MSE_r - MSE_u$.[8] 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 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 2 indicate that the funding liquidity model outperforms both benchmarks for 8 out of 9 advanced countries and 6 out of 14 emerging countries.

2.4. 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 in dollar funding liquidity forecast dollar appreciations, then one would expect increases in (say) euro funding liquidity to forecast euro appreciations.

Table 3 displays the results from simple monthly fixed-effects panel regressions using short-term credit aggregates from the euro and yen repo markets and the exchange rates of our 9 developed countries. 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

[7] The results are also robust to tests against other common benchmarks such as random walk with a drift.

[8] See Diebold and Mariano (1995) and West (1996).

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.

2.5. 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. 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. In addition, it is possible that the policy actions (such as the FX swap agreements among central banks) in response to the malfunctioning of foreign exchange markets lead to significantly different determination of risk premia in the crisis compared to normal times.

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 (i), recursively for 1/1993-11/2009 and plot the t-statistics of lagged repo and

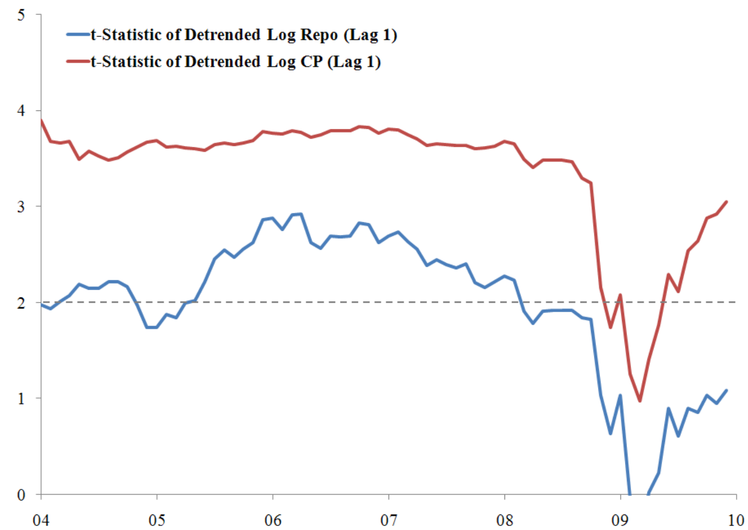

Figure 2.4: Statistical significance of lagged U.S. credit aggregates as predictors of the U.S. dollar exchange rate growth. The t-statistics are obtained from recursive panel regressions of exchange rate growth on lagged repo and lagged commercial paper with standard errors clustered by currency and month (see column (i) of Table 1B). The critical value 1.96 corresponds to significance at 5% level.

lagged financial commercial paper from these regressions. The figure confirms our result that both repo and commercial paper are statistically significant forecasters of the U.S. dollar exchange rate growth over the baseline period. Following the Lehman bankruptcy, however, the statistical significance of lagged repos deteriorates substantially. The statistical significance of lagged commercial paper, on the other hand, revives in 2009.

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 recent 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 showed that the forecasting power of our funding liquidity variables is separate from the usual "carry trade" explanation for exchange rates, which emphasizes the relative attractiveness of currencies of high interest rate countries. In particular, we showed that expansions in U.S. dollar funding aggregates forecast appreciations of the dollar against both high and low-yielding currencies. Thus, the rationale for our findings is very different 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 funding conditions are favorable that 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 asset pricing framework, which illustrates how fluctuations in balance sheet constraints may lead to time-variation in effective risk aversion. We look at the world from the perspective of U.S. dollar-based financial investors who can trade freely in both international and domestic markets.[9] In particular, we assume that this group is spanned by two types of investors: highly leveraged financial intermediaries such as large investment banks (active investors), and less leveraged financial institutions such as commercial banks,

[9] That is, we conduct the analysis in a partial equilibrium setting.

insurance companies, and finance arms of non-financial corporations (passive investors). We think that this set of financial institutions constitutes a reasonably realistic representation of the dollar-based investors who hold internationally diversified securities portfolios and have substantial presence in foreign exchange markets. For expositional simplicity, we seek to remain agnostic about the actual assets held on the investors' balance sheets and begin by assuming that the foreign portfolio is invested in riskless bonds. This assumption allows us to isolate the risk that stems from fluctuations in exchange rates.

3.1. Leveraged Financial Intermediaries

Consider a leveraged financial intermediary (A) that manages its leverage Actively in the U.S. dollar funding market and trades freely in domestic and international assets. Suppose that the foreign portfolio is invested in riskless bonds with holding period rate of return $r^i_{f,t}$, and that U.S. dollar funding is riskless at rate $r^{US}_{f,t}$. Thus, the only risk in this investment strategy is the movement of the exchange rate of the foreign currency relative to the U.S. dollar, ε^i_t. Note that ε^i_t denotes the dollars that can be bought with foreign currency, and is the reciprocal of the definition of exchange rate used so far.[10] The excess return to this strategy is given by:

$$r^i_{t+1} \equiv \left(1 + r^i_{f,t}\right) \frac{\varepsilon^i_{t+1}}{\varepsilon^i_t} - \left(1 + r^{US}_{f,t}\right), \tag{3.1}$$

We suppose that intermediaries 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).[11] Denoting by y^A_i the share of the active intermediary's wealth w^A_t in position i, the investment problem is:

$$\max_{\mathbf{y}^A_t} E_t \left(\mathbf{y}^{A\prime}_t \mathbf{r}_{t+1}\right) \quad s.t.\ VaR_t \leq w^A_t,$$

[10] This change of notation is made for expositional purposes.

[11] Adrian and Shin (2008a) provide a microeconomic foundation for the Value-at-Risk constraint.

where $\mathbf{r}_{t+1}$ is a vector of excess returns. If VaR_t is a multiple κ of equity volatility, the risk constraint becomes $w_t^A \kappa \sqrt{Var_t\left(\mathbf{y}_t' \mathbf{r}_{t+1}\right)} \leq w_t^A$. By risk-neutrality, this constraint binds with equality. It follows that the Lagrangian is:

$$\mathcal{L}_t = E_t\left(\mathbf{y}_t^{A\prime} \mathbf{r}_{t+1}\right) - \phi_t \left[\kappa \sqrt{Var_t\left(\mathbf{y}_t^{A\prime} \mathbf{r}_{t+1}\right)} - 1\right],$$

with the first order condition:

$$\mathbf{y}_t^A = \frac{1}{\kappa \phi_t} \left[Var_t\left(\mathbf{r}_{t+1}\right)\right]^{-1} E_t\left(\mathbf{r}_{t+1}\right). \tag{3.2}$$

From (3.2), we see that the asset demands of the leveraged intermediaries are identical to the standard CAPM choices, but where the risk-aversion parameter is the scaled Lagrange multiplier $\kappa \phi_t$ 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 $\kappa \phi_t$. As the balance sheet constraint binds harder, leverage must be reduced.[12]

Note that, by the binding VaR constraint,

$$\kappa \sqrt{Var_t\left(\mathbf{y}_t^{A\prime} \mathbf{r}_{t+1}\right)} = \frac{1}{\phi_t} \sqrt{E_t\left(\mathbf{r}_{t+1}\right)' \left[Var_t\left(\mathbf{r}_{t+1}\right)\right]^{-1} E_t\left(\mathbf{r}_{t+1}\right)} = 1,$$

which implies that the Lagrange multiplier is given by:

$$\phi_t = \sqrt{E_t\left(\mathbf{r}_{t+1}\right)' \left[Var_t\left(\mathbf{r}_{t+1}\right)\right]^{-1} E_t\left(\mathbf{r}_{t+1}\right)}. \tag{3.3}$$

That is, the tightness of the balance sheet constraint is proportional to the generalized Sharpe ratio in the economy.

3.2. Equilibrium Pricing

We assume that the passive (P) group of dollar-based investors has constant relative risk aversion γ. Their portfolio choice is:

$$\mathbf{y}_t^P = \frac{1}{\gamma} \left[Var_t\left(\mathbf{r}_{t+1}\right)\right]^{-1} E_t\left(\mathbf{r}_{t+1}\right). \tag{3.4}$$

[12] Danielsson, Shin and Zigrand (2008) solve for the rational expectations equilibrium of a continuous time dynamic model along these lines.

Market clearing implies:

$$\mathbf{y}_t^A \frac{w_t^A}{w_t^A + w_t^P} + \mathbf{y}_t^P \frac{w_t^P}{w_t^A + w_t^P} = \mathbf{s}_t, \tag{3.5}$$

where the vector $\mathbf{s}_t$ denotes the net supply of investment opportunities absorbed by dollar-based investors. Plugging the two asset demands (3.2) and (3.4) in the market clearing condition, and rearranging, one obtains:

$$\begin{aligned} E_t(\mathbf{r}_{t+1}) &= Var_t(\mathbf{r}_{t+1})\,\mathbf{s}_t \frac{w_t^A + w_t^P}{w_t^A/(\kappa\phi_t) + w_t^P/\gamma} \\ &= Cov_t\left(\mathbf{r}_{t+1}, r_{t+1}^W\right)\Gamma_t, \end{aligned} \tag{3.6}$$

where $r_{t+1}^W = \mathbf{r}'_{t+1}\mathbf{s}_t$ is the return on the aggregate wealth portfolio and $\Gamma_t = \frac{w_t^A + w_t^P}{w_t^A/(\kappa\phi_t) + w_t^P/\gamma}$ denotes the effective risk aversion of dollar-based investors.

With the equilibrium returns in hand, we can express Γ_t in terms of observable balance sheet components. Begin by rewriting Γ_t as:

$$\Gamma_t = \gamma\left[1 + \frac{w_t^A}{w_t^P}\left(1 - \frac{\Gamma_t}{\kappa\phi_t}\right)\right]. \tag{3.7}$$

In order to obtain an expression for $\frac{\Gamma_t}{\kappa\phi_t}$, we plug the equilibrium expression (3.6) in the intermediary's portfolio choice (3.2), which yields:

$$\mathbf{y}_t^A = \frac{\Gamma_t}{\kappa\phi_t}\left[Var_t(\mathbf{r}_{t+1})\right]^{-1} Cov_t\left(\mathbf{r}_{t+1}, r_{t+1}^W\right) = \frac{\Gamma_t}{\kappa\phi_t}\mathbf{s}_t.$$

Summing over individual positions, we get:

$$\frac{\Gamma_t}{\kappa\phi_t} = \frac{\sum_i y_{i,t}^A}{\sum_i s_{i,t}}. \tag{3.8}$$

By balance sheet identity, the value of risky securities holdings must equal the value of equity (wealth) plus the value of debt:

$$w_t^A \sum_i y_{i,t}^A = w_t^A + debt_t^A,$$

which implies that one can define the financial leverage of active intermediaries as:

$$lev_t^A \equiv 1 + \frac{debt_t^A}{w_t^A} = \sum_i y_{i,t}^A,$$

and the leverage of all financial institutions as:[13]

$$lev_t^{A\&P} \equiv 1 + \frac{debt_t^A + debt_t^P}{w_t^A + w_t^P} = \sum_i s_{i,t}.$$

Using this notation, we substitute (3.8) into (3.7) to get:

$$\Gamma_t = \gamma \left[1 + \frac{w_t^A}{w_t^P} \left(1 - \frac{lev_t^A}{lev_t^{A\&P}} \right) \right]. \tag{3.9}$$

Equation (3.9) states that the time-variation in the effective risk aversion of dollar-based investors can be represented by fluctuations in the leverage of highly leveraged intermediaries relative to the leverage of the market, scaled by the wealth of leveraged intermediaries relative to the wealth of passive investors. Specifically, an increase in active intermediaries' leverage is associated with a decrease in effective risk aversion (since $lev_t^A > lev_t^{A\&P}$). The greater the wealth share of intermediaries, the greater the impact of their leverage on Γ_t.

Plugging (3.1) into (3.6), one obtains:

$$E_t \left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i} \right) = \frac{1 + r_{f,t}^{US}}{1 + r_{f,t}^i} + Cov_t \left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}, r_{t+1}^W \right) \Gamma_t, \tag{3.10}$$

and by (3.9):

$$E_t \left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i} \right) = \frac{1 + r_{f,t}^{US}}{1 + r_{f,t}^i} + Cov_t \left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}, r_{t+1}^W \right) \underbrace{\gamma \left[1 + \frac{w_t^A}{w_t^P} \left(1 - \frac{lev_t^A}{lev_t^{A\&P}} \right) \right]}_{\Gamma_t} \tag{3.11}$$

Thus, an increase in the leverage of dollar-funded leveraged intermediaries forecasts an appreciation of the dollar against currencies that comove positively with their wealth portfolio.

[13] In a closed economy, the leverage of the aggregate, $lev_t^{A\&P}$, would be one.

To the extent that our short-term dollar credit aggregates—primary dealer repos and financial commercial paper outstanding—measure the availability of U.S. dollar leverage in the financial system, one may expect them to be linked to the effective risk aversion of dollar-funded investors, Γ_t, and hence, to the equilibrium returns on dollar-funded positions, including risky positions in foreign currencies. In the following section, we conduct cross-sectional asset pricing tests to more formally investigate this hypothesis.

4. Reconciling Theory and Empirics

Can our simple no-arbitrage pricing model explain why U.S. primary dealer repos and financial commercial paper forecast the dollar exchange rate? To answer this question, we proceed in three steps: First, as a preliminary step, we use data on U.S. financial institution balance sheets to construct a measure of Γ_t. We relate this theoretically motivated measure of effective risk aversion to our short-term credit aggregates—repos and financial commercial paper—which are observable at a higher frequency. Second, we investigate the extent to which effective risk aversion, as captured by Γ_t, explains the forecasting ability of repos and commercial paper for dollar cross rates. We do this by conducting simple time-series tests for individual exchange rates. Third, we estimate the asset pricing model (3.11) in the cross-section of dollar exchange rates and test if the idiosyncratic (residual) exchange rate variation remains predictable by primary dealer repos and financial commercial paper outstanding. If the idiosyncratic variation is not predictable, we may conclude that our higher-frequency measures of funding liquidity forecast exchange rates because they are related to the effective risk aversion of dollar-funded investors; that is, they contain information about systematic risk premia.

4.1. Measuring Effective Risk Aversion

Taking the cue from our theoretical model (3.9), we construct the following measure of effective risk aversion:

$$\hat{\Gamma}_t = 1 + \frac{\text{Primary Dealer Equity}_t}{\text{All Financials Equity}_t - \text{Primary Dealer Equity}_t}\left(1 - \frac{\text{Primary Dealer Leverage}_t}{\text{All Financials Leverage}_t}\right). \quad (4.1)$$

That is, we let the primary dealers represent the active leveraged intermediaries while the other U.S. financial institutions represent the passive investors (who are nevertheless diversified internationally). We obtain the data on market equity from CRSP and merge them with data on market leverage from Compustat.[14] Since the leverage data is available only at a quarterly frequency, we interpolate it to obtain a monthly time series.

We investigate the extent to which our forecasting variables from Section 2 are related our theoretically motivated measure of effective risk aversion by projecting $\hat{\Gamma}_t$ onto detrended log repos and detrended log financial commercial paper. A simple OLS regression shows that repo and commercial paper explain over 62% of the variation in $\hat{\Gamma}_t$ with both regressors being highly statistically significant (t-statistics in parentheses):

$$\hat{\Gamma}_t = \underset{(463.03)}{0.890} - \underset{(-6.58)}{0.082}\text{Repo}_t - \underset{(-18.10)}{0.177}\,\text{CP}_t + \text{error}_{t+1}.$$

Consistent with our intuition, the negative slope coefficients of repo and commercial paper indicate that an increase in either short-term credit aggregate is associate with a decrease in effective risk aversion. The fitted values from this regression are displayed in Fig. 4.1 along with the original series $\hat{\Gamma}_t$. Note the sharp peaks in effective risk aversion during the recent financial turmoil as well as during the Long Term Capital Management crisis in 1998.

[14] Note that "All Financials" refer to all U.S. financial firms reported in the CRSP/Compustat database.

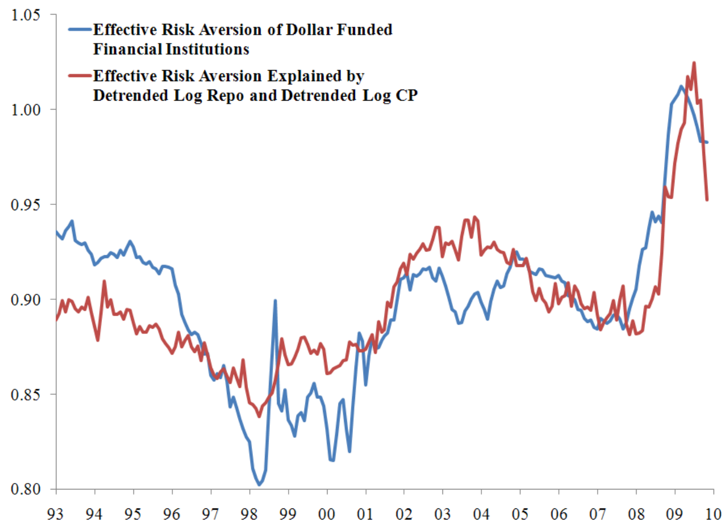

Figure 4.1: Effective risk aversion of U.S. dollar funded financial institutions and the projection of this series onto primary dealer repos and financial commercial paper outstanding

4.2. Effective Risk Aversion and Predictability of Individual Exchange Rates

Do primary dealer repos and financial commercial paper forecast exchange rates because they contain information about effective risk aversion? To answer this question, we regress the exchange rate growth $\frac{\varepsilon_{t+1}^i - \varepsilon_t^i}{\varepsilon_t^i}$ currency by currency on a constant and the one-month lag of effective risk aversion $\hat{\Gamma}_t$, and save the residuals (denoted by ν_{t+1}^i). We then use these residuals as dependent variables in the regressions (for each currency i):

$$\nu_{t+1}^i = a_0^i + a_i^\nu \nu_t^i + a_i^{\mathrm{Repo}} \mathrm{Repo}_t + a_i^{\mathrm{CP}} \mathrm{CP}_t + \mathrm{error}_t^i, \tag{4.2}$$

where Repo_t and CP_t again denote detrended log repos and detrended log financial commercial paper, respectively.

We test the Granger restriction that $a_i^{\mathrm{Repo}} = a_i^{\mathrm{CP}} = 0$. If this hypothesis is not rejected at conventional confidence levels, we may attribute the forecasting ability of primary dealer repos and financial commercial paper to their association with effective risk aversion, as suggested by our simple theoretical framework.

The test results are dislayed in Table 4, which reports the F-statistics along with the associated p-values for individual exchange rates. The p-values are greater than 0.10 for 9 out of 9 advanced country currencies, which suggests that repos and commercial paper forecast exchange rate growth of advanced countries only because they contain information about effective risk aversion, $\hat{\Gamma}_t$. For emerging country currencies, the p-values are greater than 0.10 for 10 out of 14 currencies, suggesting that most of the forecasting ability of repo and commercial paper can be attributed to effective risk aversion also outside of advanced currencies.

4.3. Risk or Mispricing? Estimating Cross-Sectional Prices of Risk

Estimation Framework. Can the forecasting ability of primary dealer repos and commercial paper be explained by systematic fluctuations in risk premia? To answer this question, we estimate (3.11) in the cross-section of exchange rates. Replacing expectations in (3.11) by realizations, one obtains:

$$\underbrace{\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}}_{\substack{\text{Exchange Rate}\\ \text{Appreciation}}} - \underbrace{\frac{1+r_{f,t}^{US}}{1+r_{f,t}^i}}_{\substack{\text{Interest Rate}\\ \text{Carry}}} = \underbrace{Cov_t\left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}, r_{t+1}^W\right)\Gamma_t}_{\substack{\text{FX Risk}\\ \text{Premium}}} + \underbrace{z_{t+1}^i}_{\substack{\text{FX}\\ \text{Risk}}}, \tag{4.3}$$

where the FX risk is defined as $z_{t+1}^i \equiv r_{t+1}^i - E_t\left(r_{t+1}^i\right)$. We can go further by decomposing the FX risk z_{t+1}^i into a component that is correlated with unpredictable returns to the wealth portfolio, $r_{t+1}^W - E_t\left(r_{t+1}^W\right)$, and a return pricing error ξ_{t+1}^i that is orthogonal to $r_{t+1}^W - E_t\left(r_{t+1}^W\right)$:[15]

$$z_{t+1}^i = \beta_t^i\left[r_{t+1}^W - E_t\left(r_{t+1}^W\right)\right] + \xi_{t+1}^i, \tag{4.4}$$

for some β_t^i. Note that, by construction, $E\left[\xi_{t+1}^i | X_t, r_{t+1}^W - E_t\left(r_{t+1}^W\right)\right] = 0$. It follows that,

$$\begin{aligned} Cov_t\left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}, r_{t+1}^W\right)\Gamma_t &= Cov_t\left(\beta_t^i\left[r_{t+1}^W - E_t\left(r_{t+1}^W\right)\right] + \xi_{t+1}^i, r_{t+1}^W\right)\Gamma_t \\ &= \beta_t^i Cov_t\left(r_{t+1}^W - E_t\left(r_{t+1}^W\right), r_{t+1}^W\right)\Gamma_t \\ &= \beta_t^i Var_t\left(r_{t+1}^W\right)\Gamma_t, \end{aligned} \tag{4.5}$$

such that $\beta_t^i = Cov_t\left(\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}, r_{t+1}^W\right)/Var_t\left(r_{t+1}^W\right)$ is the beta of currency i with the wealth portfolio. We can now use (4.4) and (4.5) to rewrite (4.3) as:

$$\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i} - \frac{1+r_t^{US}}{1+r_t^i} = \beta_t^i Var_t\left(r_{t+1}^W\right)\Gamma_t + \beta_t^i\left[r_{t+1}^W - E_t\left(r_{t+1}^W\right)\right] + \xi_{t+1}^i, \tag{4.6}$$

By (3.9) and (4.1), we can express Γ_t in terms of a constant and a time-varying component:

$$\Gamma_t = \gamma + \gamma\left(\hat{\Gamma}_t - 1\right).$$

[15] See Adrian and Moench (2008).

Assuming constant conditional variances and covariances, the price of risk $Var\left(r^W_{t+1}\right)\Gamma_t$ can be written as:

$$Var\left(r^W_{t+1}\right)\Gamma_t = \lambda_0 + \lambda_1\left(\hat{\Gamma}_t - 1\right). \tag{4.7}$$

where our theory predicts that:

$$\lambda_0 = \lambda_1 = Var\left(r^W_{t+1}\right)\gamma. \tag{4.8}$$

Note that the first of the two equalities in (4.8) is a statement about the level of the risk premium while the second is a statement about the responsiveness of the risk premium to time-variation in effective risk aversion. Since our aim is not to explain the forward premium puzzle—i.e., why high interest rate currencies often fail to depreciate relative to low interest rate currencies—but merely to explain why our short-term credit aggregates forecast exchange rate growth, we will not impose the restriction $\lambda_0 = \lambda_1$ in our estimation. Using this notation, we express (4.6) as:

$$\frac{\varepsilon^i_{t+1}}{\varepsilon^i_t} - \frac{1 + r^{US}_t}{1 + r^i_t} = \underbrace{\beta^i\left[\lambda_0 + \lambda_1\left(\hat{\Gamma}_t - 1\right)\right]}_{\substack{\text{FX Risk} \\ \text{Premium}}} + \underbrace{\beta^i\left[r^W_{t+1} - E_t\left(r^W_{t+1}\right)\right]}_{\substack{\text{Systematic} \\ \text{FX Risk}}} + \underbrace{\xi^i_{t+1}}_{\substack{\text{Idiosyncratic} \\ \text{FX Risk}}}. \tag{4.9}$$

We estimate the cross-sectional pricing model (4.9) 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). We consider two alternative proxies for r^W_{t+1}, the wealth portfolio of our internationally diversified dollar-based investor: First, we proxy r^W_{t+1} by the excess U.S. dollar return on the MSCI world equity index, which we believe is a reasonable measure of the systematic risk faced by dollar-funded global financial institutions. Second, we proxy r^W_{t+1} by the excess return on a dollar-funded FX portfolio given by the first principal component of carry returns across all countries in our sample. This latter proxy emphasizes our focus on foreign exchange risk.

Estimation Results. Table 5 summarizes the results from the estimation of (4.9) by reporting the point estimates that determine the prices of risk in our two specifications. The first row shows that the world equity market risk is priced in the cross-section of currency returns and the price of risk varies significantly over time: as predicted by the theory, both the constant λ_0 and the loading λ_1 on the measure of effective risk aversion are positive and highly significant. However, the last column shows that one can reject the theory's restriction (4.8) that the two coefficients are equal.[16] The second row reports the results for the specification where the risk factor is the return on the dollar-funded FX portfolio. The qualitative message of the results is similar: The price of FX risk exhibits significant variation over time as given by the significant positive loading on the measure of effective risk aversion. Also, we can again reject the hypothesis that $\lambda_0 = \lambda_1$.

The results of Table 5 confirm our earlier intuition that the effective risk aversion of dollar-based investors matters for the pricing of U.S. dollar cross rates due to its association with the market-wide risk premium. Specifically, an increase in effective risk aversion, as captured by observable changes in key balance sheet components of dollar-based financial institutions, is associated with an increase in required returns on risky positions held by these institutions, including those in foreign currencies.[17]

To test the hypothesis that the idiosyncratic FX risk ξ_{t+1}^{i} in (4.9) is not fore-

[16] It is nonetheless possible to use the estimates of λ_0, λ_1 and the time-series average of $\hat{\Gamma}_t$ to infer sample estimates for the average price of risk, the average effective risk aversion, and the coefficient of relative risk aversion of passive investors: By (4.7), $E\left[Var\left(r_{t+1}^{MSCI}\right)\Gamma_t\right]$ $= 0.097 + 0.800 \times (0.889 - 1) = 0.00046$. Since $Var\left(r_{t+1}^{MSCI}\right) = 0.00029$ over our sample, we get $E\left[\Gamma_t\right] = \frac{0.00046}{0.00029} = 1.59$. Furthermore, since $\gamma = \frac{E[\Gamma_t]}{E\left[\hat{\Gamma}_t\right]}$, one obtains $\gamma = \frac{1.59}{0.889} = 1.79$.

[17] It is important to note that despite our success in explaining the forecasting ability of U.S. dollar repos and financial commercial paper by their association with systematic risk premia, our no-arbitrage framework could not accurately price the cross-section of returns if we imposed the theory's restriction that $\lambda_0 = \lambda_1$. In other words, the restricted pricing kernel could explain the time-variation in the forward risk premium but not the average level of the risk premium.

castable by U.S. dollar repos and financial commercial paper outstanding, we run the predictive regression:

$$\xi^i_{t+1} = b^0_i + b^\xi_i \xi^i_t + b^{\text{Repo}}_i \text{Repo}_t + b^{\text{CP}}_i \text{CP}_t + \text{error}_{t+1}, \tag{4.10}$$

and test Granger causality by the joint test $b^{\text{Repo}}_i = b^{\text{CP}}_i = 0$. The results are displayed in Table 6 for each currency (rows) and for each of the two model specifications (columns). The first column corresponds to the specification where the risk factor is the excess dollar return on the MSCI world equity index and the second column corresponds to the specification with the dollar-funded FX portfolio. The large p-values in brackets indicate that U.S. primary dealer repos and financial commercial paper have little forecasting ability for idiosyncratic changes in dollar cross rates: In both specifications, the joint significance test rejects the null hypothesis only for two or three of the 23 dollar cross rates, implying that our arbitrage-free framework does a good job in explaining the forecasting ability of repos and commercial paper for the rest of the cross-section.

In sum, the cross-sectional evidence supports our view that the forecastability of exchange rate growth uncovered in Tables 1-2 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 short-term 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 dollar cross rates. We have shown how this result goes beyond the usual "carry trade" story, in favor of a parallel funding liquidity channel as expressed in short-term credit aggregates. 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.

Second, motivated by our new empirical evidence on forecastability, we have constructed a simple asset pricing framework where the effective risk aversion of financial investors varies over time with observable balance sheet components. Estimation of the model in the data suggests that the forecastability of exchange rates by our short-term credit aggregates is linked to time-variation in systematic risk premia.

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 is needed to explore this hypothesis further.

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Table 1A: 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 Newey-West t-statistics (using 4 lags) in parentheses; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. The sample period is 1/1993- 12/2007.

Dep. Variable	Independent Variables					
Exchange Rate Growth	Detrended Log Repo (Lag 1)		Detrended Log CP (Lag 1)		Constant	R^2
Australia	4.669**	(2.394)	3.419***	(3.077)	-0.105	6.6%
Canada	1.382	(0.914)	2.022**	(2.550)	-0.121	4.1%
Germany	1.320	(0.625)	2.977***	(2.698)	-0.071	4.5%
Japan	4.686**	(2.035)	0.993	(0.754)	-0.001	2.0%
New Zealand	6.252***	(2.742)	4.034***	(3.487)	-0.175	8.3%
Norway	1.516	(0.791)	2.824***	(2.704)	-0.083	3.5%
Sweden	2.773	(1.270)	3.127***	(2.881)	-0.025	4.3%
Switzerland	2.143	(0.975)	2.480**	(2.115)	-0.106	2.7%
UK	2.260	(1.526)	1.839**	(2.422)	-0.137	3.2%
Chile	-0.129	(-0.063)	2.459**	(2.233)	0.169	3.7%
Colombia	-3.532	(-1.365)	3.727***	(2.880)	0.596***	7.0%
Czech Republic	0.050	(0.019)	3.703**	(2.525)	-0.203	4.3%
Hungary	0.556	(0.264)	4.673***	(4.509)	0.453**	7.9%
India	0.787	(0.510)	1.677***	(3.069)	0.248	2.3%
Indonesia	9.130	(1.189)	9.714	(1.439)	1.328	2.6%
Korea	2.540	(0.729)	2.851	(1.215)	0.187	1.4%
Philippines	-0.425	(-0.157)	2.476*	(1.696)	0.315	2.3%
Poland	-2.028	(-1.067)	3.302***	(2.723)	0.288	4.2%
Singapore	1.090	(0.662)	1.472**	(2.200)	-0.059	3.0%
South Africa	3.494	(1.056)	4.195**	(2.432)	0.532*	3.8%
Taiwan	2.202*	(1.715)	1.131	(1.489)	0.147	3.3%
Thailand	-1.209	(-0.289)	2.927	(1.404)	0.226	2.0%
Turkey	-5.009	(-1.174)	11.580***	(5.627)	2.957***	10.1%

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

This table uses panel regressions with currency and time fixed effects to forecast exchange rate growth. The dependent variable is the monthly growth of the U.S. dollar bilateral exchange rate against 9 advanced-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 and month 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)	3.000**	2.952**		2.775**	3.131**	3.399**	3.473**	3.723***
	(2.281)	(2.218)		(2.062)	(2.230)	(2.406)	(2.518)	(2.663)
Detrended Log CP (Lag 1)	4.231***	4.191***		3.949***	3.973***	4.980***	5.081***	5.115***
	(3.685)	(3.588)		(3.371)	(3.383)	(2.965)	(3.043)	(3.087)
Exch. Rate Growth (Lag 1)		0.005	0.034	0.004	-0.005	-0.005	-0.006	-0.007
		(0.133)	(0.885)	(0.110)	(-0.111)	(-0.129)	(-0.143)	(-0.171)
Interest Rate Differential (Lag 1)			-0.103***	-0.037*	-0.048***	-0.057***	-0.054***	-0.061***
			(-3.377)	(-1.649)	(-3.511)	(-3.670)	(-3.553)	(-4.122)
Stock Mkt. Ret. Dif. Ann. (Lag 1)					-0.006	-0.005	-0.005	-0.004
					(-1.037)	(-0.871)	(-0.883)	(-0.734)
U.S. Interest Rate						-0.119	-0.118	-0.119
						(-0.766)	(-0.759)	(-0.764)
VIX Growth Annual (Lag 1)							-0.001	0.001
							(-0.233)	(0.148)
Signed VIX Growth Ann. (Lag 1)							-0.001	-0.002
							(-0.369)	(-0.680)
TED Growth Annual (Lag 1)								-0.003
								(-1.421)
Signed TED Growth Ann. (Lag 1)								0.001**
								(2.096)
Constant	-0.038	-0.047	-0.056	-0.035	-0.032	0.436	0.436	0.490
	(-0.310)	(-0.376)	(-0.411)	(-0.276)	(-0.247)	(0.697)	(0.702)	(0.786)
# Countries	9	9	9	9	9	9	9	9
Adjusted R^2	3.7%	3.7%	0.9%	3.8%	4.3%	4.4%	4.2%	4.6%

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

This table uses panel regressions with currency and time 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 and month 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)	1.501	1.173		1.963	1.948	2.073	2.075	2.294*
	(1.140)	(0.927)		(1.612)	(1.536)	(1.519)	(1.500)	(1.647)
Detrended Log CP (Lag 1)	4.259***	3.671***		3.739***	3.734***	4.145**	4.142**	4.130**
	(3.551)	(3.447)		(3.480)	(3.463)	(2.337)	(2.300)	(2.326)
Exch. Rate Growth (Lag 1)		0.120***	0.081***	0.062**	0.061**	0.061**	0.061**	0.061**
		(3.031)	(2.675)	(2.227)	(2.163)	(2.169)	(2.158)	(2.141)
Interest Rate Differential (Lag 1)			0.053***	0.051***	0.050***	0.050***	0.050***	0.049***
			(14.711)	(15.071)	(11.988)	(12.601)	(12.742)	(12.344)
Stock Mkt. Ret. Dif. Ann. (Lag 1)					-0.001	-0.001	-0.001	-0.000
					(-0.285)	(-0.257)	(-0.245)	(-0.127)
U.S. Interest Rate						-0.046	-0.046	-0.040
						(-0.327)	(-0.327)	(-0.289)
VIX Growth Annual (Lag 1)							-0.000	0.002
							(-0.026)	(0.417)
Signed VIX Growth Ann. (Lag 1)							0.001	0.000
							(0.421)	(0.098)
TED Growth Annual (Lag 1)								-0.003*
								(-1.693)
Signed TED Growth Ann. (Lag 1)								0.001
								(0.740)
Constant	0.303*	0.258	-0.062	-0.011	-0.003	0.179	0.176	0.200
	(1.694)	(1.634)	(-0.508)	(-0.088)	(-0.019)	(0.310)	(0.306)	(0.347)
# Countries	23	23	23	23	23	23	23	23
Adjusted R^2	2.4%	3.8%	6.2%	7.7%	7.5%	7.4%	7.4%	7.5%

Table 1D: Forecasting Quarterly and Weekly Exchange Rate Growth

This table uses panel regressions with currency and time fixed effects to forecast exchange rate growth. The dependent variable is the growth of the U.S. dollar bilateral exchange rate against 9 advanced-country 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 and time 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)		-0.062		-0.030
		(-0.913)		(-1.543)
Detrended Log Repo (Lag1)	5.120	5.914	0.782**	0.800**
	(1.219)	(1.360)	(1.997)	(2.041)
Detrended Log CP (Lag 1)	10.316***	11.265***	1.008***	1.035***
	(2.939)	(2.971)	(3.627)	(3.719)
Constant	-0.105	-0.118	-0.021	-0.022
	(-0.278)	(-0.305)	(-0.709)	(-0.726)
# Countries	9	9	9	9
Adjusted R^2	8.9%	9.2%	0.8%	0.9%

Table 2: 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		
	ΔMSE	$\Delta MSE - Adj.$	p-value	ΔMSE	$\Delta MSE - Adj.$	p-value
Australia	0.466	0.899**	0.005	0.487	0.858***	0.003
Canada	0.051	0.528**	0.020	0.075	0.449**	0.028
Germany	0.116	0.588**	0.035	0.171	0.544**	0.037
Japan	-0.129	0.358	0.186	-0.033	0.344	0.156
New Zealand	0.576	1.030***	0.006	0.656	1.023***	0.002
Norway	0.099	0.584*	0.066	0.174	0.548*	0.069
Sweden	0.235	0.683**	0.018	0.275	0.647**	0.020
Switzerland	-0.019	0.469*	0.099	0.042	0.417*	0.099
UK	0.059	0.523**	0.031	0.097	0.471**	0.027
Chile	0.130	0.502**	0.012	0.138	0.508**	0.035
Colombia	0.429	1.464***	0.002	0.260	0.628**	0.014
Czech Republic	0.100	0.650	0.102	0.252	0.630*	0.080
Hungary	-0.203	1.363***	0.006	0.477*	0.847**	0.011
India	-0.067	0.575***	0.006	0.030	0.402***	0.003
Indonesia	0.335	4.572	0.220	1.697	2.078*	0.074
Korea	-0.518	0.203	0.434	0.112	0.488	0.211
Philippines	-0.192	0.416	0.273	-0.095	0.279	0.223
Poland	-0.613	0.636	0.137	0.135	0.511*	0.075
Singapore	-0.281	0.141	0.307	-0.112	0.267	0.151
South Africa	0.687	1.545***	0.009	0.667	1.033**	0.032
Taiwan	-0.125	0.350	0.123	-0.074	0.301	0.102
Thailand	-0.759	-0.081	0.474	-0.168	0.219	0.333
Turkey	0.894	21.730***	0.000	1.289	1.637***	0.001

Table 3: Evidence from Euro and Yen Repo Markets

This table uses panel regressions with currency and time 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 advanced-country currencies. The forecasting variable is the one-month lag of the year-over-year 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 and month 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)	-0.005	0.148
	(-0.086)	(1.366)
Euro Repos (Annual Growth, Lag1)	0.023**	
	(2.524)	
Yen Repos (Annual Growth, Lag1)		0.010**
		(1.972)
Constant	-0.001	0.850***
	(-1.174)	(7.960)
# Countries	9	9
Adjusted R^2	1.2%	4.2%

Table 4: Predictability of Residual Variation in Exchange Rate Growth

Do primary dealer repos and financial commercial paper forecast exchange rate growth only because they contain information about effective risk aversion? In this table, we test the hypothesis that lagged repo and lagged commercial paper forecast exchange rate growth cannot forecast exchange rate growth beyond what is predictable by lagged effective risk aversion (see equation (4.2). The rows of the table report the F-statistics and the associate p-values (in brackets) for the joint significance of lagged repo and lagged commercial paper, currency by currency; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. The sample period is 1/1993- 12/2007.

	Predictability of Residual FX Variation by Lagged Repo and Lagged CP $H_0 : a^{\mathrm{Repo}} = a^{\mathrm{CP}} = 0$	
Currency	F-statistic	p-value
Australia	1.166	[0.3140]
Canada	1.204	[0.3024]
Germany	0.794	[0.4539]
Japan	1.051	[0.3520]
New Zealand	1.799	[0.1686]
Norway	0.651	[0.5230]
Sweden	0.996	[0.3716]
Switzerland	0.288	[0.7503]
UK	1.469	[0.2330]
Chile	0.762	[0.4681]
Colombia	3.191**	[0.0436]
Czech Republic	1.183	[0.3088]
Hungary	4.963***	[0.0080]
India	1.016	[0.3643]
Indonesia	0.018	[0.9824]
Korea	0.157	[0.8544]
Philippines	0.318	[0.7280]
Poland	2.577*	[0.0789]
Singapore	0.010	[0.9902]
South Africa	0.692	[0.5020]
Taiwan	0.634	[0.5319]
Thailand	0.490	[0.6132]
Turkey	4.026**	[0.0195]

Table 5: Cross-Sectional Prices of Risk

This table reports the results from the estimation of a cross-sectional arbitrage-free asset pricing model (4.9) for U.S. dollar funded investments in 23 foreign currencies. We consider two alternative model specifications. The first row corresponds to the specification where the risk factor is the excess dollar return on the MSCI global equity index. In the second row, the risk factor is the excess return on a dollar-funded foreign exchange portfolio (first principal component from the cross-section of excess carry returns). The two columns display point estimates for the loadings of each alternative risk factor on a constant and lagged effective risk aversion as measured by $\hat{\Gamma}_t - 1$ (see 4.1). The third column tests the hypothesis that the two coefficients are equal. 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.

Price of Risk	λ_0	λ_1	$H_0 : \lambda_0 = \lambda_1$
MSCI Global Equity Return	0.097***	0.800***	[0.0123]**
	(2.814)	(2.514)	
FX Portfolio Return	0.032***	0.027***	[0.0021]***
	(2.786)	(2.667)	

Table 6: Predictability of Idiosyncratic Foreign Exchange Risk

Can the forecasting ability of primary dealer repos and financial commercial paper be explained by systematic fluctuations in risk premia? This table test the hypothesis that the arbitrage-free forecast residuals are not predictable by lagged repos and commercial paper outstanding (see equation (4.10)). We conduct the test for each currency return (rows) and two model specifications (columns). The first column corresponds to the specification where the risk factor is the excess dollar return on the MSCI global equity index. In the second column, the risk factor is the excess return on a dollar-funded foreign exchange portfolio. p-values for the joint significance of lagged repo and commercial paper are reported in brackets; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. The sample period is 1/1993- 12/2007.

	Predictability of Idiosyncratic FX Risk by Lagged Repo and Lagged CP $H_0 : b^{\mathrm{Repo}} = b^{\mathrm{CP}} = 0$	
Currency	Risk Factor = r^{MSCI}	Risk Factor = r^{FX}
Australia	[0.3441]	[0.1942]
Canada	[0.6599]	[0.3551]
Germany	[0.1312]	[0.5652]
Japan	[0.3530]	[0.2996]
New Zealand	[0.0723]*	[0.1104]
Norway	[0.1797]	[0.4285]
Sweden	[0.5932]	[0.7621]
Switzerland	[0.1453]	[0.5378]
UK	[0.2595]	[0.7748]
Chile	[0.7961]	[0.9023]
Colombia	[0.2509]	[0.5174]
Czech Republic	[0.6457]	[0.0118]**
Hungary	[0.1173]	[0.6691]
India	[0.0385]**	[0.0818]*
Indonesia	[0.4988]	[0.9065]
Korea	[0.9553]	[0.9776]
Philippines	[0.6261]	[0.8400]
Poland	[0.5471]	[0.0526]*
Singapore	[0.7745]	[0.7989]
South Africa	[0.5448]	[0.7061]
Taiwan	[0.4070]	[0.3370]
Thailand	[0.4838]	[0.1889]
Turkey	[0.3516]	[0.2989]