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Abstract

We present evidence that fluctuations in the aggregate balance sheets of financial intermediaries forecast exchange rate returns—at weekly, monthly, and quarterly frequencies, both in and out of sample, and for a large set of countries. We estimate prices of risk using a cross-sectional, arbitrage-free asset pricing approach and show that balance sheets forecast exchange rates because of the latter's association with fluctuations in risk premia. We provide a rationale for an intertemporal equilibrium pricing theory in which intermediaries are subject to balance sheet constraints.

Key words: asset pricing, financial intermediaries, exchange rates

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1. Introduction

In market-based financial systems, banking and capital market developments are inseparable. At the margin, all financial intermediaries (including commercial banks) have to borrow in capital markets, since deposits are insufficiently responsive to funding needs. Market-based credit aggregates such as the stock of repurchase agreements (repos) or financial commercial paper outstanding can be expected to provide a window on liquidity in the sense of the availability of credit. To the extent that such market aggregates reflect the risk appetite of financial intermediaries via the associated leverage constraints they face, we may expect the prices of financial assets to impound the pricing consequences of balance sheet constraints. In short, we may conjecture that financial market prices will reflect liquidity conditions through their association with market-wide risk premia.

In this paper, we show that foreign exchange markets are influenced by such funding liquidity conditions. In particular, we show that balance sheet aggregates for US financial intermediaries have forecasting power for the future returns on the US dollar across a wide cross-section of currencies – both for developed countries as well as for developing countries. The forecasting power of our liquidity variables is surprisingly strong. The fluctuations in market-based liability series of financial intermediaries can be shown to explain subsequent returns on exchange rates at weekly, monthly, and quarterly frequencies, both in and out-of sample.

In part, our liquidity channel is related to the familiar forward risk premium for exchange rates and associated "carry trade" incentives.[1] We show, for instance, that the interest rate differential of a currency vis-à-vis the US dollar has forecasting power for the future evolution of its exchange rate against the US dollar. However, liquidity conditions as reflected in balance sheet variables have

[1] 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.

forecasting power beyond such carry trade channels. Controlling for interest rate differentials and for the absolute level of US short-term interest rates, balance sheet growth for US financial intermediaries have incremental value in forecasting future appreciations of the US dollar.

Our favored rationalization for the empirical findings in our paper is in terms of the fluctuations in the risk-bearing capacity of financial intermediaries in the United States. As balance sheets expand and leverage rises, the constraints faced by financial intermediaries loosen, thereby increasing their risk appetite. 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, *growth* of intermediary balance sheets will be associated with *innovations* in risk appetite. When balance sheets expand, there is an increase in risk appetite and risky asset prices are driven up. This drives down the equilibrium risk premium on risky assets, including risky holdings of foreign currency, implying a future depreciation of such risky currencies (i.e. a dollar *appreciation* against such risky currencies). In short, we would expect to see growth of US financial intermediary balance sheets being followed by subsequent dollar appreciations. This is exactly what we find in our empirical investigation. The growth of key balance sheet components forecast future appreciation of the US dollar at weekly, monthly and quarterly horizons, and across a wide range of currencies.

Additional corroboration for our favored hypothesis comes from a comparison of the predictions of our favored hypothesis with an asset pricing model of exchange rates. In particular, we examine the betas obtained from a simple OLS regression of excess FX returns on balance sheet growth, and compare them with the predictions that arise from an asset pricing model of FX returns. We show indeed that the betas from the OLS regression line up closely with the covariances between excess returns on individual currencies with the FX market excess return.

In this sense, our favored explanation for our findings is in the spirit of the asset pricing approach to foreign exchange markets of Fama (1984), Hodrick (1989) and Dumas and Solnick (1995) who approached the problem of foreign exchange movements in terms of compensation for risk. Our twist is that liquidity conditions add an additional element to the analysis. Balance sheet constraints and the consequent risk appetite of market participants in the foreign exchange market fluctuate in line with funding conditions. A similar logic is shown to hold in the commodities market by Etula (2008), who shows that fluctuations in U.S. broker-dealer balance sheets forecast commodity returns at quarterly horizons; broker-dealer risk appetite is reflected in the risk premia that speculators require for providing insurance to producers and end-users of commodities in the futures market.

The fluctuations in leverage resulting from shifts in funding conditions are closely associated with epochs of financial booms and busts. Figure 1.1 plots the leverage of US primary dealers—banks that have a daily trading relationship with the Fed. They consist of US bank holding companies with large broker subsidiaries (such as Citigroup, JP Morgan Chase, Bank of America), the (former) investment banks (Goldman Sachs and Morgan Stanley), as well as foreign banks with large US broker dealers (such as Deutsche Bank, UBS, Credit Suisse, among others).[2] Each of the peaks in leverage is associated with the onset of a financial crisis (the peaks are 1987Q2, 1998Q3, 2007Q4). Financial crises tend to be preceeded by marked increases of leverage.[3]

As a foretaste of our main empirical results, Figure 1.2 plots the dollar-yen exchange rate[4] together with a summary measure of liquidity conditions given by the lagged quarterly growth of the ratio of broker dealer total assets to household

[2] The current and historical list of primary dealers can be found at `http://www.newyorkfed.org/markets/pridealers_listing.html`.

[3] The plot uses balance sheet leverage of only U.S. primary dealers, as differences in accounting rules and regulation for the foreign dealers lead to vastly different leverage numbers.

[4] It is a convention in the FX market that "dollar-yen" refers to the number of yen that can be bought with one dollar (i.e. the yen/dollar ratio).

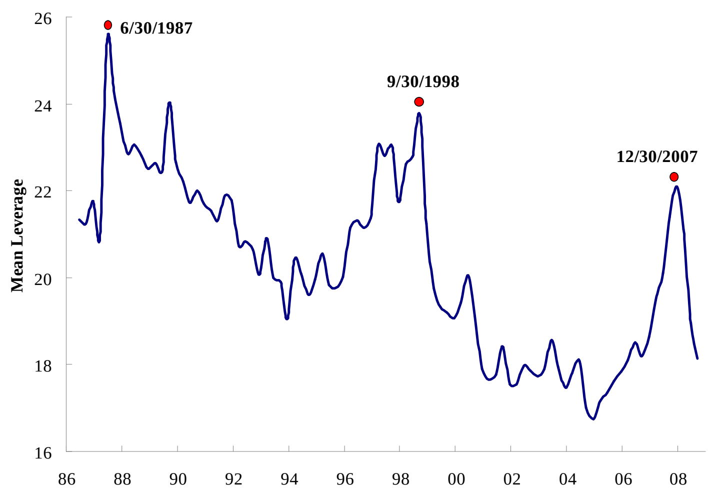

Figure 1.1: Mean leverage of U.S. primary dealers, 6/1986-9/2008 (Source: SEC 10-K and 10-Q filings)

total assets in the US. We see some suggestions already from this chart that liquidity conditions and the value of the dollar against the yen have tracked each other closely. In particular, comparing Figure 1.2 with Figure 1.1 is instructive. The balance sheet growth of US financial intermediaries appears to reflect both leverage conditions and the future value of the US dollar. In what follows, we will show that these initial suggestions have firmer analytical and empirical substance.

Our approach is notable in that it uses only US balance sheet variables. We are able to explain the US dollar's movements against a wide cross-section of currencies by reference only to US financial conditions. Thus, our results can be seen as further confirmation of the central importance of the US capital market in the global financial system.

However, we also acknowledge the limitations of an empirical analysis that focuses just on US financial conditions. The limitations will become important when exchange rate movements are due to shifts in risk appetite of non-US finan-

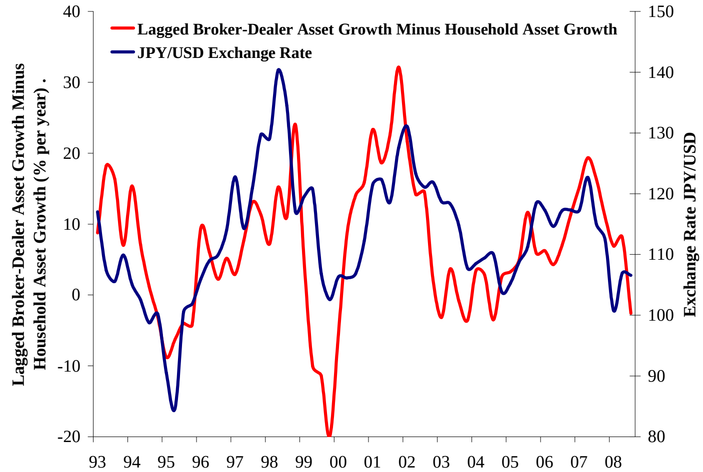

Figure 1.2: Lagged balance sheet growth and the yen/dollar exchange rate, Q1/1993-Q3/2008

cial entities. For instance, if non-US banks have large dollar liabilities, then a global liquidity crisis will lead to a dollar appreciation even though US financial intermediaries may also be slowing their balance sheet growth. The global liquidity crisis in the second half of 2008 following the Lehman Brothers collapse has such a flavor. The issue here is that US intermediary balance sheets are contracting, but they may be contracting *less* than foreign balance sheets. Thus, in relative terms, dollar balance sheets may be increasing, and hence is consistent with an appreciation of the US dollar. In any case, events during a global liquidity crisis may not be easily captured by a standard asset pricing model, and so we urge caution in interpreting our results.

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 liquidity channel explored in our paper differs from the standard logic underpinning the carry trade literature.

Having established the forecasting power of liquidity variables, we then focus on providing a possible rationalization of the role of liquidity variables in terms of balance sheet constraints and the fluctuations in 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. Much still remains to be done in reconciling the strong empirical empirical findings with a coherent theoretical framework, but we believe that our analysis provides some basic building blocks in this direction.

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. 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 incorporating liquidity constraints, as proxied by the growth of financial intermediary balance sheets. We show that the balance sheet variables have robust forecasting power for the bilateral movements of the US 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.

Our approach is notable in that it uses only U.S. variables. We are able to explain the US dollar's movements against a wide cross-section of currencies by reference only to US financial conditions. Thus, our results can be seen as further confirmation of the central importance of the US capital market in the global financial system.

2.1. Data

The empirical analysis that follows uses weekly, monthly, and quarterly data on the nomimal exchange rates of 23 countries over 1993-2007. The countries include nine advanced countries (Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, UK) and 14 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.

In the computation of the pricing kernel or as controls in the panel specification, 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.

Our monthly and weekly liquidity variables are constructed from the outstanding stocks of repurchase agreements and financial commercial paper of the Federal Reserve's primary dealers. The primary dealers are a group of designated banks 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. The Federal Reserve publishes the previous week's repo and commercial

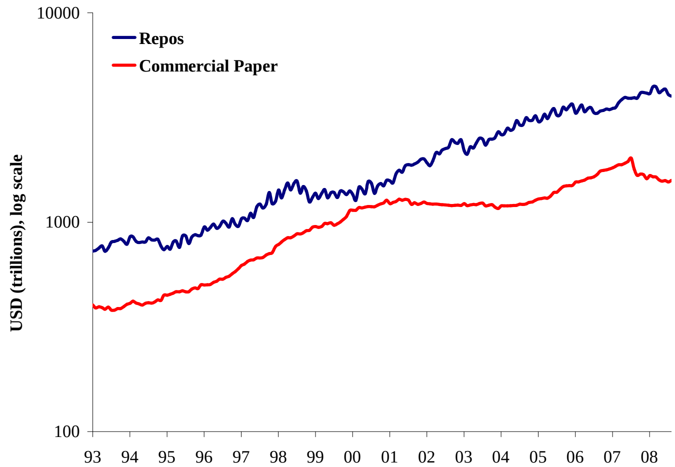

Figure 2.1: Primary dealer repos and financial commercial paper outstanding, 1/1993-8/2008

paper contracts on its website every Wednesday – we incorporate this one-week announcement lag in all of our regression specifications to ensure real-time implementability. 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 annual growth rates of the two variables. The time series suggest that periods of low repo growth tend to be associated with high commercial paper growth, and vice versa. Indeed, the monthly correlation between the annual growth rates of repos and commercial paper is -0.17 over 1/1993 - 8/2008.

In addition to the monthly and weekly forecasts, we also consider predictability of exchange rate returns at quarterly horizons. In the quarterly forecasts, we include a third liquidity proxy, which is computed from the total financial assets of U.S security brokers and dealers and the total financial assets of U.S. households.

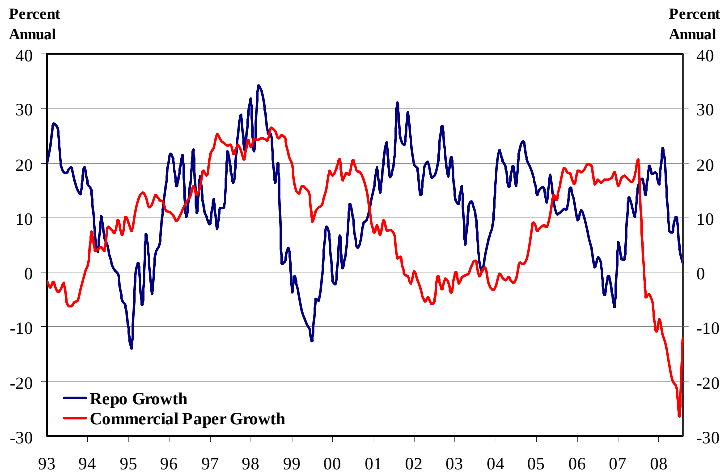

Figure 2.2: Annual growth rates of US primary dealer repos (mean = 12.0%, standard deviation = 9.9%) and financial commercial paper outstanding (mean = 10.0%, standard deviation = 9.5%), 1/1993-8/2008

These data are published quarterly as part of the Federal Reserve's Flow of Funds data release.

2.2. In-Sample Regressions

We begin by considering a panel regression with country fixed effects of the monthly change in the log nominal exchange rate against the US dollar of the sample of 23 countries, with the focus on two lagged forecasting liquidity variables – the annual growth rates of U.S. dollar repurchase agreements (repos) and the stock of commercial paper outstanding. The time period under consideration is 1/1993-12/2007. We also include controling variables, such as the level of US short-term interest rate and the interest rate differential between a particular currency against the US dollar.

The regresssion 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 balance sheet variables have explanatory power for future exchange rate changes. Expansions of balance sheets this month tends to be followed by US dollar appreciation next month. The baseline monthly panel specification is displayed in columns (i)-(ii) of Table 1, which demonstrate that both lagged liquidity variables are highly significant forecasters of monthly exchange rate changes 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 commercial paper growth. Economically, annual repo growth of 20% forecasts a roughly 0.5% appreciation in the U.S. dollar over the following month; similarly, annual commercial paper growth of 20% forecasts a 1% appreciation of the dollar over the following month.

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 higher US short-term interest rate explains a future *appreciation* of the US dollar. The interest rate differential is defined as the difference between the foreign (non-U.S.) short-term interest rate against that for the US dollar. For the sample of all countries (Table 1A), the US 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 incentives that rely on high interest rate differentials.

However, when the sample is restricted to the set of 9 advanced countries only, 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. It also ties together nicely with the fact that high US dollar interest rates are associated with appreciations of the US dollar. We regard the negative sign on the sample of 9 advanced countries as being more credible, due to greater scope of market prices to adjust to the external environment in the absence of explicit policies to peg the exchange rate, or more implicit policies of currency management.

Finally, we conduct two OLS regressions for each country: one with only a constant and another including an autoregressive term. The results are report in Table 3. The autoregressive specification shows that at least one of the two balance sheet variables is statistically significant at 10% level for 16 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. liquidity this month forecasts U.S. dollar appreciations over the next month.

2.3. Out-of-Sample 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 misspecification 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 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 proposed 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.[5] 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 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 liquidity model outperforms both benchmarks at 5% significance level for 11 out of 23 countries.

Among the sample of advanced countries, the largest improvements in forecasts due to the inclusion of liquidity variables is for Australia, Japan and New Zealand, with a smaller effect for the Canadian dollar and Swedish krona. This list is notable for the fact that it includes both the funding currency for the carry trade (the Japanese yen) as well as the destination currencies for the carry trade (Australian and New Zealand dollars). The fact that liquidity variables enter with the same sign in all three cases suggests that the forecasting power of the liquidity variables derive from a different source from the more familiar carry trade incentives. Among the emerging countries, Chile, Columbia and Turkey see the most significant improvements in forecasting power.

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

3. Toward a Theoretical Framework

Having established our benchmark empirical findings, we now turn our attention to how our empirical results can be rationalized.

It is illuminating to begin by taking the cue from the fact that our empirical results differs from the "carry trades" explanation for currency movements. In our panel regression for 23 countries, the coefficient of the interest rate differential is positive, and hence is at variance with the carry trades explanation, which emphasizes the attractiveness of high interest rate currencies. However, if one runs the same regression for a single "carry trade currency" such as the New Zealand dollar or Australian dollar, the coefficient is negative and significant at 5% level. This is consistent with the previous literature: the currencies of *developed* carry countries tend to appreciate rather than depreciate, in violation of uncovered interest parity. Thus, our approach is based on a very different rationale from the carry trades literature.

Liquidity conditions provide a possible explanation for why the US dollar strengthens with falls in US interest rates. 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). As balance sheets expand and leverage rises, the constraints faced by financial intermediaries loosen, thereby increasing their risk appetite. To an outside observer, it would be as if the preferences of market participants were changing toward greater willingness to take on risk. To the extent that foreign currencies are regarded as risky assets by US investors, high US financial intermediary risk appetite relative to foreign risk appetite should be associated with low equilibrium expected returns on these assets. That is, increases in US risk appetite should forecast US dollar appreciations.

We now proceed to work out an equilibrium asset pricing framework in order to investigate the liquidity hypothesis more systematically. We begin with a small illustrative example, which shows how balance sheet constraints lead to fluctuations in risk appetite.

3.1. Balance Sheet Constraints and Asset Prices

Consider a leveraged institution such as a security broker-dealer. The dealer finances the holding of a risky asset (security 1) by holding a short positions in another risky asset (security 2). Let y_1 be the dollar value of security 1 and y_2 is the dollar value of security 2, where $y_2 < 0$. Dealers hold cash of c, and the rate of return on cash is r^f. Then, the balance sheet can be depicted as:

Assets	Liabilities
y_1	$-y_2$
c	w

where w is the equity of the leveraged institution. The balance sheet identity is $y_1 + y_2 + c = w$. Suppose that dealers are risk neutral and aim to maximize returns on their portfolios 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).

We assume that world asset prices depend on a vector of state variables x. If we denote by $r(x)$ the return on the dealer's portfolio, and $\sigma_r(x)$ is the standard deviation of r, the investment problem is:

$$
\begin{aligned}
J(x_t, w_t) \quad &= \quad \max_{y_t} E_t \int_t^{\infty} e^{-\rho s} r(x_s)\, ds \\
&\quad\ \text{subject to} \\
(1) \quad &: \quad VaR_t(r(x_t)) \leq w_t \ \forall t \\
(2) \quad &: \quad dw_t = y_t \left(dP/P - r^f dt\right) + \left(r^f w_t - r(x_t)\right) dt
\end{aligned}
$$

where $VaR(r(x_t))$ is the Value-at-Risk of the portfolio, which we suppose is some multiple α of $\sigma_r(x)$. Due to the risk neutrality, the VaR constraint binds. We assume that returns evolve according to:

$$dP/P = \mu(x_t)\,dt + \sigma(x_t)\,dZ \tag{3.1}$$

where $\mu(x_t)$ is the conditional mean of asset returns, and $\sigma(x_t)$ the conditional volatility. Both depend on the economy's state variables. Because the risk managment constraint is binding, it can be transformed to $r(x_t)^2 \leq \left(\frac{w_t}{\alpha}\right)^2$, so the Hamilton-Jacobi-Bellman equation is:

$$0 = \max_{y_t} e^{-\rho t} r(x_t) - \Phi_t \left(r(x_t)^2 - \left(\frac{w_t}{\alpha}\right)^2 \right) + E_t[dJ]/dt \tag{3.2}$$

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

$$J(x_t, w_t) = e^{-\rho t + f(x_t)} w_t \tag{3.3}$$

which implies:

$$E_t[dJ]/J = -\rho + f_x E_t[dx_t] + E_t[dw/w] + f_x' \left\langle dx \frac{dw}{w} \right\rangle + \frac{1}{2}\left(f_x'' + (f_x')^2\right) E_t \langle dx \rangle^2 \tag{3.4}$$

The first order conditions for portfolio choice are:

$$E_t\left[\frac{dP}{P}\right] + f_x' \left\langle dx \frac{dP}{P} \right\rangle = 2\frac{\Phi_t}{J_t} Y \sigma_P \sigma_P' \tag{3.5}$$

We can then define $\tilde{\Phi}_t = 2\frac{\Phi_t}{J_t}$, so that portfolio choice is:

$$y_t = \frac{1}{\tilde{\Phi}_t} \left(\mu_P + f_x' \sigma_x \sigma_P'\right) \left(\sigma_P \sigma_P'\right)^{-1} \tag{3.6}$$

The Lagrange multiplier associated with the transformed constraint can be solved from:

$$\begin{aligned}
\sigma_r^2 &= y_t' \left(\sigma_P \sigma_P'\right) y_t \\
&= \frac{1}{\tilde{\Phi}_t^2} \left(\mu_P + f_x' \sigma_x \sigma_P'\right)' \left(\sigma_P \sigma_P'\right)^{-1} \left(\mu_P + f_x' \sigma_x \sigma_P'\right) \\
&= (w_t/\alpha)^2 \qquad \text{by the constraint}
\end{aligned}$$

and the Lagrange multiplier is $\tilde{\Phi}_t = \frac{\alpha}{w_t}\sqrt{\left(\mu_P + f_x'\sigma_x\sigma_P'\right)'\left(\sigma_P\sigma_P'\right)^{-1}\left(\mu_P + f_x'\sigma_x\sigma_P'\right)}$.

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 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 market conditions. In other words, the hedge fund's risk appetite fluctuates with shifts in the Lagrange multiplier $\tilde{\Phi}_t$. As the balance sheet constraint binds harder, leverage must be reduced. Figure 1.1 seen earlier should be understood in terms of such fluctuations in risk appetite.[6]

Since "as if" preferences are changing with market conditions, we would expect market prices to be affected by such changing conditions. Our liquidity variables are those associated with fluctuations in balance sheet size—such as primary dealer repos and financial commercial paper. Our approach is to write down an otherwise standard asset pricing model, but where the pricing kernel incorporates explicitly such balance sheet effects.

3.2. The Equilibrium Pricing Kernel

Denote the vector of weights of each asset on the broker-dealer balance sheet by π. Given the asset demands of equation (3.6), market clearing implies:

$$\begin{aligned} \mu_P &= \tilde{\Phi}_t\left(\sigma_P\sigma_P'\right)\pi - f_x'\sigma_x\sigma_P' \\ &= \tilde{\Phi}_t\left(\sigma_P\sigma_E'\right) - f_x'\sigma_x\sigma_P' \\ &= \tilde{\Phi}_t Cov_t\left(\frac{dP^i}{P^i}, \frac{dP^E}{P^E}\right) - f_x' Cov_t\left(\frac{dP^i}{P^i}, dx\right) \end{aligned} \tag{3.7}$$

where $\frac{dP^E}{P^E}$ corresponds to the intermediary's portfolio. So the expected return on each asset is proportional to the Lagrange multiplier of the balance sheet

[6] Danielsson, Shin and Zigrand (2008) solve for the rational expectations equilibrium of a continuous time dynamic model along these lines. Adrian and Shin (2008a) provide a microeconomic foundation for the Value-at-Risk constraint.

constraint, and the prices of risk of the state variables of the economy. So we can see that the state variables of the pricing kernel X_t, and prices of risk λ_t, are:

$$X_t = \left\{ \frac{dP^E}{P^E}, x_t \right\} \tag{3.8}$$

$$\lambda_t = \left\{ \tilde{\Phi}_t, f_x' \right\} \tag{3.9}$$

Our empirical analysis will be in discrete time, and so we can go further by specifying a pricing kernel. We assume that the pricing kernel is exponentially affine in the state variables X_t:

$$\frac{M_{t+1}}{M_t} = \exp\left(-r_t^f - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t' v_{t+1} \right) \tag{3.10}$$

$$\Sigma_t \lambda_t = \lambda_0 + \lambda_1 X_t \tag{3.11}$$

where

$$X_{t+1} = \mu + \phi X_t + \Sigma_t v_{t+1} \tag{3.12}$$

In general, the volatility of the state variables Σ_t can be stochastic, so we assume:

$$vec\left(\Sigma_t \Sigma_t'\right) = S_0 + S_1 X_t \tag{3.13}$$

We further assume that $v_{t+1} \sim N\left(0, I_k\right)$.

4. Pricing Liquidity Risk

4.1. Asset Pricing Approach

Consider investing in foreign bonds of country i with gross holding period interest rate R_t^i, financed by borrowing at the $U.S.$ interest rate R_t. The only risk in this strategy is the movement of the foreign exchange rate $\varepsilon_{t+1}^i / \varepsilon_t^i$. The payoff to the strategy is:

$$R_t^i \cdot \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i} - R_t \tag{4.1}$$

Under the risk neutral measure, the payoff to this strategy is zero. Denoting the pricing kernel M_{t+1}/M_t, the expected payoff is:

$$E_t\left[\frac{M_{t+1}}{M_t}\left(R_t^i \cdot \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i} - R_t\right)\right] = 0 \tag{4.2}$$

Expected exchange rate changes equal relative interest rates, plus a risk premium. Using the definition of covariance, we find the uncovered interest rate parity:

$$\frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i} = \frac{R_t}{R_t^i} + \mu_t + u_{t+1}^i \tag{4.3}$$

where $E_t\left[u_{t+1}^i\right] = 0$, i.e. u_{t+1}^i is exchange rate risk, and

$$\mu_t = -Cov_t\left[\frac{M_{t+1}/M_t}{E_t\left[M_{t+1}/M_t\right]}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}\right] \tag{4.4}$$

is the risk premium. So exchange rate appreciation is due to three components:

$$\underbrace{\frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}}_{\substack{\text{Exchange Rate}\\\text{Appreciation}}} = \underbrace{\frac{R_t}{R_t^i}}_{\substack{\text{Interest Rate}\\\text{Carry}}} + \underbrace{\mu_t}_{\substack{\text{FX Risk}\\\text{Premium}}} + \underbrace{u_{t+1}^i}_{\substack{\text{FX}\\\text{Risk}}}$$

4.2. Estimating Prices of Risk

We assume that $u_{t+1}^i \sim N(0,1)$ for all i. Using Stein's lemma, the FX risk premium (4.4) becomes

$$\mu_t = -Cov_t\left[\frac{M_{t+1}/M_t}{E_t\left[M_{t+1}/M_t\right]}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}\right] = Cov_t\left[v_{t+1}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}\right]\Sigma_t^{-1}\left(\lambda_0 + \lambda_1 X_t\right). \tag{4.5}$$

So the pricing equation reduces to:

$$\frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i} = \frac{R_t}{R_t^i} + \beta_t^{i\prime}\left(\lambda_0 + \lambda_1 X_t\right) + u_{t+1}^i, \tag{4.6}$$

where $\beta_t^{i\prime} = Cov_t\left[v_{t+1}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}\right]\Sigma_t^{-1}$. Exchange rate risk u_{t+1}^i can further be decomposed into a systematic component $\beta_t^{i\prime} v_{t+1}$, and an idiosyncratic component e_{t+1}^i, so that exchange rates are determined by:

$$\underbrace{\frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}}_{\substack{\text{FX}\\ \text{Appreciation}}} - \underbrace{\frac{R_t}{R_t^i}}_{\text{Carry}} = \underbrace{\beta_t^{i\prime}\left(\lambda_0 + \lambda_1 X_t\right)}_{\substack{\text{FX Risk}\\ \text{Premia}}} + \underbrace{\beta_t^{i\prime} v_{t+1}}_{\substack{\text{Systematic}\\ \text{FX Risk}}} + \underbrace{e_{t+1}^i}_{\substack{\text{Idiosyncratic}\\ \text{FX Risk}}} \tag{4.7}$$

4.3. Pricing FX Liquidity

The cross-sectional model in (4.7) 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. Following (3.8), we include three state variables:

$$X_t = \begin{pmatrix} \text{FX Market Excess Return} \\ \text{Repo Growth} \\ \text{Commercial Paper Growth} \end{pmatrix}$$

where we proxy the FX market excess return by the first principal component extracted from the cross-section of foreign exchange excess returns.

Table 4 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 a significant negative loading on lagged CP growth. This result confirms our earlier intuition that 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 to repo and CP growth 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

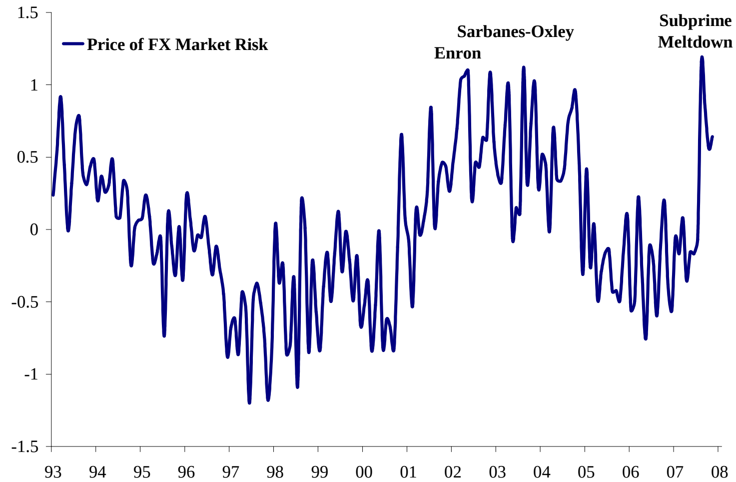

Figure 4.1: Time-variation in the price of FX market risk

the subprime mortgage meltdown in late 2007.

We also investigate the significance of currency-specific factor loadings. Column (i) of Table 5 tests the joint significance of betas for each currency. The bootstrapped p-values in brackets indicate that all currencies except for Colombia have highly 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 10% level for 12 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 Norway, 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 of 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 liquidity correspond to times when balance sheet constraints are relatively loose, enabling 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 the waxing and waning of risk appetite.

In sum, the cross-sectional evidence supports our view that the forecastability of exchange rate returns uncovered in Tables 1-3 is in fact a reflection of systematic changes in risk premia. As U.S. financial intermediary balance sheets expand, U.S. investor risk appetite increases, which decreases the equilibrium returns on all risky dollar-funded positions, including those in foreign currencies. This puts appreciation pressure on the dollar going forward.

4.4. Corroboration from Asset Pricing Model of FX Returns

Our equilibrium asset pricing model in (3.7) rests on the premise that an increase in effective risk aversion $\tilde{\Phi}_t$ forecasts higher expected returns on those positions that covary positively with the relevant benchmark portfolio. We provide further corrobation of our main hypothesis that fluctuations in balance sheet components are associated with innovations in risk appetite.

We present a comparison between two measures of co-movement. On the one hand, we obtain betas from simple OLS regressions of excess FX returns on negative commercial paper growth. We compare these simple OLS betas with the covariance of the excess returns with the FX market excess return. Our favored

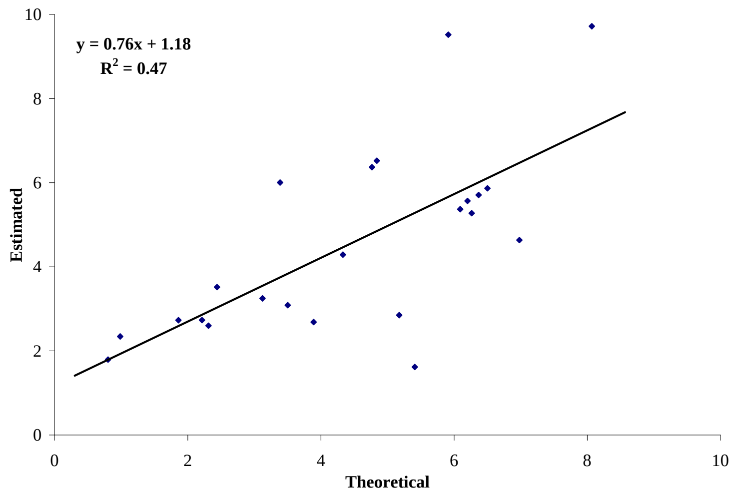

Figure 4.2: All countries. Theoretical vs. estimated coefficients of financial intermediary risk appetite, as proxied by commercial paper growth.

hypothesis that balance sheet changes are associated with shifts in risk appetite (and hence risk premia) would imply that the simple OLS betas should line up with the covariances with the market excess return in the FX market.

Figures 4.2 and 4.3 plot the simple betas obtained from the OLS regressions with the covariance of the excess returns with the FX market excess return. Both scatter plots lend substantial support to our favored hypothesis. The empirical coefficient estimates line up cleanly with the model-predicted coefficients both for our sample of 23 countries as well as for the 9 advanced countries. For the advanced countries, the slope of the relationship is remarkable 0.96 and for the whole sample the slope is still close to unity at 0.76. The R^2-values of both regressions are close to 50%.

We see these findings as lending support to the main theme of our paper that liquidity is priced in the FX market, and liquidity operates through shifts in risk premia.

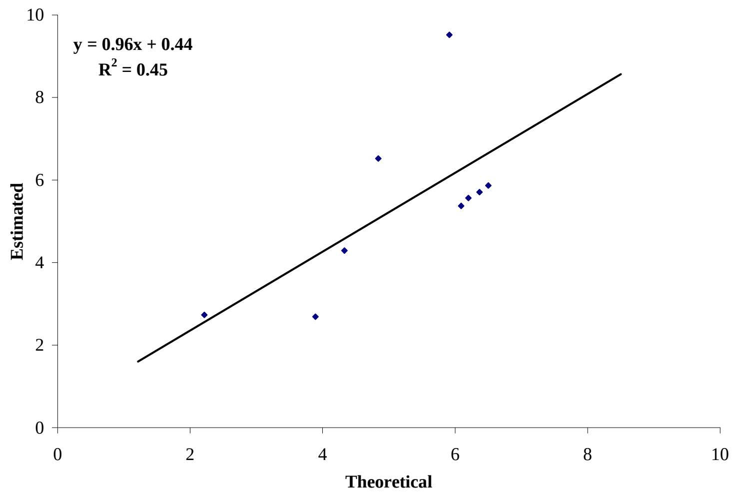

Figure 4.3: Advanced countries. Theoretical vs. estimated coefficients of financial intermediary risk appetite, as proxied by commercial paper growth.

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 found strong evidence that growth in the balance sheets of financial intermediaries have a role in explaining future exchange rate movements. Growth in US dollar components of financial intermediary balance sheets explain future appreciations of the US dollar, both in sample and out of sample. The results hold over horizons as short as one week for a wide range of cross rates. We have shown how this result goes beyond the usual "carry trade" story, in favor of 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 relative liquidity conditions are important in the foreign exchange market is further bolstered by our supporting evidence on other measures of risk appetite such as the VIX index.

Taken together, our two related contributions are the first steps toward an overall framework for thinking about exchange rate movements and how liquidity matters for such movements. The fluctuations in the size of financial intermediary balance sheets is the common thread that ties together exchange rate movements with shifts in risk premia. Thus, the predictable changes in exchange rates may be accompanied by shifting risk premia that are consistent with forward-looking portfolio decisions of investors. 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.

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TABLE 1A (ALL COUNTRIES): Monthly In-Sample Panel with Country Fixed Effects (1/1993-12/2007)

	Dependent Variable: Change in Exchange Rate (%)							
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)
Repo Growth Annual (Lag1)	0.031***	0.029***		0.026***	0.024***	0.029***	0.029***	0.030***
	(3.768)	(3.931)		(3.771)	(3.390)	(3.806)	(3.487)	(3.622)
CP Growth Annual (Lag 1)	0.048***	0.045***		0.046***	0.051***	0.029***	0.029***	0.023***
	(5.594)	(5.668)		(7.268)	(7.737)	(4.286)	(4.603)	(3.421)
Change in Log Exch.Rate (Lag 1)		0.100***	0.078***	0.061***	0.058***	0.056***	0.056***	0.056***
		(4.455)	(5.391)	(3.974)	(3.408)	(3.238)	(3.144)	(3.151)
Interest Rate Differential (Lag 1)			0.054***	0.054***	0.052***	0.052***	0.051***	0.050***
			(23.226)	(26.676)	(16.251)	(17.101)	(14.640)	(14.149)
Stock Mkt. Ret. Dif. Ann. (Lag 1)					-0.001	-0.001	-0.001	-0.001
					(-0.486)	(-0.485)	(-0.465)	(-0.304)
U.S. Interest Rate						0.168***	0.167***	0.197***
						(3.810)	(3.572)	(4.186)
VIX Growth Annual (Lag 1)							-0.001	0.002
							(-0.345)	(1.032)
Signed VIX Growth Ann. (Lag 1)							0.003***	0.001
							(2.852)	(1.053)
TED Growth Annual (Lag 1)								-0.003***
								(-4.590)
Signed TED Growth Ann. (Lag 1)								0.002**
								(2.446)
Constant	-0.706***	-0.664***	-0.052***	-0.913***	-0.972***	-1.508***	-1.493***	-1.519***
	(-3.403)	(-3.609)	(-3.774)	(-5.710)	(-5.865)	(-5.759)	(-5.161)	(-5.298)
# 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
Time-Series Adjusted R^2	2.1%	3.1%	3.7%	5.4%	5.6%	5.8%	5.8%	6.0%

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$; robust t-statistics in parentheses.

TABLE 1B (DEVELOPED COUNTRIES): Monthly In-Sample Panel with Country Fixed Effects (1/1993-12/2007)

	Dependent Variable: Change in Exchange Rate (%)							
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)
Repo Growth Annual (Lag1)	0.023***	0.022***		0.024***	0.025***	0.028***	0.029***	0.031***
	(5.874)	(5.401)		(5.214)	(4.990)	(5.612)	(5.447)	(5.909)
CP Growth Annual (Lag 1)	0.044***	0.044***		0.034***	0.030***	0.020**	0.020**	0.013
	(6.067)	(6.116)		(3.836)	(3.701)	(2.356)	(2.220)	(1.407)
Change in Log Exch.Rate (Lag 1)		0.013	0.024	0.011	-0.006	-0.007	-0.010	-0.011
		(0.727)	(1.621)	(0.621)	(-0.364)	(-0.385)	(-0.551)	(-0.683)
Interest Rate Differential (Lag 1)			-0.190***	-0.088**	-0.182***	-0.147***	-0.162***	-0.185***
			(-6.297)	(-2.165)	(-4.644)	(-3.714)	(-4.466)	(-4.794)
Stock Mkt. Ret. Dif. Ann. (Lag 1)					-0.011***	-0.010***	-0.010***	-0.009***
					(-3.934)	(-3.612)	(-3.969)	(-3.404)
U.S. Interest Rate						0.101***	0.102**	0.125***
						(2.776)	(2.281)	(2.737)
VIX Growth Annual (Lag 1)							-0.001	0.001
							(-0.728)	(0.647)
Signed VIX Growth Ann. (Lag 1)							0.002**	0.000
							(2.477)	(0.149)
TED Growth Annual (Lag 1)								-0.003***
								(-6.635)
Signed TED Growth Ann. (Lag 1)								0.003***
								(3.872)
Constant	-0.903***	-0.912***	-0.110***	-0.828***	-0.853***	-1.197***	-1.198***	-1.199***
	(-7.216)	(-7.223)	(-74.012)	(-6.240)	(-6.452)	(-6.936)	(-5.748)	(-5.580)
# 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
Time-Series Adjusted R^2	3.1%	3.2%	1.5%	3.3%	4.5%	4.5%	4.5%	5.2%

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$; robust t-statistics in parentheses.

TABLE 1C: Quarterly and Weekly In-Sample Panels with Country Fixed Effects (1993-2007)

	Quarterly Exch. Rate Growth			Weekly Exch. Rate Growth	
	(i)	(ii)	(iii)	(iv)	(v)
Exch. Rate Growth (Lag 1)		0.114	0.113		0.001
		(1.525)	(1.511)		(0.026)
Repo Growth Ann. (Lag1)	0.065***	0.058***	0.065***	0.006***	0.006***
	(4.044)	(4.320)	(4.443)	(4.280)	(4.759)
CP Growth Ann. (Lag 1)	0.122***	0.112***	0.116***	0.011***	0.011***
	(6.890)	(7.253)	(7.282)	(8.224)	(9.860)
B-D Asset Growth Ann. (Lag 1)			−0.021*		
			(−1.700)		
Constant	−1.759***	−1.672***	−1.635***	−0.132***	−0.132***
	(−4.115)	(−4.700)	(−4.617)	(−4.514)	(−5.100)
# Countries	23	23	23	23	23
# Observations	1,357	1,334	1,334	19,363	19,357
Time-Series Adjusted R^2	4.5%	5.9%	6.0%	0.4%	0.4%

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$; robust t-statistics in parentheses.

TABLE 2: Out-of-Sample Regressions: In-Sample 1/1993-12/1996, Out-of-Sample 1/1997-12/2007

	Random Walk Benchmark			AR(1) Benchmark		
	ΔMSE	ΔMSE-*Adj.*	p-value	ΔMSE	ΔMSE-*Adj.*	p-value
Australia	0.447	0.892***	0.009	0.487	0.843***	0.003
Canada	-0.117	0.369*	0.071	-0.050	0.315*	0.069
Germany	-0.174	0.331	0.115	-0.065	0.303	0.108
Japan	0.305	0.783**	0.021	0.349	0.706**	0.016
New Zealand	0.706	1.116***	0.003	0.796	1.145***	0.001
Norway	-0.143	0.355	0.137	-0.018	0.348	0.117
Sweden	-0.018	0.473*	0.059	0.092	0.458**	0.046
Switzerland	-0.206	0.260	0.203	-0.132	0.234	0.201
UK	-0.259	0.167	0.227	-0.147	0.221	0.144
Chile	0.172	0.699***	0.005	0.215	0.578***	0.006
Colombia	0.093	1.485***	0.004	0.041	0.404*	0.084
Czech Republic	-0.176	0.372	0.210	-0.009	0.367	0.176
Hungary	-0.701	1.137**	0.026	0.098	0.467**	0.048
India	-0.189	0.609**	0.020	-0.070	0.295**	0.044
Indonesia	2.737	7.482	0.148	4.431	4.843**	0.043
Korea	-0.324	0.528	0.364	0.263	0.638	0.190
Philippines	0.113	0.961	0.124	0.156	0.511	0.164
Poland	-1.003	0.485	0.216	-0.166	0.212	0.247
Singapore	-0.092	0.326	0.197	0.046	0.410	0.122
South Africa	0.499	1.633**	0.030	0.398	0.756**	0.044
Taiwan	0.043	0.588*	0.059	0.091	0.451*	0.065
Thailand	-0.445	0.404	0.394	0.097	0.477	0.250
Turkey	-0.585	21.663***	0.000	-0.264	0.104	0.391
# In-Sample Obs.	50	50	50	50	50	50
# Out-of-Sample Obs.	130	130	130	130	130	130

Adj. = Clark-West (2006) adjustment.

p-values are calculated from a one-sided test.

TABLE 3: Monthly In-Sample OLS Regressions (1/1993-12/2007)

Dependent Variable	Independent Variables										
Exchange Rate Growth	Repo Growth Annual (Lag 1)		CP Growth Annual (Lag 1)		Growth in Exch. Rate (Lag 1)		Constant		Obs.	R^2
Australia	0.024	(1.319)	0.052***	(2.838)			-0.993**	(-2.542)	186	4.60%
	0.025	(1.358)	0.052***	(2.759)	-0.017	(-0.235)	-1.000**	(-2.494)	185	4.53%
Canada	0.017	(1.396)	0.012	(0.929)			-0.457*	(-1.706)	186	1.33%
	0.017	(1.400)	0.011	(0.870)	0.008	(0.113)	-0.446*	(-1.645)	185	1.34%
Germany	0.017	(1.056)	0.035**	(2.164)			-0.702**	(-1.989)	186	2.76%
	0.017	(1.043)	0.034**	(2.060)	0.095	(1.291)	-0.695**	(-1.978)	185	3.90%
Japan	0.037*	(1.863)	0.032	(1.549)			-0.876**	(-1.998)	186	2.70%
	0.038*	(1.889)	0.028	(1.345)	0.040	(0.545)	-0.812*	(-1.844)	185	2.87%
New Zealand	0.037*	(1.898)	0.064***	(3.269)			-1.312***	(-3.105)	186	6.46%
	0.038*	(1.938)	0.067***	(3.281)	-0.043	(-0.585)	-1.356***	(-3.133)	185	6.60%
Norway	0.014	(0.795)	0.037**	(2.082)			-0.682*	(-1.783)	186	2.43%
	0.013	(0.727)	0.040**	(2.215)	-0.021	(-0.288)	-0.720*	(-1.878)	185	2.72%
Sweden	0.016	(0.852)	0.037**	(1.985)			-0.663	(-1.635)	186	2.26%
	0.014	(0.749)	0.039**	(2.096)	0.038	(0.518)	-0.691*	(-1.722)	185	2.82%
Switzerland	0.015	(0.809)	0.038**	(1.989)			-0.736*	(-1.802)	186	2.25%
	0.014	(0.765)	0.038**	(1.961)	0.044	(0.599)	-0.740*	(-1.805)	185	2.64%
UK	0.007	(0.522)	0.013	(0.962)			-0.362	(-1.220)	186	0.58%
	0.006	(0.416)	0.017	(1.244)	-0.090	(-1.242)	-0.419	(-1.437)	185	1.61%
Chile	0.024	(1.502)	0.014	(0.884)			-0.312	(-0.896)	186	1.44%
	0.024	(1.520)	0.015	(0.957)	0.153**	(2.082)	-0.363	(-1.046)	185	3.75%
Colombia	-0.006	(-0.302)	0.043**	(2.277)			0.159	(0.396)	186	2.99%
	-0.005	(-0.252)	0.039**	(2.100)	0.156**	(2.115)	0.085	(0.212)	185	5.43%
Czech Republic	0.012	(0.587)	0.055***	(2.680)			-1.007**	(-2.280)	186	3.79%
	0.010	(0.514)	0.058***	(2.765)	-0.041	(-0.555)	-1.041**	(-2.339)	185	4.06%
Hungary	0.002	(0.104)	0.059***	(3.111)			-0.266	(-0.650)	186	5.10%
	0.001	(0.031)	0.063***	(3.233)	-0.035	(-0.468)	-0.298	(-0.728)	185	5.60%
India	0.039**	(2.059)	0.071***	(3.692)			-1.281***	(-3.081)	186	7.98%
	0.031*	(1.713)	0.052***	(2.796)	0.386***	(5.217)	-0.998**	(-2.533)	185	20.02%
Indonesia	0.192***	(2.667)	0.192***	(2.629)			-3.410**	(-2.167)	186	6.20%
	0.180**	(2.409)	0.183**	(2.442)	0.053	(0.708)	-3.201**	(-1.989)	185	6.46%
Korea	0.040	(1.357)	0.030	(0.992)			-0.653	(-1.006)	186	1.33%
	0.037	(1.252)	0.026	(0.851)	0.117	(1.588)	-0.596	(-0.916)	185	2.68%
Philippines	0.027	(1.465)	0.025	(1.345)			-0.296	(-0.737)	186	1.83%
	0.022	(1.170)	0.021	(1.136)	0.116	(1.549)	-0.224	(-0.555)	185	3.11%
Poland	0.012	(0.628)	0.032	(1.624)			-0.304	(-0.725)	186	1.50%
	0.012	(0.624)	0.031	(1.584)	0.101	(1.376)	-0.337	(-0.807)	185	2.66%
Singapore	0.020*	(1.945)	0.028***	(2.791)			-0.649***	(-2.962)	186	5.23%
	0.019*	(1.846)	0.028***	(2.661)	0.023	(0.314)	-0.636***	(-2.804)	185	5.31%
South Africa	0.036	(1.352)	0.049*	(1.795)			-0.454	(-0.779)	186	2.34%
	0.035	(1.294)	0.049*	(1.789)	0.011	(0.149)	-0.461	(-0.784)	185	2.39%
Taiwan	0.017*	(1.702)	0.024**	(2.429)			-0.365*	(-1.724)	186	4.03%
	0.012	(1.249)	0.021**	(2.112)	0.146**	(1.970)	-0.299	(-1.399)	185	6.15%
Thailand	0.026	(1.105)	0.040	(1.639)			-0.561	(-1.080)	186	1.83%
	0.018	(0.763)	0.032	(1.318)	0.188**	(2.564)	-0.405	(-0.783)	185	5.27%
Turkey	0.074*	(1.757)	0.068	(1.587)			1.118	(1.219)	186	2.58%
	0.067*	(1.649)	0.055	(1.308)	0.254***	(3.556)	0.598	(0.661)	185	8.95%

TABLE 4: Cross-Sectional Prices of Risk

Residual	λ_0	λ_1^{FX}	λ_1^{Repo}	λ_1^{CP}	$\lambda_0 = \lambda_1^{FX} = \lambda_1^{Repo} = \lambda_1^{CP} = 0$
FX Market Excess Return	0.293***	0.110***	0.005	−0.032***	[0.000]***
	(3.15)	(5.49)	(1.06)	(−6.01)	
Repo Growth	−4.870	−2.691**	0.447*	0.065	[0.105]
	(−0.87)	(−2.48)	(1.75)	(0.23)	
CP Growth	0.991	−0.562	0.043	−0.042	[0.459]
	(0.58)	(1.20)	(0.59)	(−0.51)	

Note: bootstrapped t-statistics in parentheses, p-values in brackets; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

TABLE 5: Significance of β's, $\beta'\lambda$'s and Excess Predictability

Test Asset	$\beta^{FX} = \beta^{Repo} = \beta^{CP} = 0$	$\beta'\lambda_0 = \beta'\lambda_1^{FX} = \beta'\lambda_1^{Repo} = \beta'\lambda_1^{CP} = 0$	Predictability of Forecast Residuals
Australia	[0.000]***	[0.008]***	[0.149]
Canada	[0.000]***	[0.234]	[0.696]
Germany	[0.000]***	[0.000]***	[0.621]
Japan	[0.000]***	[0.059]*	[0.233]
New Zealand	[0.000]***	[0.004]***	[0.140]
Norway	[0.000]***	[0.000]***	[0.071]*
Sweden	[0.000]***	[0.000]***	[0.669]
Switzerland	[0.000]***	[0.000]***	[0.561]
UK	[0.000]***	[0.000]***	[0.272]
Chile	[0.000]***	[0.123]	[0.433]
Colombia	[0.272]	[0.690]	[0.550]
Czech Republic	[0.000]***	[0.000]***	[0.336]
Hungary	[0.000]***	[0.000]***	[0.017]**
India	[0.043]**	[0.815]	[0.061]*
Indonesia	[0.002]***	[0.455]	[0.638]
Korea	[0.010]***	[0.719]	[0.612]
Philippines	[0.002]***	[0.108]	[0.711]
Poland	[0.000]***	[0.000]***	[0.117]
Singapore	[0.000]***	[0.002]***	[0.515]
South Africa	[0.000]***	[0.188]	[0.105]
Taiwan	[0.000]***	[0.150]	[0.446]
Thailand	[0.000]***	[0.105]	[0.736]
Turkey	[0.001]***	[0.795]	[0.273]

Note: Bootstrapped p-values in brackets; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.