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Abstract

Theories of financial frictions in international capital markets suggest that financial intermediaries' balance-sheet constraints amplify fundamental shocks. We present empirical evidence for such theories by decomposing the U.S. dollar risk premium into components associated with macroeconomic fundamentals and a component associated with financial intermediary balance sheets. Relative to the benchmark model with only macroeconomic state variables, balance sheets amplify the U.S. dollar risk premium. We discuss applications to financial stability monitoring.

Key words: foreign exchange risk premium, financial stability monitoring, financial intermediaries, asset pricing

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1 Introduction

Theories of financial frictions in international capital markets suggest that shocks to macroeconomic fundamentals are amplified by the presence of funding constraints of financial intermediaries. Such theories of amplification have been proposed by Caballero and Krishnamurthy (2001, 2004) in the context of international financial markets. Brunnermeier and Pedersen (2009) provide a theory based on the "margin spiral," which leads to a spillover of distress across financial market participants. More recently, Bacchetta, Tille, and van Wincoop (2010) and Korinek (2010b) provide additional equilibrium theories of balance sheet amplification. The common thread of this literature on financial amplification is that limited funding liquidity of intermediaries leads to limits of arbitrage, which in turn gives rise to excess movements in asset prices relative to fundamentals.[1] Within an asset pricing context, such excess volatility will generate time variation in effective risk aversion due to changes in the tightness of intermediaries' funding constraints. Balance sheet components thus enter the equilibrium pricing kernel explicitly (e.g. Adrian, Etula, Shin, 2010 and Adrian, Etula, Muir 2010). From a normative point of view, the pricing of intermediary funding constraints gives rise to an externality, as individual financial institutions do not take into account the cost of excessive risk taking for the financial sector as a whole (e.g. Korinek, 2010a).

In this paper, we estimate foreign exchange risk premia associated with both macroeconomic fundamentals and funding liquidity conditions. We start by extracting the common components of expected U.S. dollar funded carry trade returns by applying a partial least squares regression approach to a large number of potential state variables. This produces three common state variables: two are associated with global macroeconomic fundamentals (an inflation state variable and a real state variable), and one is associated with balance sheet components of U.S. financial institutions and the U.S. affiliates of foreign financial institutions.[2] Within the context of a dynamic asset pricing model, we then estimate the price of foreign exchange risk as a function of these estimated state variables. The model allows us to empirically decompose the compensation for systematic foreign exchange risk into components associated with global macroeconomic fundamentals and a component associated with funding liquidity.[3]

Our main finding is that the balance sheet state variable associated with funding liquidity conditions tends to amplify the volatility of the foreign exchange risk premium. Our rationalization for this empirical finding is in terms of the theories of amplification men-

[1] The term "limits of arbitrage" was coined by Shleifer and Vishny (1997) and refers to equilibrium asset price movements that are excessively risky due to financial constraints that arbitrageurs face. Note that a setting with "limits to arbitrage" does not imply that "no arbitrage" fails. In fact, in the setup studied in the paper, we implicitly assume that limits to arbitrage give rise to volatility of the pricing kernel, but that no riskless arbitrage opportunities exist.

[2] Our focus on U.S. financial institutions is due to limited availability of foreign balance sheet data. Hence, our results are expected to underestimate the impact of funding liquidity conditions on the foreign exchange risk premium.

[3] Since the partial least squares methodology allows the balance sheet state variable to be correlated with the macroeconomic state variables, we also investigate the extent to which balance sheets amplify underlying macroeconomic shocks.

tioned earlier—in a frictionless world, we would not expect financial intermediaries' balance sheet components to significantly impact the foreign exchange risk premium. The component of the risk premium associated with the balance sheet variables may also capture some sources of independent shocks emanating from the financial sector, in addition to nonlinear amplification of the macro risk factors by financial institutions. Only a structural general equilibrium model would allow the independent identification of financial sector shocks from the amplification of underlying macroeconomic shocks (Enders, Kollmann, and Müller, 2010, provide an example of such an approach).

Our analysis demonstrates that the excess volatility of the U.S. dollar risk premium associated with balance sheet variables (the "balance sheet risk premium") is tightly linked to three episodes of sharp declines in our indicator. The first decline within our sample started in 1988 (shortly after the signing of the Louvre Accord), and it continued until the beginning of the Gulf War and the 1990 spike in the price of crude oil, which led much of the world into a recession in 1991. The second dramatic compression in the balance sheet risk premium occurred in the run-up to the LTCM crisis, between early 1995 and 1998. The balance sheet premium then reversed sharply in August and October 1998, around the LTCM crisis, and the well documented unwinding of carry trades (see Brunnermeier, Nagel, and Pedersen, 2009 and the references therein). Our analysis indicates that the risk premium associated with macroeconomic fundamentals (the "macro risk premium") played a lesser role in these historical episodes.

The behavior of the foreign exchange risk premium in the early part of our sample stands in contrast with the fluctuations during the global financial crisis of 2007-09, which constitutes our third episode of sharply deteriorating balance sheet capacity. The recent financial crisis featured unusually strong shifts in the components of the foreign exchange risk premium associated with both macroeconomic fundamentals and balance sheets. However, the broad themes of the previous crisis episodes were featured clearly. In particular, the balance sheet risk premium exhibited a prolonged decline between July 2002 and June 2008, while the decline in the risk premium associated with macroeconomic fundamentals was substantially less pronounced. The balance sheet risk premium then increased sharply at the onset of Lehman's bankruptcy. Starting in July 2009 and continuing into early 2010, the balance sheet premium declined rapidly as funding conditions improved. The macro risk premium, however, continued to increase until September 2009. We interpret this lagged response in the macro premium as evidence for a link between financial sector conditions and broader macroeconomic fundamentals. Adrian, Moench and Shin (2010) provide an investigation of this channel for a broad cross-section of financial assets.

The remainder of the paper is organized as follows. In Section 2, we provide a brief overview of the related literature. The method of extraction of state variables via partial least squares is explained in Section 3. We discuss the asset pricing model and the empirical decomposition of the price of foreign exchange risk into components linked to macroeconomic fundamentals and funding liquidity conditions in Section 4. Implications for financial stability monitoring are drawn in Section 5. Finally, Section 6 concludes.

2 Related Literature

Since Fama (1984) we know that uncovered interest rate parity (UIP) is strongly violated for floating currencies. That is, a regression of subsequent relative nominal exchange rate changes on the forward premium typically yields a negative parameter estimate. The most dominant explanation for this phenomenon put forward in the literature is the presence of time-varying risk premia. However, as Engel (1996) notes, many of the existing structural and reduced form models of the foreign exchange rate risk premium are not able to generate estimates of the risk premium that are sufficiently variable to explain the observed deviations from the UIP.

The literature most relevant to this paper has employed asset pricing approaches to analyze the determination of risk premia. Early studies (e.g. Mark, 1985) use a consumption Euler equation. Later studies employ a more flexible approach where a data generating process for the pricing kernel is assumed and estimated. This approach often yields estimates of the foreign exchange risk premium with more realistic dynamics; examples include Groen and Balakrishnan (2006) who use a global conditional factor model, as well as Wolff (1987), Nijman et al. (1993), and Bams et al. (2004) who employ more agnostic time series models based on unobserved component techniques. However, none of these risk premia proxies are able to fully explain away UIP deviations, and if they do, it is based on an implausibly high degree of risk aversion.

Mahieu and Schotman (1994) and Lustig et al. (2010) report substantial success in modeling the pattern of excess currency returns within panels of dollar-based exchange rates by assuming that the UIP deviations are driven by a small number of common components that can be interpreted as risk factors. Our paper follows a similar approach but our aim is not to explain the cross-section of carry returns (i.e., UIP deviations). Instead, the analysis in this paper focuses on explaining the dynamics of the risk premium on a U.S. dollar-funded equal-weighted portfolio of foreign exchange positions. More specifically, we allow the U.S. dollar risk premium to depend on state variables linked to global real activity, inflation and U.S. dollar funding liquidity. Our choice of funding liquidity proxies builds on the study of Adrian, Etula and Shin (2009), who demonstrate that fluctuations in aggregate short-term U.S. dollar liabilities forecast the U.S. dollar exchange rate, and Adrian, Etula, Muir (2010), who study and test a dynamic asset pricing model with financial sector funding constraints in the cross-section of stock returns. Etula (2009) investigates the impact of financial intermediary funding constraints on risk premia in commodity markets.

3 Data and Extraction of State Variables

We use monthly data on exchange rates, global macroeconomic fundamentals and aggregate balance sheet components of U.S. financial institutions. Our focus on U.S. dollar denominated balance sheet components is due to the limited availability of foreign balance sheet data with sufficiently long monthly time series. The sample period runs from January 1988 to March 2010, the beginning of which is dictated by the availability of balance sheet data.

3.1 Measuring the Foreign Exchange Risk Premium

We take the perspective of a U.S. dollar funded investor who measures wealth in U.S. dollars.[4] For simplicity, we suppose that the investor's 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 stems from the movement of the spot exchange rate, ε^i_t, defined as the number of U.S. dollars that can be bought with one unit of foreign currency i.[5] The *excess* return to this strategy is given by:

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

We use monthly data on 1-month spot and forward U.S. dollar-based exchange rates for up to 35 currencies from January 1987 to March 2010. Note, however, that at the start of the sample, we have no more than 13 currencies available, a number that increases to 35 in the second half of the 1990s, and then decreases again to 24 after the introduction of the Euro. At a maximum, we have data on currencies relative to the U.S. dollar for Australia, Austria, Belgium, Canada, Hong Kong, Czech Republic, Denmark, Euro area, Finland, France, Germany, Greece, Hungary, India, Indonesia, Ireland, Italy, Japan, Kuwait, Malaysia, Mexico, Netherlands, New Zealand, Norway, Philippines, Poland, Portugal, Saudi Arabia, Singapore, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Thailand, United Kingdom. The currency data are extracted from Datastream and bid-ask spreads are used to correct the currency data for transaction costs.[6]

As Engel (1996) shows, future currency excess returns generally remain predictable based on current forward premia. We interpret this predictability as compensation for risk, and link the source of predictability to observable macroeconomic and financial intermediary variables. As test assets, we do not use individual currency pairs, but rather portfolios of currencies. We follow Lustig et al. (2010) and form six currency portfolios on the basis of the forward premium at the end of each month assuming that covered interest rate parity holds. That is,

$$\frac{\left(1 + r^{US}_{f,t}\right)}{\left(1 + r^i_{f,t}\right)} = \frac{F^i_t}{\varepsilon^i_t}, \tag{2}$$

where F^i_t is the 1-month ahead forward exchange rate for the U.S. dollar relative to currency i.

More specifically, at the end of each month we sort the available currencies for that month in ascending order based on the current value of the corresponding log forward premium, $\ln\left(F^i_t\right) - \ln\left(\varepsilon^i_t\right)$, for the U.S. dollar *vis-à-vis* currency i.[7] We then allocate these currencies

[4] This implies that risk premia, including the foreign exchange risk premium, are measured from the perspective of a U.S. dollar based investor. Our choice of base currency stems from the superior availability of U.S. balance sheet data, which allows us to measure funding liquidity denominated in U.S. dollars.

[5] That is, an increase in ε^i_t corresponds to an appreciation of the foreign currency relative to the U.S. dollar.

[6] Adrian Verdelhan makes these data available through his website.

[7] The log forward premia approximate the interest rate differentials under (2).

Panel A	Portfolio 1	Portfolio 2	Portfolio 3	Portfolio 4	Portfolio 5	Portfolio 6
Observations	267	267	267	267	267	267
Mean	-0.172	-0.112	-0.064	0.142	0.213	0.314
Std. Dev.	2.068	1.989	2.032	2.028	2.227	2.679
Min	-7.379	-7.586	-5.823	-10.307	-10.008	-12.533
Max	8.930	7.000	6.175	5.076	7.323	8.278

Panel B	Portfolio 1	Portfolio 2	Portfolio 3	Portfolio 4	Portfolio 5	Portfolio 6
Portfolio 1	1					
Portfolio 2	72.8%	1				
Portfolio 3	74.5%	71.5%	1			
Portfolio 4	65.8%	69.6%	71.1%	1		
Portfolio 5	62.5%	66.4%	64.8%	75.3%	1	
Portfolio 6	48.0%	52.6%	53.3%	63.6%	68.0%	1

Table 1: Summary statistics of the returns on six currency portfolios sorted by the interest rate differential relative to the U.S.

to 6 portfolios. Portfolio 1 thus contains the currencies with the smallest log forward premia whereas portfolio 6 assembles the currencies with largest forward premia. For each portfolio, we compute the individual currency excess returns (1) using the forward rate via (2) and take the average across currencies to obtain the portfolio excess return. The portfolios are rebalanced at the end of each month. From equation (1), we can see that returns in each portfolio are proportional to the exchange rate $\varepsilon^i_{t+1}/\varepsilon^i_t$, as ε^i_t is defined as the number of U.S. dollars that can be bought with one unit of foreign currency. A larger U.S. dollar depreciation (larger $\varepsilon^i_{t+1}/\varepsilon^i_t$) thus corresponds to a more positive return. Intuitively, for given interest rates, a dollar depreciation makes investments abroad more attractive as the foreign currency can be exchanged into more domestic currency. Although Lustig et al. (2010) are mainly concerned with the cross-section of carry trade returns, their approach is useful in our context as (i) it provides a way to deal with the unbalanced panel nature of our currency data and (ii) it makes it more appropriate to assume constant risk factor loadings when constructing estimates of kernel-based risk premia. We provide summary statistics for the returns of the six carry sorted portfolios in Table 1.

3.2 Macroeconomic Fundamentals

In order to proxy for U.S. and global economic activity, we construct a panel of monthly real activity data and a panel of monthly inflation data across a range of developed and developing countries. These data are extracted from the Haver Analytics database. The sample runs from January 1988 to March 2010.

The real activity panel consists of 41 real activity series. These include industrial production data for the United Kingdom, Denmark, France, Germany, Spain, Austria, Belgium, Italy, Luxembourg, Norway, Ireland, Portugal, Taiwan, Korea,United States, Japan and capacity utilization rate data for Japan and the United States. This panel also contains consumer and business confidence indicators for the Euro area, France, Italy, Netherlands, the European Union, and the United States; and business confidence indicators for the United Kingdom, Austria, Belgium, Denmark, Luxembourg, Finland, Greece, and Portugal. For Spain, the data only include a consumer confidence indicator. We use annual growth rates of industrial production indices in order to make these series stationary. The confidence indicators are already stationary and therefore we can use the levels of these indicators in our analysis. Japanese capacity utilization rates are not stationary and thus we use annual growth rates, but the U.S. capacity utilization level is stationary.

The inflation panel consists of consumer price index (CPI) inflation data for 19 economies: the United States, the United Kingdom, Belgium, Denmark, France, Germany, Italy, Norway, Sweden, Switzerland, Canada, Japan, Finland, Greece, Ireland, Portugal, Spain, Taiwan, and Korea. As is well known from the empirical macroeconomic literature, annual inflation rates often undergo breaks in their mean, mainly due to monetary policy regime shifts; see, e.g., Sensier and Van Dijk (2004) and Groen and Mumtaz (2008). Therefore, we transform the annual inflation data such that they are guaranteed to be stationary. This is done by taking the 12-month difference in annual (year-over-year) CPI inflation rates, making the dynamic properties of the series in the inflation panel comparable to those in the real activity panel as well as those in the balance sheet panel. The data appendix provides more details on the individual real activity and inflation series.

Structural models of exchange rate determination are usually expressed in terms of relative variables, such as the inflation in the U.S. relative to (or in excess of) the inflation in foreign countries. We do not follow this approach as we are conducting the analysis from the perspective of a U.S. investor, which allows us to be less restrictive. Namely, we allow both the U.S. and foreign activity to influence the U.S. dollar exchange rates, but we do not impose the restriction that these influences must be proportional to the difference in a state variable of interest. In other words, the pricing kernel of our empirical approach will pick up such a relative specification if that is what the data indicates.

3.3 Aggregate Balance Sheet Components

In order to capture time-variation in U.S. dollar financial intermediary conditions, we use four aggregate balance sheet series for which monthly time series are available over a sufficiently long period. These series are plotted in Figure 1. All data are obtained from Haver Analytics.Our first series is the U.S. dollar financial commercial paper outstanding cleared at the Depository Trust and Clearing Corporation (DTCC). The DTCC is a limited purpose trust company chartered in the state of New York. The DTCC clears and settles commercial paper in the U.S. and reports total outstanding commercial paper by types of issuer to the Federal Reserve Board on a weekly basis. The Federal Reserve, in turn, publishes aggregate commercial paper statistics on its website.[8] We employ the dollar-denominated commercial paper issued by U.S. financial institutions and foreign financial institutions with U.S. affiliates. We take year-over-year growth rates of the data to obtain a stationary series. The plot

[8]See `http://www.federalreserve.gov/releases/cp/`.

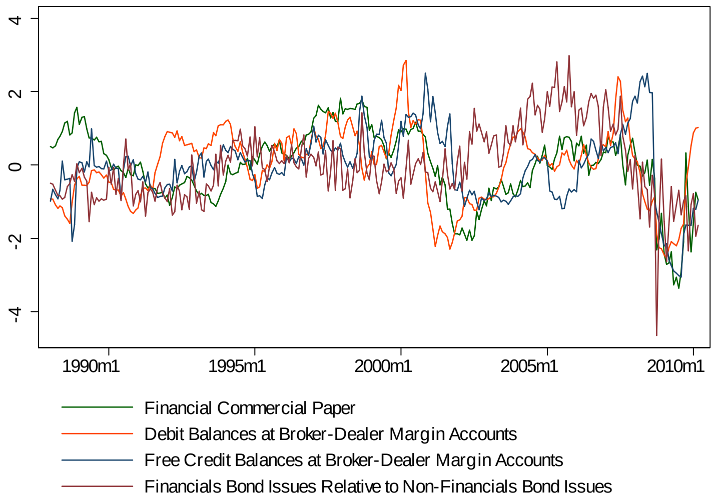

Figure 1: Balance sheet factors. We plot the standardized annual growth rates of U.S. financial commercial paper, free credit balances and debit balances at U.S. broker-dealer margin accounts, and the standardized bond issues of U.S. financial corporations relative to the bond issues of non-financial corporations.

of the standardized series in Figure 1 shows that commercial paper outstanding exhibited its most extreme declines in 1994, 2002 and 2009 with annualized contractions of -1, -2, and -3 standard deviations, respectively. We interpret the financial commercial paper series as a proxy for the short term funding liquidity of financial institutions.

Second, we use data on bond issues of U.S financial corporations and non-financial corporations. The issuance of bonds is reported monthly in the Federal Reserve Bulletin. We take the logarithm of the ratio of financial bonds issued relative to non-financial bonds issued each month. The series exhibits its sample maximum—a 3 standard deviation event—close to the peak of the housing market in 2005, and its minimum—a -4.5 standard deviation event—in the fall of 2008, following the Lehman collapse. We interpret the financial bond issuance series as a medium and longer term funding liquidity indicator of financial institutions. Note that the series only includes bonds with maturities greater one year. Thus, it complements our financial commercial paper series in terms of informational content.

The third and fourth series are the free credit balances and the debit balances at U.S. broker-dealer margin accounts. We again take year-over-year growth rates of the data to obtain stationary series. The sample maximum of free credit balances, a 3 standard deviation event, coincides with the October 1987 stock market crash. However, it is rivaled by the maxima that follow the bursting of the dot-com bubble in October 2000, and the market decline of June 2008. The credit balances bottom in the summer of 2009, following the steep decline and the April bottom of the stock market. The local extrema of debit balances tend to foreshadow the peaks and troughs of the free credit balances by a few months, potentially indicative of market timing by investors. We interpret the free credit and debit balances at

U.S. broker dealers as proxies for the balance sheet capacity of the clients of broker dealers, which are primarily hedge funds. As hedge funds—and particularly macro and emerging markets hedge funds—play an important role in exchange rate determination, we would expect this series to complement the commercial paper and longer term bond issuance series in their informational content.

Taken together, our four balance sheet variables capture the funding liquidity of core financial institutions such as banks, broker-dealers, and institutions of the shadow banking system with the financial commercial paper series (for maturities of less than a year), and with the bond issuance series (for maturities of more than a year). In addition, the free credit and debit balances at broker-dealers capture the funding liquidity of the hedge fund sector.

3.4 State Variable Extraction via Partial Least Squares

We model the U.S. dollar foreign exchange risk premium in a data-rich setting, as a multitude of domestic and foreign factors can potentially affect dollar-based bilateral exchange rates and the risk premia embedded in them. In order to allow for this flexibility in a parsimonious way, we assume that one-month ahead dollar-based risk premia are driven by a common component, which is unobservable but can be estimated from our current data on real activity, inflation, and balance sheet components. This approach generates three state variables that we employ in Section 4 to model the dynamics of the cross-sectional price of systematic foreign exchange risk. We emphasize that our focus is on systematic risk, as understood by a U.S. dollar funded investor, rather than relative or region-specific risk.

Stock and Watson (2002) propose to extract a limited number of principal components from a large panel data set to proxy these common factors. The authors, along with Bai (2003), show that—under appropriate regularity assumptions—principal components can provide consistent estimates of unobserved common factors in large data sets.[9] The drawback of the use of principal components is that it does not always guarantee that the information extracted from a large number of predictors is particularly useful in the context of a modeling exercise. Boivin and Ng (2006) make it clear that if the explanatory power for a certain target variable comes from a certain factor, this factor can be dominated by other factors in a large data set, as the principal components solely provide the best fit for the large data set and not for the target variable of interest. We therefore consider an alternative to principal components analysis in which only factors relevant for modeling the target variables—in our case a panel of dollar-based currency excess returns—are extracted from the set of predictor variables.

An alternative to the principal components approach is the usage of partial least squares (PLS) regressions. As Groen and Kapetanios (2008) show, PLS regressions outperform the

[9] One condition under which principal components provide consistent estimates of the unobserved factor structure is when the factors strongly dominate the dynamics of the data series relative to the non-factor components of the data (see Bai, 2003). However, in an international context, common factors might not dominate the non-structural dynamics because real activity and inflation cycles might not be very strongly synchronized between the U.S. and other economies.

usual principal components-based approach both in simulations and empirically, and especially when the underlying factor structure is weak. Under such circumstances, the accuracy of the factors estimated through principal components will be compromised. The PLS regression approach, on the other hand, will result in consistent estimates of the unobserved common factors relevant for currency returns, even when the factor structure in the combined data on real activity, inflation and balance sheet components is relatively weak (see Groen and Kapetanios, 2008, Theorem 2).

We standardize the $T \times N$ matrix of N indicator variables $Z = (z_1' \cdots z_T')'$ (consisting of real activity, inflation and balance sheet data) so that each variable in Z has zero mean and unit variance, resulting in the $T \times N$ matrix $\tilde{Z} = (\tilde{z}_1' \cdots \tilde{z}_T')'$. We implement PLS regression in a multivariate context by constructing the factors as linear, orthogonal combinations of the standardized predictor variables assembled in matrix $\tilde{Z}$ such that the linear combinations maximize the covariance between the demeaned 1-month ahead dollar-based currency returns (1), and each of the common components constructed from the predictor variables.[10] Specifically, we assume one common component in the dollar-based excess currency returns, and therefore PLS regression is implemented by constructing the dominant eigenvector v of the estimated squared covariance between the vector of demeaned portfolio returns and the panel of combined predictor variables:

$$\tilde{Z}' \bar{er} \bar{er}' \tilde{Z}, \tag{3}$$

where $\bar{er} = (\bar{er}^{1'} \cdots \bar{er}^{6'})'$ with $\bar{er}^i = (\bar{er}_1^i \cdots \bar{er}_T^i)'$ where $\bar{er}_t^i$ is the demeaned excess return of portfolio i. The common factor from $\tilde{Z}$ relevant for the dollar-based excess returns (1) is:

$$X_t = (\bar{v} \tilde{z}_t)', \tag{4}$$

where $\bar{v}$ is a transformation of the $N \times 1$ dominant eigenvector v of (3) such that $||v|| = 1$. This common factor X_t has zero mean and unit variance.

The common factor X_t for the U.S. dollar excess return portfolios is a convolution of developments in global real activity, global inflation, and U.S. balance sheet component data. In order to be able to interpret the movements in X_t and their effect on dollar-based currency returns, we decompose X_t into subfactors relevant for this single common component for the dollar-based excess currency returns: a global real activity subfactor X_t^{real}, a global inflation subfactor X_t^{infl}, and an aggregate U.S. balance sheet subfactor X_t^{BS}. To do this, we impose a hierarchical factor structure. The hierarchical factor structure implies that X_t is a linear combination of the aforementioned real activity, inflation and balance sheet subfactors. Each of the subfactors is extracted as the common component from the corresponding (real activity, inflation or balance sheet) subpanel so as to have the highest covariance with the dollar-based currency returns. We implement this through an iterative procedure where we first use an initial value of the common component in the excess currency returns and apply the PLS on each subpanel relative to this common component to get initial

[10] Demeaning of the dollar-based excess returns is necessary in order to avoid scale effects that can bias the factor estimates.

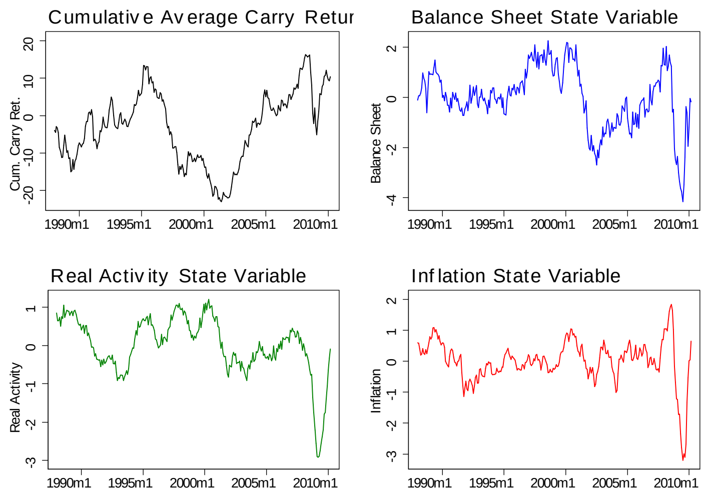

Figure 2: The four panels display the cumulative excess return to the average carry portfolio, the balance sheet state variable, and the two macroeconomic state variables (real activity and inflation).

estimates of X_t^{real}, X_t^{infl} and X_t^{BS}. We then apply the PLS again relative to the panel of excess returns to get an X_t that implies a new estimate of the common component in these excess returns. These steps are iterated until convergence. Groen and Kapetanios (2010) provide additional detail about this procedure.

3.5 Estimated State Variables

Figure 2 plots the evolution in the three PLS-based subfactors (henceforth referred to as state variables) together with the cumulative average excess carry portfolio return.[11] Table A.2 in the data appendix contains information about which of the underlying series is of most importance for each of the three state variables. For instance, Table A.2 reports the R^2 from regressions of each of the (standardized) individual series on either the real activity state variable, the inflation state variable or the balance sheet state variable.Table A.2 indicates that the real activity and the inflation state variables are both dominated by U.S. and Euro area real activity and inflation series. Specifically, the real activity state variable is dominated by perceptions about U.S. consumption and Euro area manufacturing, with the latter being heavily export-orientated. Given this result, the real activity state variable in Figure 2 exhibits a plausible pattern: when real activity is expanding, the expected dollar funded carry returns decrease, presumably because U.S. investors become more inclined to pursue overseas investments. The converse holds for decreases in the real activity state variable.

[11] We first take the cross sectional average of the six currency portfolio returns each month, and then cumulate this average return over time.

Panel A	FX Returns	Cumulative FX Returns	Real Factor	Inflation Factor	BS Factor
Observations	267	267	267	267	267
Mean	0.053	-3.286	0.000	0.000	0.000
Std. Dev.	1.823	9.548	0.762	0.697	1.217
Min	-6.947	-22.929	-2.919	-3.206	-4.160
Max	4.616	16.229	1.206	1.828	2.268

Panel B	FX Returns	Cumulative FX Returns	Real Factor	Inflation Factor	BS Factor
FX Returns	1				
Cumulative FX Returns	11.1%	1			
Real Factor	-12.2%	-24.9%	1		
Inflation Factor	-19.6%	-11.6%	58.7%	1	
BS Factor	-20.3%	-10.7%	74.3%	53.9%	1

Table 2: Summary Statistics. We report the summary statistics of the average realized and cumulative FX returns as well as the macroeconomic and balance sheet state variables. Panel A presents the the mean, standard deviation, minimum, and the maximum. Panel B presents the correlation matrix. All statistics are calculated using monthly data from January 1988 to December 2010.

For example, between 2000 and 2001, the real activity factor turned negative, coinciding with the recession in the U.S. As a result, the dollar appreciated and realized carry returns were low, possibly reflecting the higher risk premia dollar based investors demanded on their foreign investments going forward. When we look at the pattern of the inflation state variable in Figure 2, we observe sharp increases before the 2000-01 recession and particularly before the 2007-09 crisis, which signal heightened global inflation pressures. These peaks are followed by sharp disinflationary movements at the onset of the respective recessions. All told, it appears that U.S. consumption, European manufacturing and both European and U.S. inflation are the most relevant macroeconomic drivers for dollar-based expected currency returns. Notably, such expected returns do not necessarily move in proportion to growth or inflation *differentials*.

Finally, for the balance sheet state variable, we observe a pattern that is similar to that observed in the real activity state variable: stronger real activity tends to be associated with more ample U.S. funding liquidity. We will see in the analysis below that the higher funding liquidity in turn corresponds to an increase in dollar-funded investors' appetite for foreign investments, compressing the compensation for systematic U.S. dollar exchange rate risk. Note, however, that the amplitude of the swings in the aggregate balance sheet factor are larger than those observed for the real activity factor. Specifically, in the period between the 2000-01 recession and the 2007-09 crisis, the balance sheet variable exhibits a sharp upward trend, indicating that persistently more lavish funding conditions were associated with the compression of the U.S. dollar risk premium.[12] Indeed, the trend in cumulative excess returns in Figure 2 appears to be more strongly related to that in the balance sheet variable between the 2000-01 and 2007-09 events than with the trends in the macroeconomic state variables. This suggests that changes in funding conditions amplified the impact of macroeconomic developments on currency risk premia. Summary statistics of the state variables and their correlation matrix are provided in Table 2.

4 Estimating the Foreign Exchange Risk Premium

4.1 Asset Pricing Approach

Following the construction of carry portfolios in section 3, we extract the foreign exchange risk premium by considering payoffs to the carry portfolios. Suppose that the foreign portfolio is invested in riskless bonds with rate of return $r^i_{f,t}$, and that U.S. dollar funding is riskless at rate $r^{US}_{f,t}$ (see equation (1)). Under the risk neutral measure, the payoff to this strategy is zero. Denoting the pricing kernel by M_{t+1}/M_t, the expected payoff is:

$$E_t\left[\frac{M_{t+1}}{M_t}\left(\left(1+r^i_{f,t}\right)\frac{\varepsilon^i_{t+1}}{\varepsilon^i_t}-\left(1+r^{US}_{f,t}\right)\right)\right]=0. \tag{5}$$

Using the definition of covariance, we find the foreign exchange risk premium μ_t to be:

$$\mu_t=-Cov_t\left[\frac{M_{t+1}/M_t}{E_t\left[M_{t+1}/M_t\right]},\frac{\varepsilon^i_{t+1}}{\varepsilon^i_t}\right]. \tag{6}$$

U.S. dollar exchange rate depreciation $\varepsilon^i_{t+1}/\varepsilon^i_t$ thus equals the interest rate carry, the FX risk premium μ_t, and exchange rate risk ξ^i_{t+1} (with $E_t\left[\xi^i_{t+1}\right]=0$):

$$\underbrace{\frac{\varepsilon^i_{t+1}}{\varepsilon^i_t}}_{\substack{\text{Exchange Rate}\\ \text{Appreciation}}} = \underbrace{\frac{1+r^{US}_{f,t}}{1+r^i_{f,t}}}_{\substack{\text{Interest Rate}\\ \text{Carry}}} + \underbrace{\mu_t}_{\substack{\text{FX Risk}\\ \text{Premium}}} + \underbrace{\xi^i_{t+1}}_{\substack{\text{FX}\\ \text{Risk}}}. \tag{7}$$

[12] Note that this trend in our balance sheet state variable appears to coincide with the increased turnover in the global foreign exchange market over this period (see Bank for International Settlements Triennieal Central Bank Survey of Foreign Exchange and Derivatives Market Activity, 2007).

4.2 Empirical Implementation

In order to estimate (7) in the data, we assume that the pricing kernel M_{t+1}/M_t is exponentially affine in the state variables X_t:

$$\begin{aligned} \frac{M_{t+1}}{M_t} &= \exp\left(-r_t^f - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t' v_{t+1}\right), && (8)\\ \Sigma_t \lambda_t &= \lambda_0 + \lambda_1 X_t, && (9) \end{aligned}$$

where

$$X_{t+1} = \mu + \phi X_t + v_{t+1}. \tag{10}$$

We further assume that the innovations to state variables are normally distributed with $v_{t+1} \sim N(0, \Sigma_t)$. With this notation, we use Stein's lemma to express the FX risk premium (6) as:

$$\mu_t = -Cov_t\left[\frac{M_{t+1}/M_t}{E_t[M_{t+1}/M_t]}, \frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}\right] = Cov_t\left[v_{t+1}, \frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}\right]\Sigma_t^{-1}(\lambda_0 + \lambda_1 X_t). \tag{11}$$

It follows that the pricing equation reduces to:

$$\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i} = \frac{1 + r_{f,t}^{US}}{1 + r_{f,t}^i} + \beta_t^{i\prime}(\lambda_0 + \lambda_1 X_t) + \xi_{t+1}^i, \tag{12}$$

where $\beta_t^{i\prime} = Cov_t\left[v_{t+1}, \frac{1/\varepsilon_{t+1}^i}{1/\varepsilon_t^i}\right]\Sigma_t^{-1}$. The investor can cover its foreign positions by entering into a forward exchange rate contract, which locks in the return of the investment. Thus using equation (2), we can rewrite the aforementioned pricing equation as:

$$\underbrace{\frac{\varepsilon_{t+1}^i}{\varepsilon_t^i}}_{\substack{\text{FX}\\ \text{Appreciation}}} - \underbrace{\frac{F_t^i}{\varepsilon_t^i}}_{\substack{\text{Carry}\\ \text{Return}}} = \underbrace{\beta_t^{i\prime}(\lambda_0 + \lambda_1 X_t)}_{\substack{\text{FX Risk}\\ \text{Premia}}} + \underbrace{\xi_{t+1}^i}_{\text{FX Risk}}. \tag{13}$$

Equipped with the cross-sectional no-arbitrage model of (13), we next investigate the extent to which the forecasting variables identified in section 3 determine the FX risk premium. We define systematic FX risk as the unforecastable part of the return to an equal-weighted carry portfolio. More formally, we let the vector of forecasting variables be given by the three estimated state variables that result from our PLS factor extraction approach:

$$X_t = \begin{pmatrix} X_t^{\text{real}} \\ X_t^{\text{infl}} \\ X_t^{\text{BS}} \end{pmatrix}, \tag{14}$$

and consider a single risk factor:

$$v_{t+1} = \tilde{r}^{EW}_{t+1},$$

where $\tilde{r}^{EW}_{t+1} = r^{EW}_{t+1} - E_t\left[\tilde{r}^{EW}_{t+1}|1, r^{EW}_t, X_t\right]$ is the unforecastable part of the equal-weighted carry return. We estimate (13) by way of three-step OLS regressions applied to the cross-section of six carry portfolios (see Adrian and Moench, 2008, for details of the estimation methodology). For simplicity, we assume that betas are constant for each portfolio i.

4.3 Empirical Results

We begin by estimating the model ((13)) for the specification where the price of FX risk is allowed to vary with all three state variables specified in ((14)). Table 3 reports the parameter estimates of the model, where in Panel A we can observe that the model provides a good fit for the returns on our six carry trade portfolios. From Panel B, it becomes clear that our three state variables significantly affect currency risk, with economically intuitive signs. Expansions in the balance sheet and the inflation state variables are associated with compression in the FX risk premium.[13] The real activity variable has a positive but insignificant sign, consistent with previous literature documenting that real activity variables do not have forecasting power for exchange rates or currency returns.

Table 3 also reports the estimated *betas*. All of the betas are highly significant, and are close to 1. The portfolios with relatively lower carry (one and two) tend to have lower betas than the portfolios with relatively higher carry (five and six). As for the prices of risk, we estimate statistically significant, negative prices of risk for the inflation and the balance sheet factor. The real factor does not have a significant price of risk. Our estimate of the systematic U.S. dollar risk premium is plotted in Figure 3, along with the realized returns on the single risk factor, which is the equal-weighted carry return r^{EW}_{t+1}. The figure indicates that our cross-sectional no-arbitrage model is picking up the low frequency component of exchange rate returns well. The risk premium rises sharply in late 1989-90, in 2000-01, and again in late 2008, correctly forecasting the ensuing U.S. dollar depreciations.

4.3.1 Decomposition of the Foreign Exchange Risk Premium.

In order to understand the sources of variation in the compensation for FX risk, we decompose the risk premium of Figure 3 into two components. The first component captures the time variation in the risk premium due to the macroeconomic state variables X^{real}_t and X^{infl}_t. We refer to the resulting series as the "macro risk premium." The second component captures the time variation in the risk premium due to the balance sheet state variable X^{BS}_t, which we refer to as the "balance sheet risk premium." The sum of the macro and balance sheet components of the risk premium captures the time variation in the total FX risk premium.Figure 4 plots the macro risk premium along with the total FX risk premium. The wedge between the two series is due to the balance sheet component of the risk premium

[13] The significance of the balance sheet variable is holds for subsamples, particularly when the data sample ends prior to 2008. The inflation variable, on the other hand, is only significant when the 2008-2010 data is included in the estimation.

Panel A	Portfolio 1	Portfolio 2	Portfolio 3	Portfolio 4	Portfolio 5	Portfolio 6
β^i	0.958***	0.923***	0.962***	0.961***	1.037***	1.159***
t-stat	[13.907]	[12.160]	[12.569]	[15.028]	[17.921]	[10.809]
Observations	266	266	266	266	266	266
R^2	0.67	0.70	0.72	0.76	0.75	0.61

Panel B	λ_0	λ_{Infl}	λ_{Real}	λ_{BS}
Coefficient	0.077	-0.486**	0.384	-0.350**
t-stat	[0.624]	[-2.402]	[1.407]	[-2.201]

Table 3: Panel A reports the risk factor coefficients from OLS regressions of the six carry portfolio returns on the risk factor, the inflation, real and balance sheet variables. Panel B reports the prices of risk of the inflation, real and balance sheet variables. Standard errors are adjusted for heteroskedasticity and autocorrelation, and lambda standard errors are computed using a blockbootstrap. *** denotes significance at the 1 percent level, ** denotes significance at the 5 percent level, and * denotes significance at the 10 percent level.

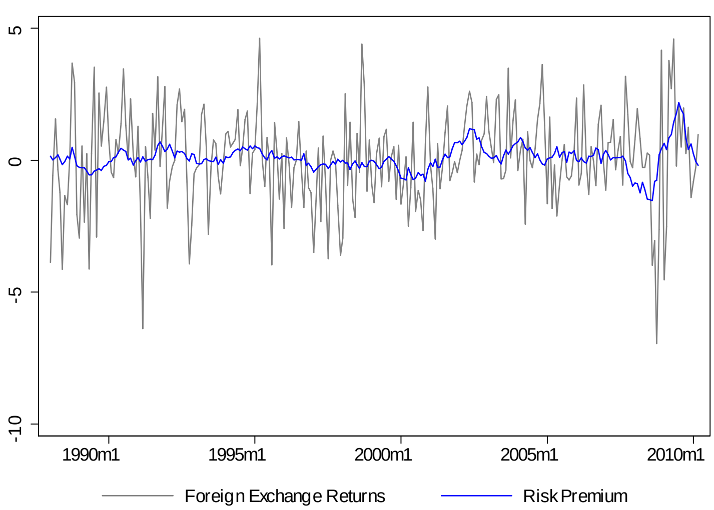

Figure 3: The risk-premium of an equal-weighted U.S. dollar funded carry portfolio and the realized returns on the portfolio.

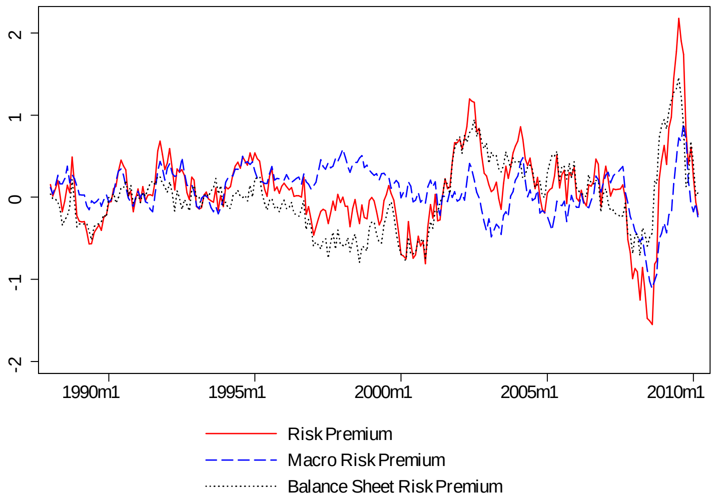

Figure 4: The components of the foreign exchange risk premium associated with macroeconomic and balance sheet variables, and macroeconomic fundamentals alone.

which is also plotted in the figure. Overall, we notice that the total FX risk premium is substantially more volatile than the component attributable to macro variables alone, reflecting a possible mechanism of balance sheet amplification. The wedge between the two series gets particularly wide during times that precede crises and during early parts of crises episodes. We will now walk through these episodes.

At the start of our sample, we observe a substantial decrease in the balance sheet premium in 1989, corresponding to the period before the 1990-91 recession. The wedge between the macro component of the risk premium and the total risk premium may be considered as a warning of the amplification mechanism at work; in particular, the low level of the total risk premium is not fully justifiable by macroeconomic fundamentals, but is in part driven by ample funding liquidity in the economy. Indeed, both components of the risk premium exhibit sharp reversals in the 1990-91 global turmoil. Prior to the Mexican peso crisis of 1994-95 both macro and balance sheet premia again decline but we observe little balance sheet amplification over this period.

Beginning in late 1996, the risk premium associated with balance sheets again dives sharply, driving a large wedge between the total risk premium and the macro risk premium, which persists until the LTCM crisis of 1998. This period is often characterized as that of "irrational exuberance," borrowing the words of former Federal Reserve Chairman Alan Greenspan. Indeed, as Figure 4 shows, hardly any of the decline in the total risk premium can be substantiated by macroeconomic fundamentals.

Following the LTCM collapse, both macro and balance sheet risk premia increased sharply, such that in 1999, all of the FX risk premium is attributable to macroeconomic fundamentals. However, in late 1999, the component associated with balance sheets decreases again rapidly, fuelling the race to the peak of the dot-com bubble in mid-2000. Note

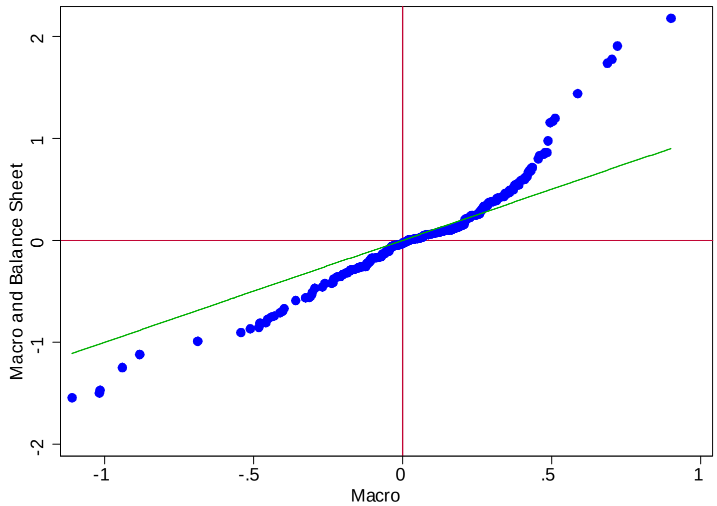

Figure 5: The scatter is a Q–Q plot of the FX risk premium associated with both macro and balance sheet variables (y-axis) against the FX risk premium associated with only macro variables (x-axis). The pattern illustrates the amplification of the risk premium when balance sheet variables are included in the estimation of the pricing kernel.

that, at this point, the total FX risk premium is as low as -1.5% per month, while the macro premium is only -0.5%. The risk premium reverses sharply as the corporate scandals of 2001-02 hit America. The reversal receives strong amplification from the balance sheet component of the risk premium, which increases to its highest level to date in 2002.

Finally, the figure illustrates how the financial crisis of 2007-09 is preceded by a long-lasting decline in the balance sheet risk premium. The long downward trend in the balance sheet risk premium begins in late 2002 and drives a large negative wedge between the total risk premium and the macro risk premium by the middle of 2008. This decline in the balance sheet risk premium is followed by a sharp reversal in the fall of 2008. The macro risk premium follows the dynamics of the balance sheet premium, albeit with a small lag. Another way to understand the mechanism of balance sheet amplification is with a Q-Q plot. We demonstrate this in Figure 5, which plots the total FX risk premium against the component associated with macroeconomic fundamentals in a Q-Q plot. The figure shows that both positive and negative macro risk premia are amplified by balance sheets, resulting in a curved scatter around the 45-degree line.

5 Implications for Financial Stability Monitoring

Systemic risk regulators monitor the evolution of risk in the financial system, develop early warning systems to detect the buildup of potential vulnerabilities, and formulate appropriate policies. This paper presents a methodology to measure the risk premium associated with the dynamics of intermediary balance sheets. The extent to which risk premia are associated with balance sheet expansions are one indicator for the buildup of financial sector risk. Consistent with theories of amplification risk, Figures 4 and 5 demonstrate that financial intermediary balance sheet variables amplify the volatility of the U.S. dollar risk premium. In this section, we explore the implications of balance sheet amplification for financial stability. First, we investigate the association of the FX risk premium with exchange rate volatility. Second, provide analysis of the extent to which the balance sheet risk premium represents an amplification mechanism of underlying macroeconomic fundamentals beyond the *linear* dynamics explored above.

5.1 FX Risk Premium and FX Volatility

Episodes of financial instability are usually accompanied with high FX volatility. In order to gain insight into how the balance sheet risk premium relates to FX volatility, Figure 6 plots the standardized balance sheet risk premium together with standardized log FX volatility. The standardization is done so that each of the variables has mean zero and standard deviation of one. We construct FX volatility from daily exchange rate data, by first computing the standard deviation of log-exchange rate changes within each month, and then taking the cross sectional average across all exchange rates.

Figure 6 shows that the relationship between FX volatility and the balance sheet risk premium is a complex one. There are some episodes in which the volatility measure and the balance sheet risk premium correlate strongly. In particular, the deleveraging in the fall of 2008 was associated with a sharp increase in both FX volatility and the balance sheet risk premium. Log volatility shot up nearly four standard deviations, and the balance sheet risk premium increased by over three standard deviations. The fall of 2008 represented a severe financial crisis where increased FX volatility was associated with an increase in the balance sheet component of the risk premium. However, several periods of sharp increases in FX volatility do not correspond to changes in the balance sheet risk premium; and likewise, several periods of sharp increases in the balance sheet risk premium do not correspond to any changes in FX volatility. For example, in November 1997, FX volatility peaked, corresponding to the Asian currency crisis. Not surprisingly, this spike in volatility was not associated with any particular change in the balance sheet risk premium, as captured by our U.S. dollar balance sheet aggregates. Thus, from the perspective of U.S. financial stability, the Asian currency crisis did not represent a financial sector risk. The converse was true around the 2001 recession. The balance sheet risk premium was at a historical low in spring 2000, just prior to the bursting of the dot-com bubble. Between mid-2000 and the end of 2001, the balance sheet risk premium increased sharply, but this increase was not associated with a change in FX volatility. Figure 3, on the other hand, shows that dollar funded carry returns changed dramatically over the 2000–01 period, which is the development picked up by the balance sheet state variable.

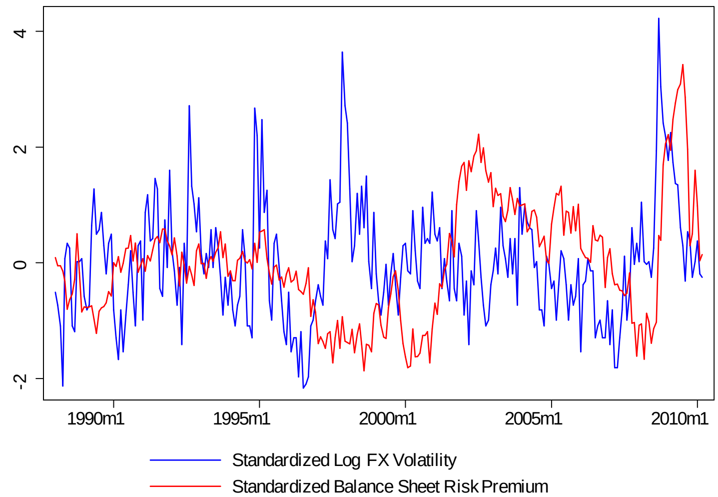

Figure 6: Standardized log FX volatility and the standardized balance sheet risk premium.

5.2 Amplification of Macroeconomic Fundamentals

In the econometric estimation of the FX risk premium, we assumed that the macro risk premium is an affine function of the real activity and the inflation state variables, while the balance sheet risk premium is an affine function of the balance sheet state variable. Our partial least squares methodology, in turn, allowed correlation between the three state variables. In order to gauge the extent to which the balance sheet risk premium captures linear or nonlinear amplification of the macroeconomic variables, we report a regression of the balance sheet risk premium onto the real and inflation variables as well as their squares and cubes. The results are reported in Table 4. Figure 7 plots the fitted value from the regression together with the original balance sheet risk premium. The regression results show that the macroeconomic variables together with their nonlinear transformations explain over 60% of the variation in the balance sheet risk premium. In addition, the plot shows that the explanatory power of the macro variables is particularly good during the crisis of 2008. In earlier times, there were several episodes of balance sheet risk premium variation that were not associated with the nonlinear transformations of macroeconomic fundamentals. Most notable examples include the aftermath of the 2001 recession and the LTCM crisis. The recent financial crisis, however, exhibits a pattern that is fully consistent with theories of balance sheet amplification where limits of arbitrage in the financial intermediary sector serve to magnify underlying macroeconomic shocks.

Inspection of the regression coefficients in Table 4 provides further insight. For the real variable, only its linear term is significant, with a t-statistic of more than 9. For the inflation state variable, only its square and cube are significant. Higher real activity is associated with an expansion of financial intermediary funding liquidity and thus a decrease in the balance sheet risk premium. Greater inflation volatility is also associated with expanding

	BS Risk Premium (%)	
	coef	t-stat
Real State Variable	-1.115***	[-8.899]
Squared Real State Variable	-0.368	[-1.146]
Cubed Real Risk State Variable	-0.356	[-0.992]
Inflation State Variable	-0.135	[-1.022]
Squared Inflation State Variable	-0.751***	[-4.720]
Cubed Inflation State Variable	0.618***	[4.992]
Constant	0.073**	[2.464]
Observations	267	
R^2	61.3%	

Table 4: Results from an OLS regression of the balance sheet risk premium on the real activity variable as well as its square and cube, and on the inflation variable as well as its square and cube. Standard errors are adjusted for heteroskedasticity and autocorrelation. *** denotes significance at the 1 percent level, ** denotes significance at the 5 percent level, and * denotes significance at the 10 percent level.

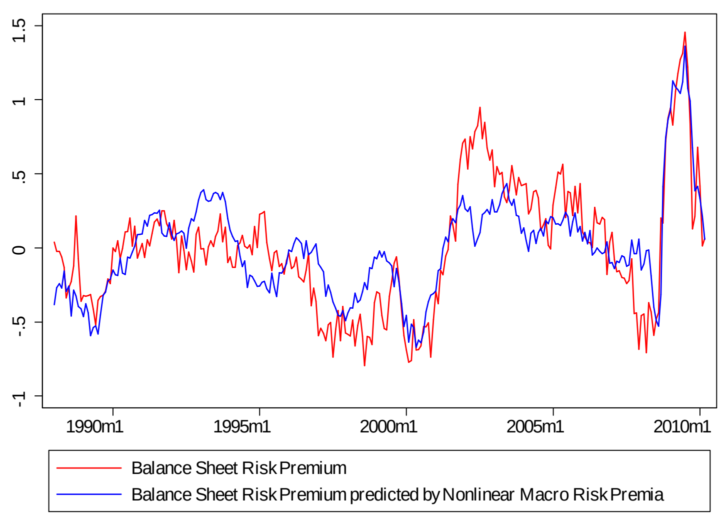

Figure 7: The balance sheet state variable and the fitted values from the regression reported in Table 4.

balance sheets and a compression in the balance sheet risk premium. The cubed term of the inflation state variable is positive, suggesting that more skewness of inflation is associated with a higher balance sheet risk premium.

6 Conclusion

This paper investigates the amplification of the U.S. dollar risk premium by changes in funding liquidity conditions. To do this, we empirically decompose the U.S. dollar risk premium into components associated with macroeconomic fundamentals and a component associated with funding liquidity conditions of U.S. financial institutions. Our results show that funding liquidity conditions have significant explanatory power for the foreign exchange risk premium above and beyond global macroeconomic fundamentals. Moreover, funding liquidity conditions tend to amplify the volatility of the risk premium. We relate these findings to theories of financial frictions in international capital markets, which suggest that shocks to macroeconomic fundamentals are amplified because of financial intermediaries' funding constraints.

The balance sheet component of the risk premium plays a particularly important role during periods of market turmoil. In the run-up to and the unwinding of major financial crises since the late 1980s, the balance sheet component dominates the components associated with global macroeconomic fundamentals. The dynamics of the balance sheet risk premium can be explained by a combination of nonlinear amplification of macroeconomic shocks by financial institutions as well as independent shocks emanating from the financial sector.

In addition to these empirical contributions, our paper develops a new two-step methodology for dynamic decompositions of risk premia. The first step implements a hierarchical partial least squares regression approach to relate the cross-section of expected returns to potential state variables, yielding a desired number of common state variables. The second step estimates cross-sectional prices of risk as a function of the common state variables. While the application here is to one particular asset class—foreign currencies—our approach is more generally applicable.

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A Data Appendix

The data set used for modeling the dollar risk premium consists of 44 domestic and foreign real activity series, 26 domestic and foreign inflation series, and 4 U.S. financial institutions' balance sheet variables; all data are retrieved from Haver Analytics. In order to have $I(0)$ predictor variables, the underlying raw series need to be appropriately transformed; Table A.1 summarizes our potential transformations for the raw series.

Transformation code	Transformation X_t of raw series Y_t
1	$X_t = Y_t$
2	$X_t = \Delta Y_{t,t-12}$
3	$X_t = \Delta Y_{t,t-12} - \Delta Y_{t-12,t-24}$
4	$X_t = \ln Y_t$
5	$X_t = \Delta \ln Y_{t,t-12} - \Delta \ln Y_{t-12,t-24}$
6	$X_t = \Delta \ln Y_{t,t-12}$

Table A.1: Transformation of the predictor variables

Hence, we are using the following series to construct our state variables through PLS regression, which span the sample January 1988 - March 2010:

Table A.2: Potential State Variables used in the PLS Common Factor Model

Index	Definition	Transformation Code	PLS Loading	R^2
Real Activity				
1	Business Confidence: Netherlands	1	1.15	0.77
2	Business Confidence: European Union	1	1.15	0.76
3	Business Confidence: Portugal	1	1.11	0.71
4	Industrial Production: Austria	6	1.10	0.70
5	Industrial Production: Italy	6	1.10	0.70
6	Business Confidence: Luxembourg	1	1.09	0.68
7	Capacity Utilization: United States	1	1.07	0.66
8	Consumer Confidence: United States	1	1.07	0.66
9	Business Confidence: Greece	1	1.07	0.66
10	Industrial Production: Belgium	6	1.05	0.65
11	Industrial Production: France	6	1.03	0.62
12	Business Confidence: Italy	1	1.03	0.61
13	Consumer Confidence: Netherlands	1	1.01	0.60
14	Business Confidence: Germany	1	1.01	0.59
15	Business Confidence: Belgium	1	1.00	0.58
16	Consumer Confidence: Euro Area	1	0.99	0.57
17	Consumer Confidence: European Union	1	0.99	0.57
18	Industrial Production: United States	6	0.98	0.56
19	Business Confidence: Euro Area	1	0.97	0.55
20	Industrial Production: Spain	6	0.97	0.55
21	Business Confidence: United Kingdom	1	0.97	0.55
22	Business Confidence: France	1	0.95	0.52
23	Consumer Confidence: Spain	1	0.93	0.51
24	Consumer Expectations: United States	1	0.91	0.48
25	Business Confidence: Austria	1	0.88	0.45
26	Industrial Production: Germany	6	0.88	0.45
27	Business Confidence: Finland	1	0.86	0.43
28	Industrial Production: Luxembourg	6	0.86	0.43
29	Industrial Production: Japan	6	0.85	0.42
30	Industrial Production: United Kingdom	6	0.85	0.42
31	Consumer Confidence: France	1	0.84	0.41
32	Industrial Production: Portugal	6	0.80	0.37
33	Industrial Production: Ireland	6	0.78	0.35
34	Capacity Utilization: Japan	6	0.76	0.34
35	Industrial Production: Denmark	6	0.67	0.26
36	Business Confidence: Denmark	1	0.61	0.21
37	Consumer Confidence: Italy	1	0.67	0.26
38	Industrial Production: Korea	6	0.44	0.11
39	Industrial Production: Norway	6	0.44	0.11
40	Business Confidence: United States	1	0.44	0.11
41	Industrial Production: Taiwan	6	0.38	0.08
Inflation				
1	Consumer Price Index: Belgium	5	1.21	0.71
2	Consumer Price Index: France	5	1.21	0.71

3	Consumer Price Index: United States	5	1.19	0.69
4	Consumer Price Index: Spain	5	1.16	0.66
5	Consumer Price Index: Finland	5	1.14	0.64
6	Consumer Price Index: Switzerland	5	1.11	0.59
7	Consumer Price Index: Italy	5	1.07	0.55
8	Consumer Price Index: Ireland	5	1.01	0.50
9	Consumer Price Index: Denmark	5	0.94	0.43
10	Retail Price Index: United Kingdom	5	0.94	0.43
11	Consumer Price Index: Japan	5	0.93	0.42
12	Consumer Price Index: Canada	5	0.93	0.41
13	Consumer Price Index: Portugal	5	0.88	0.38
14	Consumer Price Index: Sweden	5	0.86	0.36
15	Consumer Price Index: Germany	5	0.76	0.28
16	Consumer Price Index: Taiwan	5	0.68	0.23
17	Consumer Price Index: Norway	5	0.68	0.22
18	Consumer Price Index: Greece	5	0.59	0.17
19	Consumer Price Index: Korea	5	0.46	0.10
Balance Sheet Conditions U.S.-based Financial Institutions				
1	Commercial Paper Outstanding, Issued by Financial Institutions: United States	6	0.68	0.69
2	Free Credit Balances at Broker-Dealer Margin Accounts: United States	6	0.66	0.64
3	Debit Balances at Broker-Dealer Margin Accounts: United States	6	0.42	0.26
4	Bond Issues by Financial Corporations/Bond Issues by Non-Financial Corporations: United States	4	-0.13	0.03

Note: R^2 results from a regression of the standardized individual series on either the real activity state variable, the inflation state variable or the balance sheet state variable that result from the PLS regression procedure. This R^2 reflects the importance of a series for a state variable. In each category the series are sorted in descending order based on this R^2.

B Robustness to CIP Deviations

Like most models of currency risk premia, our framework assumes that the covered interest parity (CIP) ((2)) holds in the data. Generally this is indeed the case empirically, but there are occasions that large, temporary deviations from CIP ((2)) occur. Failures of CIP can be generated due to the failure of the absence of arbitrage in times of severe financial stress, or due to frictions that prevent arbitrage to work perfectly. For example, policymakers sometimes revert to capital controls or capital flow taxation ('Tobin Tax') to deal with a currency that is under severe depreciating pressure.[14] Aside from explicit foreign exchange interventions, large scale currency movements and associated CIP deviations sometimes coincide with severe financial crises that significantly increase cross-country counterparty risk. Since events like these occur within our sample, we need to check whether our results and conclusions are robust to potential CIP deviations. In order to do that, we estimate a version of our pricing model ((13)) with an extended set of state variables:

$$X_t = \begin{pmatrix} X_t^{\text{real}} & X_t^{\text{infl}} & X_t^{\text{BS}} & X_t^{\text{CIP}} \end{pmatrix}', \tag{15}$$

where X_t^{CIP} is an aggregated measure of CIP deviations.

Intuitively, CIP deviations are related to investors' inability to effectively arbitrage between foreign and domestic bond markets and including X_t^{CIP} in ((13)) allows us to measure how much currency pricing may be distorted when ((2)) fails. Note, however, that the lack of available data on money market interest rates across our cross-section of US dollar-based currency pairs limits the construction of the CIP state variable. As a result, we focus on a smaller cross-section of developed economies for which we have, simultaneously, data on both forward and spot exchange rates and 1-month LIBOR interest rates.[15] More specifically, we construct an unbalanced panel that employs data from 15 currency pairs against the U.S. dollar, including: the Euro area, Australia, Canada, Denmark, France, Germany, Italy, Japan, New Zealand, Netherlands, Portugal, Spain, Sweden, Switzerland and the U.K. In each month we compute for these currencies (when available) their deviations from ((2)) using forward and spot dollar exchange rates as well as the 1-month LIBOR interest rate differentials relative to the U.S. The number of CIP deviations in the resulting panel ranges from 5 in January 1988 to 11 in December 1998 and 9 in March 2010. Given the small number of cross-sectional observations each month, we aggregate the panel by computing the median of the observed CIP deviations within each month. This procedure yields a single time-series, which is our CIP deviation state variable X_t^{CIP}. The resulting estimates are reported in Table B.3. Panel B indicates that deviations from the CIP do not have a statistically significant impact on the U.S. dollar risk premium. Our other pricing results are also qualitatively unaffected by the inclusion of the CIP state variable. Specifically, the magnitude and the statistical significance of the prices of risk of our real activity, inflation and balance sheet state variables do not differ materially from those reported in Table 3.

[14] France and Italy, amongst others, used such measures when their currencies came under pressure within the ERM in the 1980s, early 1990s

[15] These data are obtained from Datastream.

Panel A	Portfolio 1	Portfolio 2	Portfolio 3	Portfolio 4	Portfolio 5	Portfolio 6
β^i	0.964***	0.922***	0.958***	0.960***	1.045***	1.151***
t-stat	[13.925]	[12.043]	[11.582]	[15.425]	[16.797]	[10.902]
Observations	266	266	266	266	266	266
R^2	0.67	0.70	0.72	0.76	0.75	0.61

Panel B	λ_0	λ_{Infl}	λ_{Real}	λ_{BS}	λ_{CIP}
Coefficient	0.161	-0.529**	0.305	-0.318*	-231.974
t-stat	[1.194]	[-2.454]	[1.124]	[-1.869]	[-1.438]

Table B.3: Panel A reports the risk factor coefficients from OLS regressions of the six carry portfolio returns on the risk factor, the inflation, real and balance sheet variables and median CIP deviations. Panel B reports the prices of risk of the inflation, real and balance sheet variables and median CIP deviations. Standard errors are adjusted for heteroskedasticity and autocorrelation, and lambda standard errors are computed from a block bootstrap with a moving block of three periods. *** denotes significance at the 1 percent level, ** denotes significance at the 5 percent level, and * denotes significance at the 10 percent level.

We may also investigate the impact of the CIP state variable on the time-series pattern of the dollar risk premium. This is done in Figure 8, which plots estimates of the dollar risk premium with and without the CIP state variable as well as a decomposition of these premia into the components associated with macroeconomic fundamentals, balance sheets, and the CIP state variable. These plots confirm the message of Table B.3. Broadly speaking, our results are unaffected by CIP deviations, with the exception of some additional spikes in the risk premium in the 1980s-early 1992, 1997-1998 and 2008. These transitory shocks to the risk premium are associated with currency market distortions due to restrictions imposed by policymakers to stave off currency crises (ERM in the 1980s-early 1990s) and severe market frictions during financial crises (LTCM and the Lehman Brothers' collapse).

Figure 8: Do deviations from the covered interest parity affect our results? We plot the total and decomposed dollar risk premium with and without our measure of CIP deviations.