The Non-U.S. Bank Demand for U.S. Dollar Assets

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First Draft: Dec 10, 2019

This Version: July 11, 2021

Abstract

The USD asset share of non-U.S. banks captures the relative demand for USD denominated assets by these investors. An instrumental variable strategy identifies a causal link from the USD asset share to the USD exchange rate. Furthermore, cross-sectional asset pricing tests show that the USD asset share is a highly significant pricing factor for carry trade strategies. The USD asset share also forecasts the movement of foreign currency against U.S. dollar with economically large magnitude, high statistical significance, and large explanatory power, both in sample and out of sample, pointing towards time varying risk premia. It takes 2-5 years for exchange rate risk premia to normalize in response to demand shocks.

Keywords: Exchange Rate Disconnect, Dollar Asset Demand, Intermediary Asset Pricing
JEL Classification Numbers: F3, G1

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I. Introduction

Market commentary often associates exchange rate movements with "changes in positioning" and more generally demand and supply effects. In fact, Froot and Ramadorai (2005) show that institutional investors' order flow helps explain transitory discount rate news of exchange rates. Furthermore, it is well known that exchange rates can deviate substantially and persistently from macroeconomic fundamentals (Rogoff 1996). As a result, an emerging recent literature has focused on documenting that demand and supply effects help explaining exchange rate movements (e.g., Jiang, Krishnamurthy, and Lustig 2018; Engel and Wu 2018; Lilley, Maggiori, Neiman, and Schreger 2019).

In this paper, we present a new measure of relative foreign demand for the U.S. dollar (USD) assets, constructed as the share of non-U.S. banks' USD assets. We construct the relative USD asset demand metric for 26 advanced and emerging economies, and document that this metric is consistently related to exchange rate movements contemporaneously and can help forecast exchange rates both in sample and out of sample. Furthermore, average currency excess returns are systematically related to the cross-section of betas with respect to the USD asset share.

The USD asset share is computed for each economy as the ratio of assets denominated in USD to total assets of the non-U.S. banking system on nationality basis, based on where the ultimate parent of reporting banking-system is headquartered. The USD positions include international positions, that is the USD operations outside the U.S., and U.S.-based branch operations. The non-U.S. banks serve as important marginal investors that price the foreign currency against U.S. dollar.

Time-series regressions show that the non-U.S. bank relative demand for U.S. dollar assets significantly and with an economically large magnitude contributes to the explanation of contemporaneous movements in the exchange rate of foreign currency against USD with a partial R^2 of around 37 percent. The significance and economic magnitude are unchanged by including the U.S. Treasury premium and the home minus U.S. interest differential as controls. Furthermore, the USD asset share significantly correlates with 11 out of 16 bilateral USD exchange rate movements vis-à-vis individual foreign currencies with high R^2s.

The USD asset share is an endogenous variable, and hence we establish a causal relationship using valid demand shifters. The relative demand of USD denominated assets by non-U.S. banks would be expected to be a function of the relative safety and liquidity of USD to substitute assets and the balance sheet capacity of non-U.S. banks to bear risks. One instrumental variable is the Treasury premium of non-U.S. G10 countries, which is proportional to the convenience yield on G10 currencies as substitute safe assets to USD. A second instrumental variable is the sovereign risk of substitute Treasury securities from non-U.S. G10. These two instrumental variables used as demand shifters belong to the group that captures the safety and liquidity of non-U.S. G10 currencies that is the closest substitute for the USD assets. A third instrumental variable used as demand shifter is the shock to the leverage of non-U.S. banks (orthogonal to the leverage of U.S. banks), which relates to the balance sheet capacity of non-U.S. banks to bear risks. The non-U.S. banks play as marginal investors of the foreign currency, the risk bearing capacity of which affects the pricing of foreign currency against USD via its relative demand on USD denominated assets to foreign currency denominated assets.

All three sets of instrumental variables are shown to be valid in a statistical sense, and economic intuition suggests that these are exogeneous demand shifters as they are correlated with USD asset demand, but not with the supply of USD assets. We find that an increase in the Treasury premium of non-U.S G10 countries is significantly associated with decreasing demand for USD assets by non-U.S. banks; higher sovereign risk in Treasury securities of non-U.S. G10 countries is associated with a significant increase in the foreign demand for USD denominated assets; and higher leverage of non-U.S. banks is significantly related to higher demand for USD denominated assets. In addition, an overidentification test based on the Sargan χ^2 statistic confirms that all three instrumental variables are valid exogenous demand shifters. In terms of economic significance, these instrumental variables alone account for more than 20% of the variation in contemporaneous changes of the USD asset demand from non-U.S. banks. Based on the critical values reported in Stock and Yogo (2002), we can reject the null hypothesis that the proposed instrumental variables are weak instruments.

When we investigate the relationship between the USD asset share and the USD exchange rate in the second stage, we continue to find that the higher fitted values of USD asset shares of non-U.S. banks (obtained from first stage) account for a contemporaneous depreciation of foreign currency against USD significantly at the 1% level. Note that the magnitude of the coefficient point estimates on changes in USD asset shares obtained by two-step-least square (2SLS) is around two times as large as those obtained by ordinary least squares (OLS). The F-statistic of the Durbin-Wu-Hausman test of endogeneity rejects that the USD asset share is an exogenous variable. These results suggest that the slope of supply curve of USD denominated assets is underestimated by OLS regressions because it uses equilibrium information of price and quantity. In addition, we further validate our IV strategy by looking into different alternative hypothesis and potential factors that could contaminate the IVs, the robust results further demonstrate the validity of our instrumental variable (IV) strategy. Furthermore, we have more confidence that the relative USD asset demand from non-U.S. banks is mainly driven by safe assets demand by ruling out the alternative hypothesis for speculative reasons (e.g., carry) and demonstrating it is strongly correlated with price measure of safe assets demand (e.g., Treasury premium). This result suggests that our IV strategy identifies the supply curve and corrects the coefficient point estimates on the slope due to the usage of valid exogenous demand shifters.

Perhaps more strikingly, those strong contemporaneous results are complemented with forecasting regressions. When there is a positive innovation to relative demand for USD denominated asset, the foreign currency depreciates against USD contemporaneously, and then forecasts a foreign currency appreciation against USD one, two, three, and five years out. These forecasting results are statistically significant and economically large. The R^2 for the foreign currency exchange rate changes against USD at eight quarters ahead is 40%, with the USD asset share significant at the 1% level. More strikingly, at the twenty-quarter-ahead horizon, the R^2 is 68% and the USD asset share is again significant at the 1% level. All standard errors are adjusted for autocorrelation and heterosceadsticity.

There may be complementary economic channels at play. *First*, there is clearly a "safe asset demand channel" that is also demonstrated by Jiang, Krishnamurthy, and Lustig (2018), and Engel and Wu (2018). Higher USD asset share of non-US banks can be driven by higher demand for USD denominated safe assets, which is associated with a weaker foreign currency against USD contemporaneously. In fact, non-U.S. banks' relative demand for USD is not only driven by relative safety provided by USD denominated assets but also varies with banks' risk bearing capacity. *Second*, our result coincides with the prediction by intermediary asset pricing theory that higher USD denominated asset demand from non-U.S. banks corresponds to lower balance sheet capacity of non-U.S. banks, thus being associated with a significantly negative price of risk in the cross-section of currency excess returns, indicating that these institutions are marginal investors in the foreign exchange (FX) market, especially for corresponding foreign currencies. This finding resonates with Adrian, Etula, and Muir (2014) and He, Kelly, and Manela (2017) who document that intermediary balance sheet capacity is a significant pricing factor for risky assets. *Third*, we show that higher demand for USD denominated assets from non-U.S. banks predicts corresponding foreign currencies appreciation against USD over a longer horizon, but barely predicts the exchange rate dynamics at a very shorter horizon. This suggests time variation in exchange rate risk premia. In particular, it reflects the stationary but persistent deviation from the uncovered interest parity that Duarte and Stockman (2005) and Engel (2014 and 2016) have document. Lilley, Maggiori, Neiman, and Schreger (2019) also find that the U.S. purchase of foreign bonds is associated with risk premia after the GFC, because there is a strong correlation between these flows with traditional risk measures. In summary, because the demand for USD assets driven by risk-bearing capacity of non-U.S. banks is associated with time varying currency risk premia, it helps predict exchange rate dynamics.

We follow the tradition established by Meese and Rogoff (1983) and find these forecasting results also hold out of sample using Diebold-Mariano and Clark-West statistics, which means our forecasting model significantly outperforms the random walk prediction. We document highly significant (i.e., at the 1% level) out-of-sample forecastability for the exchange rate of a basket of corresponding 16 foreign currencies against the U.S. dollar, as well as for the vast majority of bilateral exchange rates. To our knowledge, the strength of out-of-sample forecasting power from a quantity variable is unprecedented in the literature on exchange rates.

As a robustness check, we augment the 9 popular exchange rate models surveyed by Rossi (2013) with the relative USD asset demand from non-U.S. banks in out-of-sample forecasting tests. We find that the out-of-sample forecasting performance in predicting foreign currency exchange rate against USD improves after we incorporate the average relative USD asset demand for all 9 models.

The remainder of the paper is organized as follows. Section II briefly discusses the related literature. Section III describes the construction of the USD demand by non-U.S. banks. Section IV presents the main contemporaneous results, including the Two-Stage Least Squares (2SLS) regressions. Section V shows cross-sectional asset pricing evidence. Section VI gives the forecasting results and robustness checks. Section VII discusses implications for the literature. Section VIII concludes.

II. A Brief Literature Review

We now provide a brief overview of the existing related work. A more detailed discussion of implications for the literature can be found in Section VII. This paper is closely relevant to four threads of literature.

First, there is a large literature about the U.S.' special role as the provider of international reserve currency.[2] A number of papers document that there is a positive and countercyclical safety premium for the U.S. dollar (e.g., Lustig, Roussanov, and Verdelhan 2014; Du, Im, and Schreger 2018; Jiang, Krishnamurthy, and Lustig 2018; Verdelhan 2018). All these papers focus on the U.S. Treasury premium, which captures the scarcity of USD assets based on a single type of security (a 1-year government bond). Our paper first documents the counter-cyclicality in USD asset demand of marginal investors by using a quantity measure, i.e., the USD asset share of non-U.S. banks. Our paper shows that there is a "safe asset demand channel" of non-U.S. financial intermediaries at play – the non-U.S. financial intermediaries pay a safety premium to hold the USD denominated assets, which provides an insurance when negative shocks happen because the U.S. dollar appreciates against corresponding foreign currencies contemporaneously.

[2] Gourinchas, Rey, and Govillot (2010) and Farhi and Maggiori (2018) present models of the special role of the USD in the international financial system.

Second, starting with the seminal contribution of Meese and Rogoff (1983), there is a prevailing view that exchange rates follow a random-walk-like process and hence are not predictable, especially at a short horizon. Several well-known exchange rate puzzles result from this view. First, the “exchange rate disconnect puzzle” shows that the nominal exchange rate is not robustly correlated, even contemporaneously, with macroeconomic fundamentals (e.g., Mark 1995; Cheung, Chinn, and Pascual 2005; Engel and West 2005; Rogoff and Stavrakeva 2008). Second, the “UIP puzzle” implies that on average the interest differential is not offset by a commensurate depreciation of the investment currency (e.g., Fama 1984; Brunnermeier, Nagel, and Pedersen 2009). Our paper documents the demand for USD denominated assets from non-U.S. banks is associated with a time-varying currency risk premium, so it not only helps account for contemporaneous changes in exchange rates, but also helps predict exchange rate dynamics over the longer horizon both in sample and out of sample.

Third, there is a large literature about the “global financial cycle”, which is in fact a “global USD cycle”. For example, Rey (2013) and Miranda-Agrippino and Rey (2015) consider the VIX index and U.S. monetary policy as key drivers of global financial cycles, and Avdjiev, Du, Koch, and Shin (2019) find the broad USD exchange rate is a risk barometer in global capital markets. Our paper contributes to this literature by establishing a causal relationship between relative USD asset demand, the corresponding foreign currency exchange rate against U.S. dollar, the U.S. Treasury premium, and the covered interest parity (CIP) deviation with a novel instrumental variable approach. It relates to a recent literature that explores the failure of CIP for LIBOR (e.g., Ivashina, Scharfstein and Stein 2015; Du, Tepper and Verdelhan 2018) and for Treasuries (e.g., Engle and Wu 2018; Jiang, Krishnamurthy, and Lustig 2018).

Lastly, our paper also sheds light on the theory of intermediary asset pricing. Prominent examples of such theories include Brunnermeier and Pedersen (2009), He and Krishnamurthy (2012, 2013), Brunnermeier and Sannikov (2014), and Adrian and Shin (2014). Adrian, Etula, and Muir (2014) and He, Kelly, and Manela (2017) find that the leverage of U.S. securities broker-dealers possesses significant explanatory power for the cross-sectional variation of expected returns for a wide range of asset classes. This paper focuses on non-U.S. banks and finds that the average relative USD asset demand from non-U.S. banks to be a global risk factor which can help to explain currency excess returns in the cross section.

III. Measuring Relative Demand for USD Assets by Non-U.S. Banks

In this section, we first define our measure of relative demand for USD assets by non-U.S. banks and describe our data sources. We also take a glance at how the measure relates to the exchange rate of corresponding foreign currency against U.S. dollar.

The bilateral exchange rate of any currency vis-à-vis the U.S. dollar is the *relative* value of this currency with respect to the U.S. dollar, i.e., the U.S. dollar is the numéraire. To be comparable, we want to measure the relative USD asset demand within each foreign economy as the *relative* holding denominated in USD to the holding denominated in corresponding foreign currency. Given that non-U.S. banks from each foreign economy are the marginal investors for the foreign currency there, their relative demand for USD should impact the bilateral exchange rate of this specific foreign currency vis-à-vis U.S. dollar. We assume the total assets of non-U.S. banks from economy j are dominated by assets denominated either in currency J used in corresponding economy j or in USD, as the U.S. dollar historically has played a prominent role in global trade and financial flows (e.g., Farhi and Maggiori 2018; Gopinath and Stein

2018). As a result, we define the relative demand for USD assets by non-U.S. banks as the share of USD denominated assets to total assets.[3]

The analysis aims to measure the relative USD demand against corresponding foreign currency of non-U.S. banks, because they are the marginal investors for corresponding foreign currency in foreign economy. We define the non-U.S. banks from the nationality basis, which is determined by where ultimate owners of the banks are headquartered, as opposed to the domicile or residency basis. Due to data limitations on bank-level USD balance sheets, our analysis is performed by aggregating non-U.S. banks' balance sheets at economy level on nationality basis. Relying on USD balance sheets by nationality of reporting non-U.S. banks, a graphical representation of the different aggregates is shown in Figure 1.

[INSERT FIGURE 1 HERE]

International Positions (IP) are the USD operations outside the United States, which capture ① the cross-border positions of non-U.S. banks, and ② the local positions of non-U.S. banks in foreign currency (i.e., USD) outside the United States. The analysis explores the USD-denominated claims from the BIS unpublished free and restricted database of BIS Locational Banking Statistics (LBS) on nationality basis.[4] The Banks' positions are broken down by currency, by sector (bank or non-bank), by country of residency of the counterparty, by country of residence of the reporting banks, and by *nationality* of the ultimate owner of reporting banks.[5] Hence, the USD-denominated aggregates of IP positions can be constructed at the economy level on nationality basis, based on where the ultimate parents of reporting non-U.S. banks are headquartered.

Operations in the U.S. are the USD operations of non-U.S. banks within the U.S., which include both ③ non-U.S. banks' branch operations and ④ non-U.S. banks' subsidiary operations. In the BIS Locational Banking Statistics (LBS), the U.S. does not report local claims/liabilities of non-US banks (and of US banks too). Branch operations within the U.S. are constructed through a bottom-up aggregation based on nationality of balance sheet information available in FFIEC 002 Regulatory Filings. Subsidiary operations in the U.S. are similarly constructed using data from the Call Report Filings (FFIEC 031/041) of non-U.S.-owned subsidiaries, obtained via the S&P Global Market Intelligence platform. The underlying assumption is that the operations of non-U.S. banks in the U.S. are dominated by USD-denominated claims and liabilities.

In this paper, the non-U.S. banks' balance sheet aggregates encompass the International Position (IP) and ③ U.S.-based branches. We don't include U.S.-based subsidiaries, which constitutes the Foreign Position (FP), because non-U.S. banks' U.S.-based subsidiaries are subject to local regulation in the U.S. similar to domestic U.S. legal entities, the funding and liquidity management decisions of which are at least partially (if not entirely) independent of the headquarter.[6] Following the similar approach as USD assets, we

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[3] To avoid double counting, the interoffice claims are excluded from both USD denominated assets (the numerator) and total assets (denominator).

[4] While BIS Consolidated Banking Statistics (CBS) would be better suited, they do not contain the granularity required including the currency composition and the counterparty sectors (Cerutti, Claessens, McGuire 2012). A subset of BIS Locational Statistics on nationality basis is publicly available as well, however, the data with the granularity required for this analysis is only available through the BIS unpublished free and restricted database.

[5] Due to data limitations, the currency breakdowns only include USD, EUR, CHF, JPY, GBP, so it is impossible to know the holding of assets denominated in corresponding currency specifically used at nationality level.

[6] This is the key reason we use nationality basis instead of decile or residential basis of our analysis. However, the conclusion of this paper doesn't vary if we adopt Foreign Position (FP) concept rather than International Position (IP) plus US-based branches to measure USD asset demand. Relevant results can be provided upon request.

construct the total assets of non-US banks (for all currencies) on a nationality basis. However, there are data limitations for local positions in local currencies from BIS Locational Banking Statistics (LBS), which start only from 2012 for most nationalities. The aggregates of local positions in local currencies on nationality basis, which enter the denominator of USD asset share of non-U.S. banks, are obtained based on a bottom-up consolidated approach using FitchConnect and Factset database.[7] In the absence of the currency breakdown at bank-level data, we assume the share of local currency position to all currency position is equal to what is available in the Other Depository Corporations (ODCs) survey.[8]

We want to acknowledge a caveat in our aggregated measure due to data limitations. Ideally, we would like to capture the relative demand for USD assets by non-U.S. banks of each foreign economy from a pure quantity perspective. However, the aggregated non-U.S. banks data at nationality level do not allow us to completely isolate the quantity of assets purchased from the exchange rates and prices at which they are purchased. We try our best to tackle this challenge from two types adjustment throughout the whole paper. First, exchange rate adjusted outstanding claims for all series obtained from BIS Location Banking Statistics (LBS) are used following the same method as BIS recommends.[9] This helps eliminate the impact of methodological changes and more importantly exchange rate movements for those components. As a result, those components are purely isolated from exchange rate movements. Second, the local positions in local currencies are reconstructed with deliberate adjustment. To better isolate the contamination in movements of exchange rate and value, we use the share of local currency position to all currency position obtained from Other Depository Corporations (ODCs) survey of previous period *instead of* contemporaneous period to adjust. As a result, this component is purely isolated from not only movements in exchange rate but also movements in price of underlying securities of the contemporaneous period.

Lilley, Maggiori, Neiman, and Schreger (2019) also acknowledge this limitation of aggregated data and explore pure quantities of U.S. foreign bond flows using the U.S. mutual fund holdings data from Morningstar, which does allow the decomposition of market-value positions into exchange rates, prices, and quantities. However, as far as we know, given the complexity of global banks reporting, the current measure using aggregated data of non-U.S. banks at nationality level with considerate adjustment in exchange rate movement and price movement is the best we could achieve so far. We view the two approaches as complementary. Their finding confirms that the reconnect between the USD exchange rate and U.S. foreign bond flows (denominated in USD) since the GFC is either not due to the mechanical influence of exchange rate or due to the influence of price on the value of U.S. foreign bond purchases. In the end, we'd like to highlight two key distinctions between our paper and Lilley, Maggiori, Neiman, and Schreger (2019): 1) the power of USD asset share of non-U.S. banks in explaining the exchange rate dynamics is not restricted to the post-GFC sample; 2) we emphasize the role of marginal investors played by non-U.S. banks, especially on FX market of foreign currencies rather than focus on the U.S. mutual funds.

Table 1 summarizes the sample of economies used in this paper. The initial sample covers 26 economies, where USD operations are considered to be of systemic importance (based on the IMF 2019).[10] For

[7] In addition, we have further validated the robustness of this approach by using local positions from the other depository corporation surveys as well.

[8] Although the ODCs survey is based on residency basis as opposed to nationality, but this is the best proxy we could think of.

[9] The adjusted change in the outstanding is calculated by first converting US dollar-equivalent amounts outstanding into original currency using end-of-period exchange rates, then calculating the difference in amounts outstanding in the original currency, and finally converting the difference into a US dollar-equivalent change using average period exchange rates.

[10] Throughout this paper, economies are defined based on where the ultimate owners of reporting non-U.S. banks are headquartered. For example, non-U.S. banks of Germany means German banks around the world.

convenience, Germany is used as the representative for eight economies in the Euro Area, since the Euro has been adopted by most of those countries during our sample period (2001q2-2017q4).[11] China and Russia are excluded, because the relevant USD aggregates only span over 2 years. We don't include Hong Kong SAR China for its adoption of the fixed exchange rate regime.[12] The final sample covers 16 legal tender currencies with (largely) floating exchange rate regimes from 16 economies covering 10 advanced and 6 emerging economies.

[INSERT TABLE 1 HERE]

Table 1 also presents the summary statistics of key variables that we are interested in in this paper. Columns (1) to (16) report the summary statistics by foreign economies, and Column (17) reports those of the cross-sectional average across all 16 foreign economies. In the upper panel, we report the summary statistics of key variables in levels. First, the variable s_t stands for log nominal exchange rate per U.S. dollar, an increase of which means that the U.S. dollar appreciates against foreign currency. Second, the variable $D_{\$t}$ stands for the USD asset share of non-U.S. banks from specific foreign economies (nationalities of reporting banks). For example, $D_{\$t}$ in Column (1) stands for USD asset share of the Australian banks around the world, opposed to that of non-U.S. banks reside in Australia, because our measure is constructed on nationality basis. The Swiss banks around the world have the highest USD asset share standing at 27.9 percent on average over the sample period, whereas that of Brazil has the lowest USD asset share at 3.8 percent instead. The average USD asset share across those 16 foreign economies is around 9.5 percent during the sample period. Third, the variable Φ_t stands for the U.S. Treasury premium at the one-year horizon.[13] The U.S. Treasury premium is positive across almost all foreign economies, which suggests that investors are willing to pay a safety premium to hold U.S. Treasury securities relative to foreign government bonds in corresponding foreign economy. The only exception is that the U.S. Treasury has negative premium vis-à-vis Australia at around -7 bps on average.[14] The U.S. Treasury has the largest premium vis-à-vis Brazil around 189 bps on average. The average U.S. Treasury premium, i.e., the U.S. Treasury premium against a basket of 16 foreign government bonds, is around 38 bps over the sample period. The variable $y_t - y_{\$t}$ stands for the government bond yield differential between each foreign economy and the United States. Emerging markets tend to have higher yield differentials against the United States. For example, the government bond yields in Brazil and Turkey are 9-10 percent higher than that of the United States on average. On the other hand, some safe heavens (e.g., Japan and Switzerland) have even lower yields than the United States. The average yield differential between the cross-sectional average of 16 economies and the U.S. is around 2.2 percent over the sample period. In the lower panel, we report the summary statistics of the quarterly changes of the key variables. The longest sample covers the period 2001q1-2017q4 for the levels in the upper panel and, correspondingly, 2001q2 to 2017q4 for the changes in the lower panel.

[11] Eight Euro Area economies include: Austria, Cyprus, France, Germany, Italy, Luxemburg, Netherlands, and Spain. Using Germany as representative doesn't mean non-U.S. banks for other Euro Area economies are not important in terms of USD operations, the conclusion of this paper doesn't vary if we measure aggregates in Euro Area by adding up eight Euro Area economies. Relevant results can be provided upon request.

[12] The classification of regime is based on Annual Report on Exchange Arrangements and Exchange Restrictions 2018 (IMF 2018).

[13] Following Du, Im, and Schreger (2018), $\Phi_{j,t} \triangleq y_{j,t}^{Govt} - \rho_{j,t} - y_{\$,t}^{Govt}$, where $y_{j,t}^{Govt}$ and $y_{\$,t}^{Govt}$ denote 1-year government bond yield of economies *j* and the 1-year U.S. Treasury yield of the United States (all issued in respective currencies), and $\rho_{j,t}$ is the 1-year market-implied forward premium for hedging currency *j* against the U.S. dollar. The data source for government bond yields is from Bloomberg, and spot and forward exchange rates are from WM/Thomson Reuters.

[14] Du, Im, and Schureger (2018) and Cerutti, Obstfeld, and Zhou (2019) point out that there is considerable heterogeneity in the determinants of the U.S. Treasury premium and LIBOR cross-currency basis across economies and time. For example, the heterogeneity in cross-currency funding gap can be one of the sources (IMF 2019). Australia has a negative USD funding gap during the period of investigated in this paper, which is different from other economies in the sample.

[INSERT FIGURE 2 HERE]

Figure 2 shows the time series properties of quarterly changes in the log USD exchange rate against a basket of 16 foreign currencies (Δs_t) and the quarterly changes in the USD asset share ($\Delta D_{\$t}$), which measures the relative demand for USD denominated assets of non-US banks from a quantity basis. The contemporaneous correlation is 60%, much higher than the correlation between the USD and proxies of safe asset demand such as quarterly changes in the U.S. Treasury premium ($\Delta\Phi_t$) which has a 36% correlation or interest differentials ($y_t - y_{\$t}$) which have a -32% correlation. Furthermore, and perhaps more strikingly, the USD asset share highly correlates with future exchange rate changes 3-years ahead (with a correlation of 50%, again much higher than the corresponding correlation of the U.S. Treasury premium of 29% or the interest differential of 7%). These bilateral correlations are presented in Table 2.

[INSERT TABLE 2 HERE]

IV. Exchange Rates and Demand for USD Assets

In this section, we first show that there is a strong contemporaneous relationship between the exchange rate of foreign currency against USD and the relative demand for USD assets from non-U.S. banks in the time series. Second, we give a causal interpretation using a set of novel instrumental variables.

A. Time-series relationship

We examine the relationship between the relative demand for USD denominated assets from non-U.S. banks and the exchange rate over time. We regress quarterly changes in the log nominal USD exchange rate on contemporaneous quarterly changes in the relative demand for USD denominated assets, captured by USD assets as share of total assets from non-U.S. banks of corresponding economy. Our benchmark regression specification is given by:

$$\Delta s_{j,t} = \alpha_1 + \beta_1 \Delta D_{\$,j,t} + \beta_2 \Delta\Phi_{j,t} + \beta_3 \Delta\left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right) + \beta \boldsymbol{X}_{j,t} + \varepsilon_t \tag{1}$$

where the dependent variable $\Delta s_{j,t}$ is the annualized quarterly change in the log nominal exchange rate of the USD vis-à-vis currency J, an increase of which means corresponding foreign currency J depreciates against USD between quarters t-1 and t. The variable $\Delta D_{\$,j,t}$ denotes the quarterly change in the USD asset share of non-U.S. banks from economy j between quarters t-1 and t. The variable $\Delta\Phi_{j,t}$denotes the quarterly change in the U.S. Treasury premium vis-à-vis foreign economy j between quarters t-1 and t. The variable $\Delta\left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right)$ denotes the quarterly change in the interest differential of foreign economy j over the United States between quarters t-1 and t. Finally, $\boldsymbol{X}_{j,t}$ is a vector of control variables.[15]

[INSERT TABLE 3 HERE]

Table 3 shows our regression results. In Columns (1) to (7), we perform regressions by using the cross-sectional average series across 16 foreign economies. It can be interpreted as a relationship between the annualized quarterly changes in the log USD exchange rate against a basket of 16 foreign currencies and

[15] In some regressions, we control $\Phi_{j,t-1}$ and $\left(y_{j,t-1}^{Govt} - y_{\$,t-1}^{Govt}\right)$ as Jiang, Krishnamurthy, and Lustig (2018). In addition, we also control for $D_{\$,j,t-1}$.

the cross-sectional average of demand for USD assets by non-U.S. banks from those 16 economies. The coefficient point estimates on $\Delta\bar{D}_{\$,t}$ are positive and significant at the 1% level with similar magnitudes across all specifications.[16] This is consistent with theory prediction (e.g., Jiang, Krishnamurthy, and Lustig 2018; Engel and Wu 2018), suggesting that an increase in relative USD asset demand against foreign assets from non-U.S. banks is associated with U.S. dollar appreciation against a basket of foreign currencies on average. In terms of magnitude, the coefficient point estimate on $\Delta\bar{D}_{\$,t}$ from the univariate regression in Column (1) implies that a one percent quarterly increase in average USD assets (as share of total assets) of non-U.S. banks across 16 foreign economies is associated with an appreciation of the USD against a basket of 16 foreign currencies by 7.2% in the same quarter.[17] To provide a further sense of magnitudes, a positive shock to the USD asset share by one-standard deviation of 0.32% (from Table 1) is associated with an appreciation in the USD vis-à-vis a basket of 16 foreign currencies by around 2.3% in that quarter on average. In terms of economic significance, the variation in the USD asset demand of non-U.S. banks alone accounts for 37% of the variation in the contemporaneous movements of U.S. dollar vis-à-vis a basket of 16 foreign currencies. Given the well-known exchange rate disconnect puzzle (Frankel and Rose 1995; Froot and Rogoff 1995), this magnitude of R^2 is very meaningful and surprisingly high.

Column (2) of Table 3 includes the contemporaneous quarterly change in the U.S. Treasury premium, which is an alternative measure of USD asset demand advocated by Jiang, Krishnamurthy, and Lustig (2018), and Engel and Wu (2018). Our measure is different from theirs in at least four aspects: 1) we use quantity information instead of price information; 2) we focus on all USD-denominated assets rather than only one specific USD-denominated assets (U.S. Treasury at 1-year maturity); 3) we only focus on marginal investors (i.e., banks) rather than all investors; 4) we focus on more general relative demand for USD denominated assets rather than only on "safe asset demand". In Column (2), the coefficient point estimate for $\Delta\bar{\Phi}_t$ is positive as predicted but it is insignificant and the R^2 only increases marginally to 40%, which implies the additional information from this alternative measure documented in the literature is limited after controlling for the quantity measure of relative USD demand from non-U.S. banks, that is the USD asset share of non-U.S. banks.

Column (3) of Table 3 includes the contemporaneous quarterly change in the interest differential. The result is consistent with simple predictions from textbook models of exchange rate determination – interest rate increases in the U.S. relative to a basket of foreign currencies are associated with USD appreciation against foreign currencies. However, this relationship is not significant.[18] Furthermore, the additional contribution to the R^2 is negligible.

Column (4) of Table 3 includes both covariates of $\Delta\bar{\Phi}_t$ and $\Delta\left(\bar{y}_t^{Govt} - y_{\$,t}^{Govt}\right)$ along with changes in the USD asset share ($\Delta\bar{D}_{\$,t}$). The R^2 rises to 41%, and the coefficient point estimate on $\Delta\bar{\Phi}_t$ becomes marginally significant at the 10% level, and coefficients and significance for other variables are nearly

[16] The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are used for all time-series analyses in this paper. Newey and West (1987) show that their estimator for the covariance matrix is consistent if the lag length (*m*) fulfills the following two conditions: a) the lag length (*m*) grows with the sample size T, and b) the lag length (*m*) grows at a slower rate than $T^{1/4}$. Following Greene (2003), the lag length (*m*) is set to the integer part of $T^{1/4}$, i.e. 2 for our application, to ensure the estimator for the covariance matrix is consistent for heteroskedasticity and autocorrelation. All key message in this paper is robust to use different lag length (*m*) from 1-5, and the results can be provided upon request.

[17] The dependent variable Δs_t is annualized quarterly change, so the point estimate on $\Delta\bar{D}_{\$,t}$ is divided by 4 to get the magnitude for change in that quarter.

[18] Additional analysis splits the sample between advanced countries and emerging markets. The results of the former support the textbook model prediction, violating the uncovered interest parity (UIP) and are significant at the 1% level. For the latter, however, the relationship between interest rate differential and USD exchange rate supports the UIP – interest rate in U.S. increases relative to a basket of currencies is associated with USD deprecation, but statistically insignificant.

unchanged. In Column (5), we control for the one quarter lags of $\bar{D}$, $\bar{\Phi}$, and $(\bar{y}^{Govt} - y_{\$}^{Govt})$. This specification is equivalent to include the AR(1) innovations of the three factors discussed above. The reason we prefer controlling lags directly rather than constructing the AR(1) innovations obtained from regression residuals is because this can help avoid look-ahead bias introduced in the construction of residuals. Given that we don't have the luxury to obtain the residuals from an expanding-window regression approach due to limited sample, this issue needs special attention not to say we are facing the challenge of exchange rate disconnect puzzle. In this more stringent setting, we find that the coefficient point estimates and significance of $\bar{D}$ are nearly unchanged. In addition, we valid the finding of Jiang, Krishnamurthy, Lustig (2018) that the AR (1) innovations of $\Delta\bar{\Phi}_t$ has additional information on contemporaneous movements of exchange rates.[19] The coefficient point estimate on $\Delta\bar{\Phi}_t$ is about doubled to 21.90 and becomes significant at the 1% level and the R^2 rises to 52%. However, the $\Delta(\bar{y}_t^{Govt} - y_{\$,t}^{Govt})$ is still insignificant but with correct sign.

In Columns (6) and (7), we use the same specification as presented in Column (5) but split the whole sample into two halves – the 1st half is from 2001q2 to 2009q4 that includes the GFC episode, and the 2nd half is from 2010q1 to 2017q4 that is post-GFC episode. The coefficient point estimate on $\Delta\bar{D}_{\$,t}$ is about the same level of the whole sample presented in Column (5) and remains significant at the 1% level for both subsamples. However, the coefficient point estimate on $\Delta\bar{\Phi}_t$ is 50% lower in the pre-GFC sample comparing with that of post-GFC sample, and it is only significant at 5% for both subsamples.[20] The coefficient point estimate on $\Delta(\bar{y}_t^{Govt} - y_{\$,t}^{Govt})$ supports the text book prediction, but are insignificant for both episodes as in whole sample. The R^2 is above 60% for both episodes. The result for the pre-crisis episode distinguishes our findings from the recent development in exchange rate disconnect literature, which documents a reconnection between the dollar exchange rate and measures of risk or liquidity has emerged only after Global Financial Crisis in 2007-2009.[21]

In the right panel of Table 3, we adopt the same specification as in Column (5) but focus on the bilateral exchange rate of the USD vis-à-vis different individual foreign currencies in our sample (see, Columns (8) to (23)). We find that the contemporaneous relationship between the USD asset share of non-banks from foreign economy j and the exchange rate of the USD vis-à-vis foreign currency J significantly holds for 11 out of 16 cases.[22] The significant coefficient point estimates on $\Delta D_{\$,j,t}$ vary from 5.9 (CHF) to 65.8 (BRL). The changes in the U.S. Treasury premium also have explanatory power for 10 out of 16 cases. The changes in interest differential have more explanatory power for bilateral exchange rates: 9 out of 16 cases support the textbook prediction. The cross-sectional average results are mostly contaminated by currencies in emerging markets, where they either have contradictive and significant signs (MXN, TRY, ZAR) supporting the UIP condition or have insignificant results (BRL, INR, and MYR).

Taken together, these results show that the USD asset share of non-U.S. banks is a robust explanatory variable for the variation in the exchange rate of foreign currency against USD over time. The explanatory

[19] The results are uniformly weaker if we use changes in 3-month U.S. Treasury premium, which is consistent with findings from Jiang, Krishnamurthy, and Lustig (2018), likely because the 3-month U.S. Treasury premium is a noisy measure of the long-term expectation term that drives exchange rates.

[20] The finding that $\Delta\bar{\Phi}_t$ loses explanatory power (but not much) can be due to our shorter sample (starting from 2001q2) compared to the sample staring in 1988q1 of Jiang, Krishnamurthy, and Lustig (2018).

[21] Lilley, Maggiori, Neitman, and Schreger (2019) find U.S. purchases of foreign bonds, implied volatility of the S&P index (VXO), the return of the stock market (S&P 500), credit spread constructed by Gilchrist and Zakrajšek (2012), the returns of financial intermediaries composed by He, Kelly, and Manela (2017), and the U.S. Treasury premium introduced by Du, Im, and Schreger (2018) only establish economically sizable and statistically significant connection with the USD exchange rate after the Global Financial Crisis in 2007-2009.

[22] Among 5 insignificant cases, the point estimates on $\Delta D_{\$,j,t}$ for INR, NOK, SGD, and ZAR have the positive signs consistent with theory prediction, however, the only exception is GBP flips the sign but insignificant.

power doesn't change before or after Global Financial Crisis. Adding additional drivers does not affect the explanatory power of quantity measure of USD safe asset demand captured by USD asset share, especially in the case of controlling for the price measure of USD safe asset demand (the U.S. Treasury premium).

Our findings suggest that the exchange rate of individual foreign currency vis-à-vis USD of corresponding foreign economy is, to a large extent, driven by a simple story of relative demand for USD assets against individual foreign-currency denominated assets from the perspective of non-U.S. banks. Higher relative demand for the USD denominated assets against individual foreign-currency denominated assets by non-U.S. banks from that foreign economy is driving down the corresponding foreign currency where the respective banks are headquartered, especially because those non-U.S. banks play as one of the most important marginal investors for foreign currency. This finding also coincides with the earlier findings in the literature that order flows as quantity proxies for imbalance demand are significant determinants of exchange rates contemporaneously (e.g., Evan and Lyons 2002).

What is strikingly different from the existing literature is the magnitude of the contemporaneous correlations. The quantity proxy for relative USD asset demand single alone, i.e., the USD asset share of non-U.S. banks, accounts for 37% of the variation in the exchange rate of a basket of 16 foreign currencies against USD over time (in terms of the R^2), which is quantitatively beyond anything that has been documented in the literature so far.

Though striking, Table 3 is just the beginning of our empirical investigation. We next turn to an instrumental variable approach that allows us to establish a causal relationship between the relative USD demand and the exchange rate of foreign currency against USD in subsection B. Then we move to asset pricing in Section V and to forecasting relationships in Section VI.

B. Causal relationship

We now turn to establish the causal interpretation between the relative demand for USD denominated assets from non-U.S. banks and exchange rates using novel instrumental variables. It is a classical challenge in economics to separate supply and demand effects. If the demand and supply curves shift simultaneously over time, the observed data on relative quantities (USD assets as percent of total assets) and relative prices (exchange rates) reflect a set of equilibrium points on both curves. Consequently, an OLS regression of relative prices on relative quantities fails to identify or to trace out either the supply or demand relationship. Angrist and Krueger (2001) provide a comprehensive literature survey on using instrumental variables to identify causal economic relationships, especially to use exogeneous curve shifters to identify demand elasticity or supply elasticity. In spirit of it, we apply this strategy to identify the supply elasticity of USD. Furthermore, we try to justify our IV strategy captures the relative demand for USD denominated assets for its property that serve as global safe assets, and then valid our IV strategy especially in terms of exclusion restriction by eliminating several alternative hypotheses. In the end, we extend this strategy to the U.S. Treasury premium and LIBOR cross-currency basis in the next subsection, which helps us identify different roles played by the "safe asset demand channel" and the "financial intermediation channel", two channels proposed in the literature.

In the previous subsection, the underlying assumption for the OLS regression to obtain a consistent estimate of the supply curve is that the observed changes in quantities and prices are solely due to demand shifts. This might not be a very strong assumption for any given individual foreign economy, because the global supply of USD can be regarded as given for non-U.S. banks from each specific foreign economy and the USD asset share of non-U.S. banks from each specific foreign economy can be regarded as proxy for relative demand changes. However, it might be a strong assumption for average USD asset share of non-U.S. banks from 16 foreign economies as a whole. In this subsection, we try to adopt a conservative approach – find exogeneous demand shifters for USD assets of non-U.S. banks. A valid demand shifter should only affect demand conditions of USD denominated assets exogenously without affecting cost conditions of supplying USD denominated assets. Based on this principle, we propose two groups of instrumental variables:

Safety of substitute currencies. The costs of substitutes are often used as demand shifters in the literature (Angrist and Krueger 2001). One seminal application is Wright (1928), who uses exogeneous curve shifter to estimate the demand and supply elasticities of flaxseed, a source of linseed oil. He argues a valid demand curve shifter can only affect demand conditions without affecting cost or supply conditions. For example, a variable he used for demand curve shifter was the price of substitute goods, such as cottonseed. In this spirit, as the closest substitute for USD denominated assets to serve as global safe assets, we only focus on sovereign securities denominated in six non-USD currencies from non-U.S. G10 economies – that is, Canadian Dollar for Canada; Swiss Franc for Switzerland; Euro for Euro Area economies; Japanese Yen for Japan; Pound Sterling for United Kingdom; and Swedish Krona for Sweden. We consider two different variables that can capture the safety of assets denominated in substitute currencies:

(i) The first measure is the Treasury premium of substitute economies, which is proportional to the convenience yield on substitute safe assets. This is built on the work of Du, Im, and Schreger (2018) that defines the U.S. Treasury premium, which is proportional to the convenience yield on U.S. safe assets of non-U.S. investors. Similarly, the Treasury premium of substitute economy, $\Phi_{j,t}^{sub}$, is equal to $(y_{j,t}^{Govt} - \rho_{j,t} - y_{sub,t}^{Govt})$, where $y_{j,t}^{Govt}$ and $y_{sub,t}^{Govt}$ denote 1-year Treasury yield of economy j and that of substitute economy (all issued in respective currencies), and $\rho_{j,t}$ is the 1-year market-implied forward premium for hedging currency j against the substitute currency.[23] By construction, it captures the premium that investors are willing to receive less from holding the substitute Treasury securities relative to government bonds in foreign economy j.

(ii) The second measure is the sovereign CDS spread of substitute Treasury securities, which capture the sovereign risk of substitute safe assets that are denominated in substitute currencies. In fact, sovereign risk is a key component that influences the Treasury premium (Du, Im, Schreger 2018; Engel and Wu 2018). We obtain the sovereign CDS spreads from Markit for government bonds with 5-year maturity (the most liquid maturity).

These two measures of safety of substitute currencies are simple averages across six non-USD G10 currencies, including Canadian Dollar, Euro, Japanese Yen, Pound Sterling, Swedish Krona, and Swiss Franc. We would expect when the Treasury premium (sovereign CDS spread) of substitutes increases (decreases), the substitute currencies become safer – the demand for USD assets from non-U.S. banks purely driven by "safe asset demand" channel would be substituted more by demand for other safe assets denominated in substitute currencies.

[23] Substitute currency refers to the simple average across six non-USD G10 currencies (Canadian Dollar, Euro, Japanese Yen, Pound Sterling, Swedish Krona, and Swiss Franc). Correspondingly, substitute jurisdiction refers to the simple average of non-U.S. G10 jurisdictions (Canada, Euro Area, Japan, United Kingdom, Sweden, and Switzerland). The Treasury yield of German government bond is used as the Euro Area representative, so as the respective CDS spread of German bonds is used as the Euro Area representative for the second instrumental variable. The key message is robust to use different definition of substitute currency, e.g., simple average across three non-USD G10 currencies (Euro, Japanese Yen, and Swiss Franc). The results of robustness check can be provided upon request.

Balance sheet constraints of non-U.S. banks. The risk-taking capacities of non-U.S. banks should matter for their relative demand for USD denominated assets against local currency denominated assets. We build on the work of Adrian, Etula, Shin (2010) and Adrian, Etula, Muir (2014) to measure the risk-taking capacities of non-U.S. banks as the logarithm of the book leverage ratio. We use the bank balance sheet data from Fitchconnect and match with the bank ownership information from Bankscope to make sure that the leverage ratios that we construct share the same economies as USD asset share of non-U.S. banks where they are headquartered. We orthogonalize this economy-specific bank leverage measure with respect to the leverage of the U.S. banking system, so that the residual is irrelevant to the pricing kernel of U.S. financial intermediaries. We would expect that when the (orthogonalized) leverage ratio of non-U.S. banks increases, balance sheet capacity deteriorates, and the demand for USD assets increases.

Table 4 reports our Two-Stage Least Squares (2SLS) results using the two groups of instrumental variables discussed above. As before, we focus on time-series regressions of the cross-sectional average of 16 economies over the full sample period.

[INSERT TABLE 4 HERE]

In the upper panel of Table 4, we investigate the first stage relationship – between the USD asset demand from non-U.S. banks, captured by USD asset share, and three different instrumental variables proposed above.[24] In particular, we run the following regression:[25]

$$\Delta\overline{D}_{\$,t} = \alpha_1^1 + \beta_1^1 \Delta\overline{\Phi}_t^{sub} + \beta_2^1 \Delta\overline{Sovereign\ CDS}_t^{sub} + \beta_3^1 \Delta\overline{Leverage}_t + \beta^1 \boldsymbol{\overline{X}}_t + \varepsilon_t^1 \tag{2}$$

From Columns (1) to (5), the coefficient point estimates on $\Delta\overline{\Phi}_t^{sub}$ are negative with similar magnitudes across all specifications, and significant at the 10% level or above for 3 out of 5 specifications; the coefficient point estimates on $\Delta\overline{Sovereign\ CDS}_t^{sub}$ are positive and significant at the 1% level with similar magnitudes across all specifications. This is consistent with our prior, suggesting that non-U.S. banks tend to reduce relative demand for USD denominated assets when the safety and liquidity of substitute currencies increases. The coefficient point estimates on $\Delta\overline{Leverage}_t$ are positive and significant at the 1% level with similar magnitudes across all specifications. This suggests that non-U.S. banks tend to increase relative demand for USD denominated assets when their balance sheet capacity to bear risk deteriorates. In terms of economic significance, the variation from those three instrumental variables alone accounts for around 20% of the variation in the contemporaneous changes of relative USD asset demand from non-U.S. banks. Based on the critical values reported in Stock and Yogo (2002), we can reject the null hypothesis that the proposed variables are weak instruments for all specifications at the 5% significance level or above.[26] Another interesting thing to notice is that the average relative demand for USD denominated assets from non-U.S. banks ($\Delta\overline{D}_{\$,t}$) is strongly correlated to the average U.S. Treasury premium whenever we control for it in Columns (2), (4) and (5) – it suggests the our quantity measure of relative demand for USD denominated assets from non-U.S. banks is correlated with the price measure of demand for USD safe assets proxied by U.S. Treasury premium.

The lower panel of Table 4 presents the results for the second stage. Again, we investigate the relationship between USD asset demand from non-U.S. banks and the USD exchange rate by running:

[24] Controls included in the first stage are the same as the ones in the second stage.

[25] The superscripts 1 (or 2) differentiate the first stage (or the second stage) in the Equation 2 (or 3).

[26] According to Stock and Yogo (2002), the critical value for the weak instrument test based on 2SLS bias of the 5% significance level is 13.91 for parameters setting of 1 endogenous regressor, 3 instrumental variables, with desired maximal bias of IV estimator relative to OLS at 0.05.

$$\Delta\bar{s}_t = \alpha_1^2 + \beta_1^2 \Delta\widehat{\bar{D}}_{\$,t} + \beta_2^2 \Delta\bar{\Phi}_t + \beta_3^2 \Delta\left(\bar{y}_t^{Govt} - y_{\$,t}^{Govt}\right) + \beta^2 \bar{\boldsymbol{X}}_t + \varepsilon_t^2 \tag{3}$$

As a benchmark, the Column (0) of lower panel presents the OLS regression by using the observed value ($\Delta D_{\$t}$), which is the result presented in Column (1) of Table 3. Specifications in Columns (1) to (5) are the same as the ones in Table 3, except instead of the observed values ($\Delta\bar{D}_{\$,t}$) we use the *fitted* values of USD asset demand ($\Delta\widehat{\bar{D}}_{\$,t}$), which are purely driven by demand conditions of USD denominated assets of non-U.S. banks obtained from the first stage, but not by the supply of USD denominated assets. From Columns (1) to (5), the coefficient point estimates on $\Delta\widehat{\bar{D}}_{\$,t}$ are positive and significant at the 1% level across all specifications.

Note that the magnitude of the coefficient point estimates on $\Delta\widehat{\bar{D}}_{\$,t}$ obtained by 2SLS from Columns (1) - (5) are around two times as larger as those obtained by OLS in Column (0) (and similarly in Table 3), and we can reject that the USD asset share is an exogenous variable from all specifications at the 10% level or higher based on the F statistic of Durbin-Wu-Hausman test of endogeneity. This result suggests the slope of the supply curve is underestimated by the OLS estimation, because observed data in relative quantities (USD denominated assets as a percent of total assets) and relative prices (exchange rates) is a reflection of a set of equilibrium points on both supply and demand curves. In addition, we confirm that three instrumental variables together are exogenous based on inference from the Sargan χ^2 statistics of overidentification.

In order to valid our IV strategy further, we try our best to eliminate several alternative hypotheses and refine our IV strategy in the following discussion. We add refinement based on the specification in Column (4) in Table 4 and the results are presented in Table 5, where the Column (0) serves as the benchmark for comparison.[27]

First, one potential concern on our instrumental variables of safety in substitute currencies is those two measures could capture not only relative demand for USD denominated assets against foreign assets but also capture the supply condition of safe assets. For example, the Treasury premium of substitute economies may decrease because of over-supply in safe assets in both the U.S. and substitute economies (Du, Im, and Schreger 2018). To help establish a clean demand shifter, we control for the change in general government debt (consolidated) to GDP ratio of both the U.S. and substitute economies.[28] The results are presented in Column (1) of Table 5. Second, the non-U.S. banks' balance sheet capacity can be challenged by its correlation with non-US banks' propensity to provide USD denominated assets. We control for the changes in USD funding gap ratio (measured as net USD denominated assets to USD denominated assets ratio)[29]. More negative (positive) USD funding gap for non-U.S. banks implies lower (higher) propensity to serve as a USD lender (borrower) on the global USD market (Barajas et al. 2020). The results are presented in Column (2) of Table 5. Third, another concern is our IVs are subject to the global financial condition, growth prospect of US and global economy, and those could potentially affect the USD exchange rate

[27] The key message is robust if we alter the benchmark to other specifications in Table 4 or if we split the sample into two halves (2001q2-2009q4, 2009q4-2017q4), and those results could be provided upon request. The reason we prefer the specification in Column (4) and use the whole sample to present the results is because it not only controls the most important drivers of spot moves discussed in the literature but also it leaves us some degree of freedom to add other controls given the exercise is constrained by a short sample.

[28] The substitute economies are consistently selected as previous discussion as six non-USD G10 economies, including Canada, Germany (as representative of Euro Area), Japan, the United Kingdom, Switzerland, and Sweden. In addition, the results are also robust to: a) use the total government debt of Euro Area instead of Germany only, b) restrict to only Germany, Japan and Switzerland for a small set of safe heavens, c) use 1-year change in Government Debt/GDP ratio instead of 1-quarter change.

[29] Following Barajas et al. (2020), the USD funding gap ratio for each non-U.S. economies is constructed under the international position on a nationality basis consistent with our measures of USD assets share.

against foreign currencies directly. To eliminate this concern, we control the level of global FCI in Column (3), the consensus forecast of US growth and consensus forecast of Global growth in the next 12 months in Column (4), and the 1-quarter-change in VIX in Column (5). Fourth, one alternative hypothesis is the relative demand for USD denominated assets from non-U.S. banks is driven by speculative reasons (e.g., carry trade) rather than safe assets demand. To eliminate this concern, we control the 1-quarter-change in carry between the U.S. and other 16 economies in the sample, and the result is presented in Column (6) of Table 5. In the end, we want to conclude this session by one most rigorous specification in Column (7) of Table 5, which includes all controls discussed above.

From the upper panel, the coefficient of point estimates on three proposed IVs are barely different from that of the benchmark and reject the weak IVs test jointly at 5% significance level or above. More importantly, there is no evidence for statistically significant correlation between the USD asset share and those controlled variables. It suggests the impact of proposed IVs are exclusively through the relative demand for USD denominated assets of non-U.S. banks proxied by USD asset share of non-U.S. banks rather than through other factors. Although *none* of the alternative factors are insignificant statistically, there are several interesting phenomena. For example, more supply of safe government assets from the U.S. government corelates with higher USD assets to total assets of non-U.S. banks, but more supply of safe government assets from the closest substitute economies crowds out the USD assets to total assets of non-U.S. banks. In addition, the coefficient point estimate on 1-quarter changes in carry is negative but insignificant, which implies the relative demand for USD denominated assets over foreign currencies denominated assets from non-U.S. banks is mainly driven by safe assets demand or balance sheet capacity of non-U.S. banks rather than speculative reasons.[30] Note that the coefficient of U.S. Treasury premium is significant in all specifications, which suggests the USD asset share of non-U.S. banks are not only significantly correlated three IVs discussed above but also significantly correlated to the price measure of safe assets demand for U.S. 1-yr Treasuries. It further helps us to refine that our measure of USD asset share is an alternative measure for safe assets denominated in USD but in quantity measure.

From the lower panel, we find the coefficient point estimates on fitted value of USD asset share ($\Delta\widehat{\widehat{D}}_{\$,t}$) barely change or slightly lower than that is obtained in the benchmark 2SLS regression but are still at least as large as 1.7 times as that obtained from OLS regression (Column (4) in Table 3), and still significant at 5% or above. In addition, we can reject that the USD asset share is an exogenous variable from all specifications at the 10% level or higher based on the F statistic of Durbin-Wu-Hausman test of endogeneity. This result confirms our finding in the benchmark 2SLS regression that the slope of the supply curve is underestimated by the OLS estimation, because observed data in relative quantities (USD denominated assets as a percent of total assets) and relative prices (exchange rates) is a reflection of a set of equilibrium points on both supply and demand curves. In addition, we confirm that three instrumental variables together are exogenous based on inference from the Sargan χ^2 statistics of overidentification.

In summary, this result confirms that our instrumental variable (IV) strategy manages to identify the supply curve and corrects the coefficient point estimates using valid demand shifters that are exogeneous to supply conditions. Conditional on different alternative hypothesis and potential factors that could contaminate the IVs, the robust results further demonstrate the validity of our instrumental variable (IV) strategy. In addition, we have more confidence that the relative USD asset demand from non-U.S. banks

[30] Alternatively, we look at the results on average of G10 (or EM economies) instead of the average of all 16 economies after controlling the carry between the U.S. and G10 (or carry between the U.S. and EM), and the key message is robust – it doesn't matter for G10 or EM sample that the changes in USD assets share is not significant correlated to changes in carry. The results can be provided upon request.

is mainly driven by safe assets demand by ruling out the alternative hypothesis that it is driven by speculative reasons (e.g., carry) and demonstrating it is strongly correlated with price measure of safe assets demand (e.g., Treasury premium).

C. Transmission Channels

An interesting extension is to investigate the relationship between the U.S. Treasury premium, the LIBOR cross-currency basis, and the quantity of relative USD asset demand from non-U.S. banks using our IV strategy. This can add value to the discussion of the "safe asset demand channel" and the "financial intermediary channel" for the global USD cycle (Du, Im, and Schreger 2018; Jiang, Krishnamurthy, and Lustig 2018; and Du 2019). Both the LIBOR cross-currency basis, which equals the LIBOR CIP deviation, and the U.S. Treasury premium are measures of the "specialness" of USD denominated assets relative to non-USD denominated assets on a currency hedged basis. The key difference is that U.S. financial intermediaries represent key suppliers of USD denominated assets in the interbank market, while the U.S. Treasury is the ultimate supplier of U.S. Treasury securities in the U.S. Treasury market. There is quite a distinction behind this – the "safe asset demand channel" is about shifters to the demand curve, but the "financial intermediation channel" is about shifters to the supply curve. Using exogeneous demand shifters from our IV strategy, we can identify the slopes of the supply curves for these two markets.

Table 6 only presents the results of the second stage, because the first stage results are very similar to that of the upper panel in Table 4. The specification of the second stage is given by[31]:

$$\Delta\overline{\Phi}_t = \alpha_1^2 + \beta_1^2 \Delta\widehat{\overline{D}}_{\$,t} + \beta_2^2 \Delta\left(\overline{y}_t^{Govt} - y_{\$,t}^{Govt}\right) + \beta^2 \overline{\boldsymbol{X}}_t + \varepsilon_t^2 \quad (4)$$

Where instead of using average quarterly changes in log nominal exchange rate as dependent variable, we investigate average quarterly changes of the U.S. Treasury premium ($\Delta\Phi_t^{Treasury}$) and average quarterly changes of the (-1)*LIBOR cross-currency basis ($\Delta\Phi_t^{Libor}$) in Columns (1) - (2) and Columns (3) - (4), respectively.[32]

In the regressions of the U.S. Treasury premium ($\Delta\Phi_t^{Treasury}$), Column (1) shows that the 2SLS coefficient point estimate on the fitted value of the USD asset share ($\Delta\widehat{\overline{D}}_{\$,t}$) is positive and statistically significant during the full sample (2001q2-2017q4), and it remains significant and becomes even larger since the GFC (2010q1-2017q4) in Column (2). These results coincide with findings from Du, Im, and Schreger (2018), and Jiang, Krishnamurthy, and Lustig (2018) that document the U.S. Treasury premium has been positive for a long time even before the GFC. Because the fitted value of the USD asset share ($\Delta\widehat{\overline{D}}_{\$,t}$) comes from exogeneous demand shifters, it suggests the supply curve in the U.S. Treasury market identified by our IV strategy is upward sloping for both the full sample episode and the post-GFC episode, though the slope is slightly larger for the latter but insignificant in difference. On the contrary, in the regressions of the (-1)* LIBOR cross-currency basis ($\Delta\Phi_t^{Libor}$), Column (3) shows that the 2SLS coefficient point estimate on the fitted value of USD asset share ($\Delta\widehat{\overline{D}}_{\$,t}$) is not significantly different from zero during the full sample (2001q2-2017q4), however, it turns positive and statistically significant after the GFC (2010q1-2017q4) in

[31] The quarterly changes in the U.S. Treasury premium ($\Delta\Phi_t$) is not entered as control, because it becomes the dependent variable here. For the similar reason, the average Treasury premium of substitute government bonds ($\Delta\overline{\Phi}_t^{sub}$) is excluded from the instrumental variables.

[32] According to Du, Tepper, and Verdelhan (2018), the LIBOR cross-currency basis between jurisdiction j and the United States is equal to $\mathrm{x}_{j,t} \triangleq -(y_{j,t}^{Libor} - \rho_{j,t} - y_{\$,t}^{Libor})$, where $y_{j,t}^{Libor}$ and $y_{\$,t}^{Libor}$ denote 1-year interbank rate of jurisdiction j and that of the United States, respectively, and $\rho_{j,t}$ is the one-year market-implied forward premium for hedging currency j against the U.S. dollar. To make the notation consistent in the paper, the variable Φ_t^{Libor} denotes $(-1) * \mathrm{x}_{j,t}$ averaging over 16 jurisdictions.

Column (4). These results again echo the findings from Du, Tepper, and Verdelhan (2018) that CIP deviation has only become significantly negative since the GFC. In the LIBOR cross-currency swap market, we cannot reject that the supply curve is flat for the full sample episode, but we find that the supply curve becomes significantly upward-sloping in the post-GFC episode. These results suggest that U.S. financial intermediaries actively traded in the LIBOR cross-currency swap market and arbitraged away the CIP deviation pre-GFC, but they have had less balance sheet capacities to do so since the GFC. In contrast, the U.S. Treasury never exhibits active speculative trading behavior even if it has the balance sheet capacity to do so, which is because the U.S. Treasury acts according to fiscal mandates.

[INSERT TABLE 6 HERE]

As far as we know, our paper is the first one to provide comprehensive empirical evidence to understand the premise that the "safe asset demand channel" and the "financial intermediation channel" play different roles in different markets. Our findings suggest that the "safe asset demand channel" is at play for both the U.S. LIBOR cross-currency swap market and the U.S. Treasury market (Jiang, Krishnamurthy, and Lustig 2018; Krishnamurthy and Lustig 2019). In addition, our findings suggest that the "financial intermediation channel" has started to play an important role in the U.S. LIBOR cross-currency swap market, because the global capital and liquidity requirements and specific regulations at the individual economy level may have tightened the supply of US dollar funding to non-U.S. banks since the GFC (Du, Tepper, and Verdelhan 2018; Iida, Kimura, and Sudo 2018). Our findings also suggest that the "financial intermediation channel" doesn't play a critical role in the U.S. Treasury market, because the U.S. Treasury (as the ultimate supplier of U.S. government bonds) can but does not actively arbitrage away the U.S. Treasury premium when the demand for U.S. Treasury from non-U.S. banks is strong.

Overall, our novel IV strategy helps identify the causal relationship between the USD asset share of non-U.S. banks and the USD exchange rate, correcting the downward bias of the OLS estimator. Further, it sheds new light on the "safe asset demand channel" and the "financial intermediation channel".

V. Cross-Sectional Asset Pricing

In addition to the strong contemporaneous causal relationship, we turn to asset pricing in the cross-section. In this section, we find that different loadings on the average USD asset demand of non-U.S. banks help to explain currency excess returns in the cross section. Our results suggest that average global demand of USD denominated assets from non-U.S. banks acts as a risk factor in pricing currency excess returns.

The risk factor of global USD asset demand is defined as the cross-sectional *average* of annualized quarterly changes in the USD asset share of non-U.S. banks across 16 economies. The log currency excess return of foreign currency J corresponds to a long position in Treasury bonds of foreign economy j funded by borrowing in U.S. dollars:

$$R_{j,t} \triangleq y_{j,t-1}^{Govt} - \Delta s_{j,t} - y_{\$,t-1}^{Govt} \tag{5}$$

To estimate the factor price and the loading betas, we use a two-stage Fama and MacBeth (1973) regression. We conduct this analysis both at the individual-currency level and at the currency-portfolio level provided by Lustig, Roussanov, and Verdelhan (2011). We try to control for USD factor and carry factor as the most important cross-sectional pricing factors following Verdelhan (2018).[33] In the first stage, we run a time-series regression of the currency excess returns on a constant, USD factor, and global USD asset demand changes as pricing factor to estimate betas.[34] The betas for individual currencies (currency portfolios) are presented in Figure 3 below.

[INSERT FIGURE 3 HERE]

From the left panel of Figure 3, we can find the loadings on global USD asset demand are negative and significant at the 5% level or above for 15 out of 16 currencies. It suggests that higher interest rate currencies (e.g., BRL and ZAR) have higher loadings (in absolute term) than lower interest rate currencies (e.g., JPY and GBP) on average even after controlling for USD factor following Verdelhan (2018). Furthermore, we look at the loading of excess returns of carry portfolios on our global USD asset demand changes with USD factors controlled. We use the currency portfolios from Lustig, Roussanov, and Verdelhan (2011), which are ranked from low to high interest rate – portfolio 1 contains currencies with the lowest interest rate, while portfolio 5 contains currencies with the highest interest rate.[35] In the right panel, we also find higher interest rate currency portfolios tend to have higher loadings in general.[36]

In the second stage, we examine the cross-sectional relationship between the magnitude of exposure to USD asset demand and the average size of excess returns. To visualize this cross-sectional relationship, Figure 4 plots the average currency excess returns against the corresponding USD demand betas for individual currencies (in the left panel), and for currency portfolios (in the right panel), respectively. We can see a strongly negative correlation between the average excess returns and the USD demand betas with a bivariate correlation equal to -47% for the individual currencies and -76% for the currency portfolios.

[INSERT FIGURE 4 HERE]

From the left panel in Figure 4, we can see that investing in high interest rate currencies (e.g. BRL and ZAR) gives U.S. investors on average higher excess returns than investing in low interest currencies (e.g. JPY, GBP, and SGD). At the same time, high interest rate currencies have larger negative loadings on the global average USD asset demand factor than low interest rate currencies. This suggests that the demand of USD assets of non-U.S. banks is a global risk factor on the FX market. High interest rate currencies have higher systematic loadings on the USD demand factor (a more negative USD asset demand beta), which require higher on average excess returns to compensate this exposure to USD asset demand risk. On the contrary, low interest rate currencies provide a hedge against USD asset demand risk.

The right panel for carry-trade portfolios ranked by interest rate demonstrates an even stronger relationship.[37] This finding relates to the discussion of "currency crashes" by Brunnermeier, Nagel, and Pedersen (2009). It also coincides with the conclusion of Clarida, Davis, and Pedersen (2009) and Menkhoff, Sarno, Schmeling, and Schrimpf (2012) that the *ex post* returns for high interest rate currencies

[33] In addition, our results are also robust by controlling intermediary asset pricing factor (i.e., leverage ratio of U.S. banks). The results can be provided upon request.
[34] The USD factor is proxied by 1-quarter difference in logarithm of nominal effective exchange rate index (NEER) of USD from JP Morgan, and the results are also robust if we use the NEER complied by BIS or IMF.
[35] As Lustig and Verdelhan (2007) suggested, we drop the last currency portfolio, which consists of currencies with very high inflation.
[36] We should admit the cross-sectional differences cannot be fully explained by the differences in the interest rates (e.g. INR, SGD in the left panel; currency portfolio 1 in the right panel).
[37] Burnside, Eichenbaum, and Rebelo (2007 and 2011) find the excess returns on carry trade strategy cannot be explained by traditional risk factors such as growth rate of real consumption, the market return, the term structure spread, and the Ted spread, etc.

are low during times of high volatility. High interest rate currencies are risky because they have poor payoffs when a measure of global volatility is high, which coincides with strong demand for USD denominated safe assets.

The cross-sectional relationship discussed above has some linkages with the cross-sectional relationship between the broad U.S. dollar index and the LIBOR cross-currency basis. Avdjiev, Du, Koch, and Shin (2019) find currencies with more systematic exposures to the broad U.S. index have higher excess returns from CIP trade. When the global relative demand for USD denominated from non-U.S. banks rises, the USD exchange rate tends to appreciate. As a result, the currencies with more exposures to the USD asset demand risk tend to have higher expected excess returns from CIP trade.

Our finding is consistent with recent exchange rate theories based on capital flows in imperfect financial markets that links global imbalances with currency premia (Gourinchas and Rey 2007; Corte, Riddiough, and Sarno 2016). These theories suggest net debtor countries (usually high interest rate countries) offer currency risk premia to compensate investors willingness to finance negative external imbalances because their currencies depreciate in bad times, which are times of large negative shocks to risk-bearing capacity and global risk aversion inducing demand for USD assets via “safe asset demand” channel.

In summary, our findings suggest that the average USD asset demand from non-U.S. banks acts as a global risk factor that is priced in the cross section of currency excess returns. The currencies with more systematic exposures to the USD asset demand risk factor tend to have higher expected returns to compensate risks of “currency crashes” in bad times, which are times coincides with strong demand for safe USD assets.

VI. Predictability

In this section, we focus on the performance of the USD asset share from non-U.S. banks (a quantity measure) in forecasting exchange rates. In the literature, it is well known that some FX market quantities like order flow are associated with contemporaneous returns (e.g., Evans and Lyons 2002), however, it is less known whether predictive information is contained.[38]

To begin with, we illustrate the predictive power using a simple vector autoregression (VAR). The impulse response functions (IPFs) from the VAR are shown in Figure 5. In the left panel, we find that the contemporaneous appreciation in the USD exchange rate against foreign currency in response to positive shocks in USD asset share of non-U.S. banks is nearly entirely undone after five years. One of the economic interpretations is overshooting due to time varying risk premia, where relative demand for USD assets of non-U.S. banks drives up the U.S. dollar against corresponding foreign currency contemporaneously by compressing risk premia, but this contemporaneous impact gets reversed over 1-5 years as risk premia normalize. Lilley, Maggiori, Neiman, and Schreger (2019) also find that the U.S. purchase of foreign bonds is associated with risk premia after the GFC.

[INSERT FIGURE 5 HERE]

[38] A few papers show that FX order flow contains information about future currency returns, but they tend to disagree on the source of this predictive power (e.g., Evans and Lyons 2005; Froot and Ramadorai 2005; Rime, Sarno, and Sojli 2010). However, Sager and Taylor (2008) fails to find robust predictive power of exchange rates by order flow, using commercially available order flow data.

An alternative explanation is that the USD asset share of non-U.S. banks could capture a dimension of investor sentiment that is not captured by other macro fundamentals. For example, López-Salido, Stein, and Zakrajšek (2017) and Brandão-Marques, Chen, Raddatz, Vandenbussche, and Xie (2019) document that risk metrics from credit markets can help to predict reversals of financial conditions and corporate spreads. In addition, this can also reflect the extraordinary ability of non-U.S. banks to process fundamental information, especially as key marginal investors in the local foreign currency market. For example, Menkhoff, Sarno, Schmeling, Schrimpf (2016) find that customer types vary markedly in terms of their predictive ability – the flows by long-term demand-side investment managers have the strongest predictive power for exchange rates, while flows originated from the other groups only predict transitory changes in exchange rates. However, fully disentangling the extent to which the relationship between USD asset share of non-U.S. banks and the exchange rate of foreign currency against USD by any of these dimensions is beyond the scope of this paper.

In the middle panel, we find that the U.S. Treasury premium drives up the U.S. dollar contemporaneously and the induced momentum effect is short-lived for another quarter, then it depreciates over the next one-to-five-year period. This result is consistent with findings from Jiang, Krishnamurthy, and Lustig (2018). In the right panel, we find the U.S. dollar appreciates in response to positive shocks to interest differentials (home – U.S.) in the first two years, then it wanes down over the next three-year period.[39]

A. In-sample forecasting

As suggested by the theoretical models by Jiang, Krishnamurthy, and Lustig (2018) and Engel and Wu (2018), a strong relative demand for USD denominated assets from non-U.S. banks today means that foreign currency depreciates against USD today, thus lowering the expected future returns from owning USD denominated assets and inducing a subsequent appreciation of foreign currency against USD in the future.[40] In particular, we run the following regression:[41]

$$\Delta s_{j,t}^{h} = \alpha_1 + \beta_1 D_{j,\$,t} + \beta_2 \Phi_{j,t} + \beta_3 \left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right) + \varepsilon_{t+h} \quad (6)$$

where the dependent variable $\Delta s_{j,t}^{h}$ is the annualized change in the log nominal exchange rate of USD vis-à-vis foreign currency J, an increase of which means that the USD appreciates against foreign currency J over the horizon of $[t, t+h]$. The variable $D_{j,\$,t}$ denotes the USD asset share of non-U.S. banks from foreign economy j at end of quarter t. The variable $\Phi_{j,t}$ denotes the U.S. Treasury premium vis-à-vis foreign economy j at the end of quarter t. The variable $\left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right)$ denotes the interest differential between 1-year Treasury securities of foreign economy j and that of the U.S. Treasury at end of quarter t. We run this regression using quarterly data as before, but vary by horizons from one-quarter, one-year, two-year, three-year, to five-year.

[39] The same relationship is found in papers investigating the impact of monetary policy on Emerging Markets exchange rates (see, for example, Kohlscheen 2014; Hnatkovska, Lahiri and Vegh 2016; Hofmann, Shim, and Shin 2017). However, for non-USD G10 currencies, we find the U.S. dollar depreciates in response to positive shocks in interest differentials (home – U.S.) in the two quarters, then it wanes down over the next four-year period. These results coincide with the models presented by Engel (2016) and Valchev (2020) that UIP violates in the short-run and reverses in the longer horizon. To better understand different implications for G10 countries and emerging markets with respect to UIP puzzle is beyond the scope of this paper, but relevant results on non-USD G10 currencies can be provided upon request.

[40] Jiang, Krishnamurthy, and Lustig (2018), and Engel and Wu (2018) use a widening of U.S. Treasury premium to capture excessive demand for USD safe assets. In this paper, we directly use USD assets share of non-U.S. banks to capture it.

[41] The conclusion of the results is robust if we control $\Delta\Phi_{j,t}$, and $\Phi_{j,t-1}$ as Jiang, Krishnamurthy, and Lustig (2018).

Table 7 reports our forecasting results of the U.S. dollar exchange rates against a basket of 16 foreign currencies. The USD asset share doesn't help predict the USD exchange rate at the short end but has very strong predictive power over the longer horizon. This suggests time variation in exchange rate risk premia. In particular, it reflects the stationary but persistent deviation from the uncovered interest parity that Duarte and Stockman (2005) and Engel (2014 and 2016) have documented. The coefficient point estimates on the USD asset share of non-U.S. banks ($\overline{D}_{\$,t}$) start from -0.76 (not significant) at the one-quarter horizon, and then stabilizes to a level between 2.19 and 3.17 per annum from one-year to five-year horizon at 5% significance level or above. This means that a positive shock to USD asset share by one-standard deviation at 0.32% (from Table 1) is contemporaneously correlated with an appreciation of the U.S dollar against a basket of 16 foreign currencies by 8.1% on average (Column 5 of Table 3) per annum in that year, but it is expected to depreciate by 0.70% (Column 5 of Table 7) per annum over the next five years on average to wane down the overshooting.

As Jiang, Krishnamurthy, and Lustig (2018), we also find the induced momentum effect from the U.S. Treasury premium ($\overline{\Phi}_t$) is short-lived and the coefficient point estimate switches sign over the long run, although it is only statistically significant at the two-year horizon.

The Uncovered Interest Rate Parity (UIP) suggests the coefficient before interest differential equals to 1. However, numerous empirical studies consistently reject the UIP condition (see Engel 1996; Sarno 2005). This is the case if we only include non-U.S. G10 economies in our sample.[42] However, the interest differential ($\bar{y}_t^{Govt} - y_{\$,t}^{Govt}$), which is a proxy for the risk premium of carry trade excess returns, positively and significantly forecasts the USD exchange rate appreciation for one-year, and two-year horizons at 10% significance level or above. At short horizon, all these variables don't help predict the dynamics of the U.S. dollar exchange rate, where the one-quarter ahead R^2 is as low as 7%. Gradually, these regressors jointly account for more variation in the moves of exchange rate at longer horizon – the R^2 increases to 25% at the one-year horizon and strikingly reaches 68% at the five-year horizon.

[INSERT TABLE 7 HERE]

In Table 8, we report our forecasting regression results of bilateral exchange rate over the five-year horizon. For better comparison, a simple univariate regression is used here:

$$\Delta s_{j,t}^{20} = \alpha_1 + \beta_1 D_{j,\$,t} + \varepsilon_{t+20} \tag{7}$$

In Panel A, the regressor is $D_{j,\$,t}$, that is the *individual* USD asset share of non-U.S. banks in economy j at end of quarter t. The forecasting results of the bilateral exchange rate confirms our finding: the increase in USD asset share from each economy predicts the USD exchange rate depreciates against 10 out of 16 foreign currencies, where 8 out of 10 are at the 5% significance level or above. In Panel B, the regressor is $\overline{D}_{\$,t}$, that is the *cross-sectional average* of USD asset shares over 16 economies at end of quarter t. The forecasting power is even stronger – it helps predict the bilateral exchange rate at the 5% significance level or above for 15 out of 16 foreign currencies in our sample. The point estimates on $\overline{D}_{\$,t}$ vary from as low as 0.65% per annum (for MXN) to as high as 5.60% per annum (for MYR). While, the R^2 varies from 10% (for MXN) to 69% (for NOK), which is on-average higher than that of Panel A using the USD asset share from individual economies as regressors. These results suggest that not only the global component

[42] This univariate estimate is not reported, where the coefficient point estimates of interest rate differential are -0.6, -0.8, and -0.4 for non-U.S. G10 jurisdictions at horizons of 1-quarter, 1-year and 2-year.

of USD asset share of non-U.S. banks forecast the individual foreign exchange rate against USD, but also the variation in USD asset share of non-U.S. banks of individual economies also helps as well.

[INSERT TABLE 8 HERE]

Taken together, these results strongly suggest that exchange rate movements are associated with time varying risk premia, which covary with the relative demand for USD denominated assets by non-U.S. banks. The overshooting of Sections IV.A and VI.A. suggests that demand for USD assets leads to a sharp contemporaneous appreciation of the USD, which is reverted slowly over time. In addition, the results of cross-sectional asset pricing in Section V points to a coherent intermediary asset pricing story. The scarcity of USD denominated assets is driven to an important extent by risk-taking capacity of non-U.S. banks, which enters the pricing kernel of these marginal investors on the FX market, thus driving time varying currency risk premia that prices foreign currency against US dollar. We next turn to the out-of-sample exercise, where we show that the strong predictability holds even out of sample, thus, undermining one of the puzzles of the exchange rate literature (Meese and Rogoff, 1983).

B. Out-of-sample forecasting

In this subsection, we turn to the out-of-sample forecasting performance following the tradition established by Meese and Rogoff (1983). We find our forecasting model to significantly outperform the random walk prediction.

We consider an out-of-sample forecasting exercise with a period of K quarters to obtain in-sample coefficient point estimates and a period of P quarters to evaluate the model's performance out of sample. We denote the end of the model evaluation as quarter T. To estimate model parameters, we regress the annualized change of the foreign exchange rate against USD (over an h-quarter horizon) only on the USD asset share to obtain in-sample estimates of the model parameters within a window spanning K quarters:

$$\Delta s_{j,t}^{h} = \alpha_1 + \beta_1 D_{j,\$,t} + \varepsilon_{t+h} \tag{8}$$

Where, $t \in [T-P-K, T-P]$. We then use the estimated coefficients $\hat{\alpha}_1$ and $\hat{\beta}_1$, and the realized observation of the USD asset share to forecast the exchange rate change in the next period:

$$\Delta \hat{s}_{j,T-P+1}^{h} = \hat{\alpha}_1 + \hat{\beta}_1 D_{j,\$,t} \tag{9}$$

The model's forecast error is the difference between the realized exchange rate change and the model-implied forecast:

$$error_{j,T-P+1}^{h} = \Delta s_{j,T-P+1}^{h} - \Delta \hat{s}_{j,T-P+1}^{h} \tag{10}$$

We repeat this procedure for P times using a rolling regression approach until we have an out-of-sample forecast for all quarters in the period between quarters $T-P$ to T. We compute the mean-square-error (MSE) of our forecasting model and compare it against that of the random walk benchmark. First, we adopt the out-of-sample (OOS) statistic from Diebold and Mariano (1995) and West (1996) based on MSE_r – MSE_u to evaluate performance of the alternative model. Because the random walk benchmark is nested in the "unrestricted" specification, we can evaluate its performance using the inference procedure by Clark and West (2006) to adjust difference in mean squared errors: MSE_r – ($MSE_u - Adj.$), which accounts for the small-sample forecast bias.

Table 9 presents the out-of-sample forecasting performance of our exchange rate forecast model for the five-year horizon ($h = 20$).[43] In Panel A, the Diebold-Mariano difference in MSEs and the OOS-T statistic are reported.[44] In Column (1), we report the result that uses the cross-sectional average series over 16 currencies. We find our forecast model outperforms the benchmark random walk model at the 1% significance level.

In Columns (2) to (17), we repeat this exercise for individual economy. We find that our forecasting model outperforms the benchmark random model for 15 out of 16 currencies at the 10% significance level or above. In Panel B, the Clark-West adjusted difference in MSEs and the C-W statistic are reported. Similarly, our forecasting model outperforms the benchmark random walk model at the 1% significance level for the USD exchange rate against a basket of 16 currencies in Column (1). More strikingly, our forecasting model outperforms the benchmark model in 15 out of 16 currencies at 10% significance level or above after adjusting for the small-sample forecast bias from Columns (2) to (17).[45] To our knowledge, the strength of this out-of-sample forecasting power from a quantity variable is unprecedented in the literature on exchange rates.

[INSERT TABLE 9 HERE]

Overall, our results suggest a strong relative demand for USD denominated assets from non-U.S. banks today means that the USD exchange rate appreciates against foreign currency contemporaneously, which induces an expected depreciation in the future. This is an overshooting result in the spirit of Dornbusch (1976). However, the economic interpretation is one relying on relative scarcity of USD denominated assets to assets denominated in corresponding foreign currency that is driven largely by non-U.S. banks as the marginal investors in the FX market (especially for foreign currency). We also find the intermediary asset pricing factor (e.g. the leverage of non-U.S. banks) to be one of the candidate explanatory variables. Time varying FX risk premia are related to the relative demand of USD denominated assets from non-U.S. banks in a systematic fashion, thus creating forcastability of the exchange rate. The forecasting power of the USD asset share of non-U.S. banks in exchange rates over longer horizons is robust for both in-sample and out-of-sample regressions.

C. Robustness

We now turn to the robustness in forecastability of exchange rates. We compare the forecasting performance of our model with 9 exchange rate forecasting models discussed in the literature since the 1970s. More specifically, we compare the forecasting performance of augmented versions that include

[43] We set $P = round(T/2)$ and $K = T - P$ as our baseline. T (K and P) is varying by currency – the longest T covers 67 quarters from 2001q1 to 2017q4. Alternatively, we set $P = round(T * 2/3)$, and the results are robust. In addition, our key message in out-of-sample forecasting analysis is also robust to three-year or four-year horizon. Relevant robust results can be provided upon request.

[44] As suggested by Clark and West (2006), the heteroskedasticity and autocorrelation consistent (HAC) standard error are estimated to compute the OOS-T statistics and C-W statistics. They are obtained from a regression of squared error loss differential (with adjustment) on a constant. The corresponding 90%, 95%, and 99% critical values are 0.780, 1.111, and 1.784 (using $k_2 = 1$ and $\pi = 0.2$) obtained from Table 1 in McCracken (2007).

[45] Taking into account the global component of the USD asset share, we use the cross-sectional average of the USD asset share over 16 jurisdictions as predictor. If we use the USD asset share of individual jurisdiction, we find our forecast model outperforms benchmark models in 13 out of 16 currencies at the 10% significance level or above.

the USD asset share of non-U.S. banks as an additional predictor in existing model discussed in the literature.

Guided by the excellent survey of exchange rate predictability in Rossi (2013), the models we consider for the horse race include: (1) the UIP model; (2) the monetary model with flexible prices ("Frankel-Bilson" model); (3) the monetary model with sticky prices (the "Dornbusch-Frankel" model); (4) the model with productivity differentials (the "Balassa-Samuelson" model); (5) the Taylor rule model; and (6) the net foreign asset model (the "Gourinchas-Rey" model). In addition, we take into account of the most recent development in exchange rate forecast including: (7) the dollar liquidity model suggested by Adrian, Etula, and Shin (2010); (8) the Treasury premium model suggested by Jiang, Krishnamurthy, and Lustig (2018); and (9) the U.S. foreign bond flow model suggested by Lilley, Neiman, Maggiori, and Schreger (2019).

Table 10 compares the out-of-sample forecasting performance of the augmented exchange rate forecasting model by adding the USD asset share as an additional predictor to existing model against each existing model stated above. The Diebold-Mariano difference in MSEs and the OOS-T statistic are reported for each comparison. We find the $MSE_{existing} - MSE_{augmented}$ is positive across all models and that the OOS-T statistic is significant at the 10% level or above across all 9 comparisons. This suggests that including the USD asset share as an additional predictor improves the performance of out-of-sample forecasting for each of the exchange rate models discussed above.

[INSERT TABLE 10 HERE]

Finally, our findings suggest that the average USD asset demand from non-U.S. banks measured by the cross-sectional average of USD asset share from non-U.S. banks across 16 economies helps to improve the forecasting performance of 9 existing exchange rate models discussed in the literature since the 1970s. Clearly, exchange rates are deviating significantly from the random walk benchmark as a function of the USD asset demand share of non-U.S. banks, thus pointing strongly towards exchange rate overshooting as a function of time varying intermediary risk premia.

VII. Implications for the Literature

We summarize our paper's key implications for four strands of the literature.

First, our paper contributes to the literature on the United States' special role as the provider of world safe assets (e.g., Gourinchas and Rey 2007; Gourinchas, Rey, and Govillot 2010; Farhi and Maggiori 2018). Maggiori (2017) shows a model with a simple asymmetry in global risk sharing and heterogeneity in financial development can rationalize the economic role of the U.S. in the global financial architecture. As financial intermediaries in the U.S. are better able to deal with funding problems following negative shocks, the United States consumes more relative to the rest of the world (RoW) and runs a trade deficit based on higher financial income that it earns as compensation for greater risks it takes. On the contrary, the financial intermediaries in the RoW accumulate precautionary long positions in USD assets for safety and liquidity consideration in order to insulate their capitals from negative shocks. When bad time hits, capital losses on the external portfolio of the U.S. lead to a wealth transfer to the RoW. Naturally, the U.S. dollar emerges as the reserve currency because it appreciates during bad times, thus USD denominated assets represent global safe assets by providing a good hedge.[46] Our paper supports this theory by providing empirical evidence that the financial intermediaries from the ROW pay a counter-cyclical safety premium to hold USD denominated safe assets. In addition, we demonstrate that demand for USD denominated assets from financial intermediaries of the ROW increases strongly following negative shocks because of flight to safety, and the USD exchange rate appreciates contemporaneously, thus lowering the investors' expected future return from owning U.S. denominated assets.

Second, our paper sheds light on the large literature on the "exchange rate disconnect puzzle". Rossi (2013) gives an excellent review of exchange rate predictability especially for models before the first decade of the 21st century. Our paper is directly related to recent developments in this area from two perspectives. On the one hand, foreign investors' demand for safety, empirically captured by the Treasury premium, is linked to movements in exchange rates (e.g., Jiang, Krishnamurthy, and Lustig 2018; Engel and Wu 2018).[47] On the other hand, FX market quantities, like order flows, can help understand currency returns (e.g., Evans and Lyons 2002, 2005; Froot and Ramadorai 2005; Rime, Sarno, and Sojli 2010). Lilley, Maggiori, Neiman, and Schreger (2019) use security-level data on U.S. portfolios of US mutual funds and demonstrate that the reconnect of U.S. foreign bond purchases to exchange rates is largely driven by investment in USD-denominated assets. Our paper links these two directions together and finds that a quantity-based measure can well capture the relative foreign demand for USD denominated assets, especially for non-U.S. banks as the marginal investors in the FX market for foreign currencies. Furthermore, we use a novel instrumental variable approach to identify a causal relationship between relative USD denominated asset demand from non-U.S. banks and the corresponding foreign exchange rate against U.S. dollar. Last but not least, our paper documents the demand for USD denominated assets from non-U.S. banks is associated with time-varying currency risk premia, thus it not only helps account for contemporaneous changes in exchange rates, but also helps predict exchange rate dynamics over the longer horizon both in sample and out of sample.

Third, our paper also contributes to the literature on the twin sisters of "the global financial cycle" and "the global USD cycle". In an early version of Avdjiev, Du, Koch, and Shin (2019), they present a global banking model with a VaR constraint and show that a stronger USD coincides with a higher shadow cost of banks' balance sheet capacity, so it can price the cross-section of CIP deviations and is associated with lower growth of cross-border bank lending denominated in the USD. They point out when banks aggregated equity decrease, banks tend to reduce risky assets holding to ensure it meets the VaR constraint. Our study provides an empirical extension for their model under the setting, where non-U.S. banks are holding both USD denominated assets and foreign currency denominated assets – when aggregated equity decreases, non-U.S. banks would reduce their foreign currency denominated assets holding disproportionally stronger than the reduction in US denominated assets, because US denominated assets tend to be safer under negative shocks. This theoretical conjecture perfectly

[46] Traditional international macroeconomics models predict that a transfer of wealth from the United States to the RoW during bad times results in a U.S. dollar depreciation in the absence of trade costs and in the presence of home-bias in consumption, which was first highlighted in the classic Keynes and Ohlin debate on the "transfer problem". In Maggiori (2017)'s setup, however, RoW relative export costs increase when RoW financial intermediaries lose capital and decrease the availability of credit to RoW exporters, which leads to a shift in demand towards goods produced in the United States at bad times, and therefore induces appreciation in U.S. dollar. In a relevant study, Ivashina, Stein, and Scharfstein (2015) present a model and show evidence that USD lending by Eurozone banks fell during Eurozone sovereign crisis and firms who were more reliant on Eurozone banks before crisis had a more difficult time borrowing.

[47] Jiang, Krishnamurthy, and Lustig (2018) present an asset pricing model which suggests higher safe asset demand from foreign investors, captured by the U.S. Treasury premium, coincides with an immediate appreciation and a subsequent depreciation of the U.S. dollar exchange rate, and they find this phenomenon is only relevant for the United States, pointing to a special role of the U.S. as the world's provider of safe assets and to the special role of USD dollar as the world's reserve currency. Similarly, Engel and Wu (2018) present a New Keynesian model, but find contradicting empirical results that this phenomenon is not only a US-specific phenomenon. It suggests a sovereign's debt is safer on a relative basis but not merely on an absolute basis (He, Krishnamurthy, and Milbradt 2019).

coincides with the causal relationship and transmission channel identified by using non-U.S. banks' leverage ratio. Valchev (2020) provides an open-economy model to relate the quantity of U.S. Treasury bonds outstanding to the U.S. Treasury premium, to the return of the U.S. dollar, and to the failure of CIP, jointly. Our paper adopts an instrumental variable approach and establishes a causal relationship between USD asset demand from non-U.S. banks, the U.S. dollar exchange rate, the U.S. Treasury premium, and the LIBOR CIP deviation. As far as we know, our paper is the first one that provides empirical evidence to demonstrate that balance sheet constraints of non-U.S. financial intermediaries play a critical role in the U.S. interbank market but not in the U.S. Treasury market by identifying different elasticities of supply curves in both markets, respectively. In other words, we differentiate the importance of the "safe asset demand" channel from the "financial intermediation channel" in these two markets, which has been a premise in the literature (e.g., Jiang, Krishnamurthy, and Lustig 2018; Du 2019; and Krishnamurthy and Lustig 2019).

Fourthly, our paper also sheds light on the theory of intermediary asset pricing (e.g., Brunnermeier and Pedersen 2009; He and Krishnamurthy 2012, 2013; Adrian and Shin 2014; Brunnermeier and Sannikov 2014). Adrian, Etula, Shin (2010) provide a theoretical foundation for a funding liquidity channel in a global banking model and show that USD funding liquidity forecasts exchange rates because of its association with time-varying risk premia. Our paper focuses on the other side of the mirror – the USD denominated assets of non-U.S. banks rather than the USD denominated liabilities of the U.S. banking sector. As the marginal investors in the FX market especially for the corresponding foreign currency (by regarding USD as numéraire), our paper demonstrates non-U.S. banks' leverage significantly accounts for the changes in their relative demand for USD denominated assets. It suggests exchange rate fluctuations are associated with the USD asset demand from non-U.S. banks, which is a function of non-U.S. banks' balance sheet capacity. Furthermore, we contribute to cross-sectional asset pricing theory by finding evidence that average USD asset shares of non-U.S. banks can help explain the variation in currency excess returns.

VIII. Conclusion

In this paper, we construct a measure of relative demand for USD denominated assets, proxied as the USD asset share of non-U.S. banks, to explain exchange rate movements. First, we use this variable to explain contemporaneous exchange rate movements. Second, we establish a causal relationship between the USD asset demand, the exchange rate, the U.S. Treasury premium, and the LIBOR cross-currency basis. We distinguish the "safe asset demand channel" from the "financial intermediation channel" with our instrumental variable strategy. Third, we find that the USD asset demand of non-U.S. banks at the global level can help explain currency excess returns in the cross section. Fourth, the USD asset demand variable helps forecast exchange rates both in-sample and out-of-sample. For all of these four empirical exercises, we obtain statistically highly significant and economically large results. More importantly, the four empirical results taken together point to a coherent economic mechanism that significantly advances the exchange rate literature.

We find that the demand for the USD by non-U.S. banks drives down the value of corresponding foreign currency against the U.S. dollar contemporaneously with high statistical significance and in an economically large magnitude. This movement in foreign currency against the U.S. dollar reverts over the next five years. The reversal of the currency value leads to strong forecastability in exchange rates – a positive shock to the USD asset share by one-standard deviation around 0.32 percent (from Table 1) is associated with appreciation of the USD by 8.2 percent per annum on average in that same year (Column (5) of Table 3), and forecasts a depreciation of the USD by an average of 0.70 percent per annum over the next five years (Column (5) of Table 7). Importantly, the forecasting R^2s are very large, with in sample R^2s of 40 percent at the two-year horizon and 68 percent at the five-year horizon. This is an overshooting result in the spirit of Dornbusch (1976): a USD-denominated asset demand shock by non-U.S. banks drives down the exchange rate of foreign currency against the U.S. dollar contemporaneously via a compression of the risk premia, and that compression then reverts over the coming five years. The economic interpretation is one relying on relative scarcity of USD denominated assets to assets denominated in corresponding foreign currency that is driven largely by non-U.S. banks as the marginal investor on the FX market, especially for the foreign currency.

The in-sample forecasting results are strongly suggestive of a risk premium mechanism. Time varying FX risk premia are related to the relative demand of USD denominated assets from non-U.S. banks in a systematic fashion, thus creating forcastability of the exchange rate. Indeed, in the cross-sectional Fama-MacBeth regressions, we find that currencies and carry-trade portfolios earn significant excess returns with respect to higher loadings on the global USD asset demand factor. Put differently, the price of risk of the USD asset demand factor is negative, and relatively high interest rate currencies (or high carry portfolios) have more negative exposure to the USD asset demand factor, thus earning a positive and large excess return on average.

Importantly, the strength of the out-of-sample forecasting power from a quantity variable is unprecedented in the literature on exchange rates. We clearly reject the random walk hypothesis, thus again confirming that a time varying risk premium embedded in the USD asset demand of non-U.S. banks is the lead explanation for our findings. Importantly, we significantly improve relative to 9 traditional exchange rate forecasting models that have been presented in the literature since the 1970s, suggesting that our USD asset demand channel in forecasting exchange rates is very robust.

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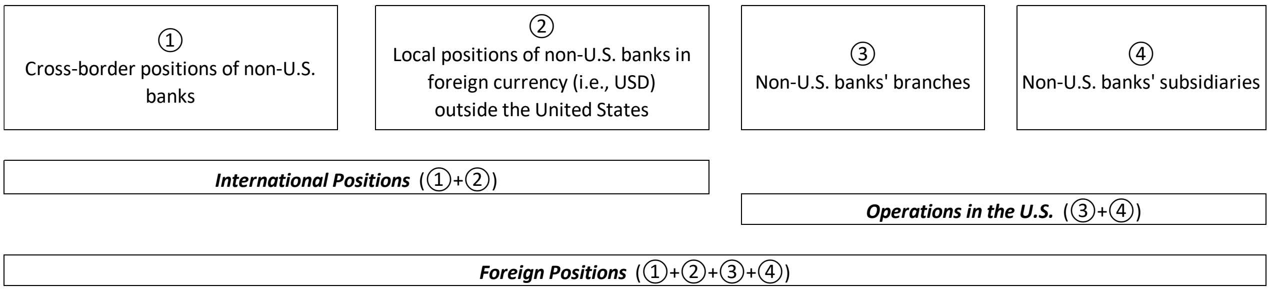
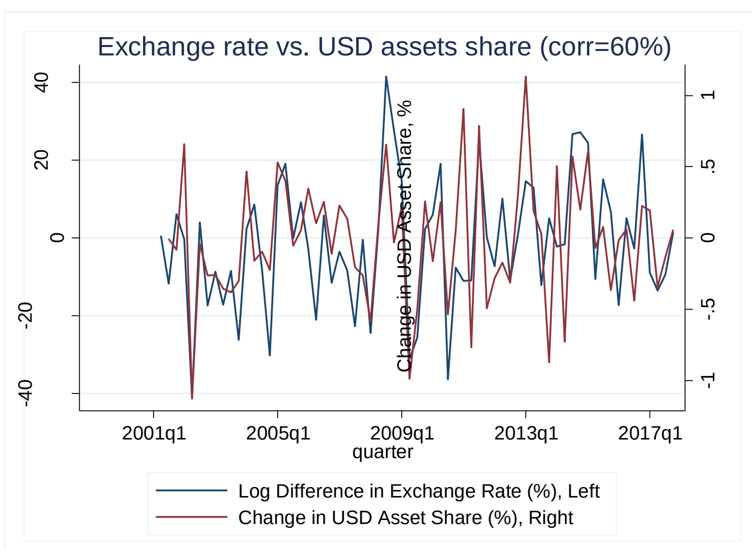
Figures and Tables

FIGURE 1. USD INTERNATIONAL POSITIONS vs. FOREIGN POSITIONS



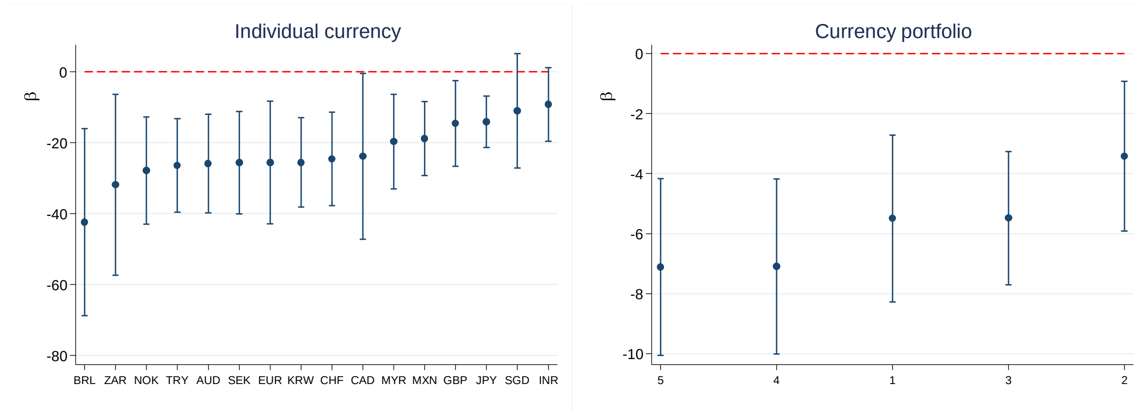
Notes: ① corresponds to the U.S. dollar cross-border claims; ② refers to U.S. dollar local positions outside the U.S. (i.e., U.S. dollar is the foreign currency of the economy); ③ and ④ refer to the local positions in U.S. dollar in the U.S for non-U.S. banks' branches and subsidiaries. ① and ② are constructed based on the BIS unpublished free and restricted database of Locational Banking Statistics (LBS) on nationality basis. Due to data limitations, ③ is constructed through a bottom-up aggregation based on nationality of balance sheet information available in FFIEC 002 Regulatory Filings. Similarly, ④ is constructed using data from the Call Report Filings (FFIEC 031/041) of non-U.S.-owned subsidiaries, obtained via the S&P Global Market Intelligence platform. The underlying assumption is that the operations of non-U.S. banks in the U.S. are dominated by USD-denominated claims and liabilities.

FIGURE 2. U.S. DOLLAR ASSET SHARE AND EXCHANGE RATE



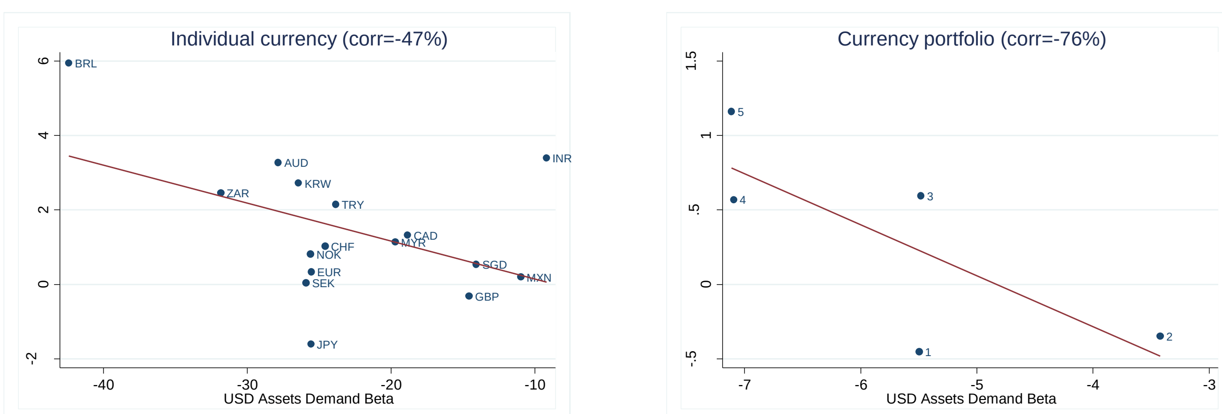
Notes: The figure plots the annualized quarterly changes in log nominal exchange rate and quarterly changes in U.S. dollar asset share of non-U.S. banks over time. All series are cross-sectional averages over non-U.S. G10 currencies and valued at quarter end. An increase in the exchange rate means the U.S. dollar appreciates against a basket of non-U.S. G10 currencies. The contemporaneous correlation between the quarterly changes in USD exchange rate and quarterly changes in U.S. dollar asset share of non-U.S. banks is 60% over time. In the non-USD G10 currencies sample, the USD is excluded as the numéraire and NZD is excluded because of data availability for the USD asset share of non-U.S. banks from the Bank of International Settlements (BIS). The sample period is from 2001q2 to 2017q4.

FIGURE 3. USD ASSET DEMAND LOADING



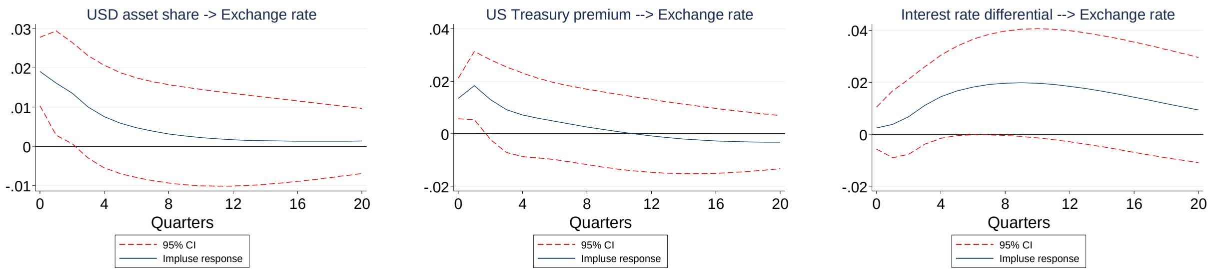
Notes: This figure reports the coefficient point estimates of $\Delta D_{\$,t}$ in the following regression specification: $R_{j,t} = \alpha_j + \beta_j \Delta D_{\$,t} + \gamma_j \Delta USD_t + \varepsilon_{j,t}$, where $R_{j,t}$ is the currency excess return of individual currency J from investing in 1-year government bonds in jurisdiction j and shorting 1-year U.S. Treasury Bill at end of quarter t (in the left Panel), or stands for that of currency portfolio (in the right Panel). $\Delta D_{\$,t}$ is the average quarterly changes in the USD asset share of non-U.S. banks across 16 jurisdictions at end of quarter t. ΔUSD_t is the quarterly changes in logarithm of nominal effective exchange rate (NEER) of USD at the end of quarter t. The blue dots indicate the coefficient point estimates of $\widehat{\beta_j}$ for individual currencies (in the left panel) and for currency portfolios (in the right panel), respectively. The blue bars indicate 95% confidence interval bands. The excess returns of currency portfolios are from Lustig, Roussanov, and Verdelhan (2011). As Lustig and Verdelhan (2007) suggest, we drop the last currency portfolio, which consists currencies with very high inflation. The sample period is from 2001q2 to 2017q4.

FIGURE 4. CURRENCY EXCESS RETURNS AND USD ASSET DEMAND BETAS



Notes: The vertical axis in the left panel shows the average excess return of currency J average over time, defined as investing in 1-year government bonds in jurisdiction j from shorting the 1-year U.S. Treasury, while the horizontal axis indicates the regression beta of running a quarterly regression of currency excess returns on average changes in the USD asset demand of non-U.S. banks for 16 currencies. The vertical axis in the right panel shows the average excess return of currency portfolios over time, while the horizontal axis indicates the regression beta of running quarterly regression of currency excess returns on average changes in USD assets demand of non-U.S. banks for 5 currency portfolios from Lustig, Roussanov, and Verdelhan (2011). As Lustig and Verdelhan (2007), we drop the last currency portfolio, which consists of currencies with very high inflation. The sample period is from 2001q2 to 2017q4.

FIGURE 5. IMPULSE RESPONSES FROM A VAR



Notes: This figure presents impulse responses from a four-variable VAR in the USD asset share, the U.S. Treasury premium, the interest rate differential (home – U.S.), and the exchange rate. All series are cross-sectional average over 16 economies and valued at quarter end. An increase in the exchange rate means that the U.S. dollar appreciates against a basket of 16 currencies. The left panel presents the exchange rate response to a USD-asset-share shock. The middle panel shows the exchange rate response to a U.S. Treasury premium shock. The right panel shows the exchange rate response to an interest rate differential shock. The VAR is estimated in levels and the shocks are identified using a Cholesky decomposition with the following ordering: exchange rate, U.S. Treasury premium, interest rate differential, USD assets share. Red dashed lines represent 95% confidence intervals and blue solid lines represent impulse response function. The sample period is from 2001q2 to 2017q4.

TABLE 1—SAMPLE COVERAGE AND SUMMARY STATISTICS

		(1) Australia	(2) Brazil	(3) Canada	(4) Switzerland	(5) Euro Area	(6) UK	(7) India	(8) Japan	(9) Korea	(10) Mexico	(11) Malaysia	(12) Norway	(13) Singapore	(14) Sweden	(15) Turkey	(16) South Africa	(17) Average
										Levels								
s_t [per USD]	mean	0.24	0.82	0.17	0.10	-0.21	-0.48	3.93	4.65	6.99	2.56	1.25	1.90	0.34	2.02	0.62	2.30	1.70
	std.	0.19	0.27	0.14	0.18	0.13	0.12	0.16	0.14	0.09	0.19	0.12	0.16	0.10	0.14	0.34	0.27	0.12
$D_{\$t}$ [%]	mean	6.61	3.75	14.59	27.86	8.57	9.03	7.79	8.93	4.19	4.69	4.70	5.19	5.82	10.81	12.54	7.89	9.49
	std.	1.41	0.63	1.72	2.68	1.63	1.73	1.93	2.64	0.35	1.65	1.01	1.91	1.31	2.48	1.40	0.87	1.09
Φ_t [%]	mean	-0.07	1.89	0.16	0.41	0.24	0.03	1.17	0.44	1.39	0.08	0.85	0.11	0.12	0.16	0.56	0.01	0.38
	std.	0.26	0.85	0.20	0.24	0.19	0.20	1.34	0.31	1.08	0.70	0.74	0.23	0.29	0.31	0.87	0.48	0.35
$(y_t - y_{\$t})$ [%]	mean	2.39	9.65	0.43	-0.97	-0.08	0.67	5.09	-1.39	1.73	4.36	2.33	1.10	-0.41	0.15	9.60	5.92	2.19
	std.	1.21	2.31	0.72	1.02	1.09	1.09	2.12	1.45	1.24	1.34	0.63	1.67	0.91	1.43	2.54	0.71	1.01
N. Obs.		67	44	67	67	67	67	65	67	52	62	41	67	56	67	52	34	67
										First differences								
Δs_t [%]	mean	-2.56	4.42	-1.13	-3.69	-2.10	0.28	1.75	-0.61	0.39	4.30	2.00	-0.78	-1.64	-1.71	8.11	6.06	0.13
	std.	25.06	35.30	17.61	18.99	20.38	19.22	14.89	22.40	19.62	21.10	17.43	24.23	11.62	23.03	26.35	23.97	15.37
$\Delta D_{\$t}$ [%]	mean	-0.06	-0.03	0.06	0.02	-0.07	-0.05	0.04	0.10	0.00	0.01	0.01	0.00	0.07	-0.13	-0.09	0.01	-0.01
	std.	0.61	0.47	0.68	1.33	0.50	0.47	0.39	0.41	0.28	0.96	0.52	2.12	0.31	0.93	0.88	0.64	0.32
$\Delta\Phi_t$ [%]	mean	0.00	-0.01	0.00	0.00	-0.01	0.00	0.02	0.00	-0.01	0.01	-0.02	0.00	0.00	-0.01	-0.01	0.01	0.00
	std.	0.26	0.77	0.15	0.24	0.17	0.23	1.06	0.31	0.64	0.66	0.63	0.22	0.33	0.29	0.64	0.37	0.24
$\Delta(y_t - y_{\$t})$ [%]	mean	-0.01	-0.04	-0.01	-0.03	-0.04	-0.04	-0.01	0.03	0.00	-0.04	0.03	-0.07	0.00	-0.05	-0.03	-0.05	-0.02
	std.	0.47	1.06	0.30	0.40	0.38	0.39	0.59	0.43	0.41	0.67	0.44	0.52	0.39	0.48	1.55	0.40	0.38
N. Obs.		66	43	66	66	66	66	64	66	51	61	40	66	55	66	51	33	66

Notes: This table describes the sample of 16 economies (based on nationalities of ultimate owners of reporting non-U.S. banks) used in this paper. It also presents the summary statistics of key variables. Columns (1) - (16) report by economies, and Column (17) reports the cross-sectional average over all 16 economies. In the upper (lower) panel, we report the summary statistics in levels (first differences) for key variables, including the log nominal exchange rate per U.S. dollar (s_t); the USD asset share of non-U.S. banks ($D_{\$t}$); the U.S. Treasury premium (Φ_t); home – U.S. government bond yield differential ($y_t - y_{\$t}$). All series are valued at quarter end. The longest sample covers 2001q1 to 2017q4 for the upper panel and 2001q2 to 2017 q4 for the lower panel.

TABLE 2—CORRELATION BETWEEN EXCHANGE RATE AND DETERMINANTS

Panel A: Correlation with Contemporaneous Changes				
	Δs_t	$\Delta D_{\$t}$	$\Delta\Phi_t$	$\Delta(y_t\text{-}y_{\$t})$
Δs_t	1.00***			
$\Delta D_{\$t}$	0.60***	1.00***		
$\Delta\Phi_t$	0.36***	0.28**	1.00***	
$\Delta(y_t\text{-}y_{\$t})$	-0.32***	-0.28**	-0.18	1.00***
Panel B: Correlation with Future Changes				
	Δs^{12}_t	$D_{\$t}$	Φ_t	$y_t\text{-}y_{\$t}$
Δs^{12}_t	1.00***			
$D_{\$t}$	-0.49***	1.00***		
Φ_t	-0.29**	0.33***	1.00***	
$y_t\text{-}y_{\$t}$	0.07	-0.26**	-0.19	1.00***

Notes: The upper panel shows the contemporaneous correlations between annualized quarterly changes in the log nominal exchange rate (Δs_t) and quarterly changes in selected determinants. The lower panel shows the correlations between annualized future changes in the log nominal exchange rate between t and t+12 quarters ahead (Δs_t^{12}) and selected determinants. The variable $D_\$$ denotes the USD asset share of non-U.S. banks. The variable Φ denotes the U.S. Treasury premium. The variable $(y - y_\$)$ denotes home – U.S. government bond yield differentials. All series are cross-sectional averages over non-USD G10 currencies and valued at quarter end. The USD is excluded as the numéraire, and NZD is excluded because of data availability for the USD asset share of non-U.S. banks from the Bank of International Settlements (BIS). The sample period is from 2001q2 to 2017q4.

TABLE 3—USD ASSETS SHARE AND EXCHANGE RATES

Regressors	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)	(23)
	Dependent variable = Δs_t [Annualized %]																						
	Cross-sectional Average							Bilateral Relationship															
	Full Sample					1st half	2nd half	AUD	BRL	CAD	CHF	EUR	GBP	INR	JPY	KRW	MXN	MYR	NOK	SEK	SGD	TRY	ZAR
$\Delta D_{\$t}$ [%]	28.82***	26.45***	28.60***	25.86***	25.53***	27.20***	25.33***	25.17***	65.75***	10.57***	5.89***	20.05***	-3.03	7.23	31.90***	23.19*	8.08***	14.66**	1.11	4.77**	3.30	13.98***	17.09
	(5.57)	(5.32)	(5.60)	(5.41)	(5.26)	(7.03)	(5.62)	(4.07)	(6.69)	(2.53)	(1.48)	(3.71)	(2.33)	(5.71)	(6.48)	(12.45)	(3.01)	(6.92)	(0.81)	(1.97)	(4.15)	(4.42)	(10.25
$\Delta\Phi_t$ [%]		12.48*		13.56**	21.90***	19.39***	29.77**	18.89**	7.15	31.43**	36.93***	35.38***	22.13***	4.40**	1.50	20.47***	-2.59	19.49***	42.34**	41.67***	5.48	9.43*	11.73
		(6.72)		(6.28)	(6.00)	(5.76)	(11.02)	(9.03)	(4.67)	(12.05)	(9.09)	(11.59)	(7.97)	(1.86)	(4.85)	(4.01)	(6.69)	(4.60)	(21.15)	(8.25)	(8.75)	(5.50)	(17.57
$\Delta(y_t - y_{\$t})$ [%]			-2.26	-3.81	-3.06	-0.32	-14.48	-15.80***	-4.21	-18.42***	-19.02***	-13.26***	-18.90**	-1.54	-12.81***	-11.60**	4.96	-2.72	-18.62***	-15.35***	-1.22	5.55***	2.19
			(3.57)	(3.87)	(3.09)	(2.76)	(9.57)	(4.94)	(2.82)	(5.61)	(6.25)	(4.89)	(9.40)	(2.55)	(3.88)	(5.00)	(3.93)	(6.50)	(6.86)	(4.66)	(5.39)	(1.95)	(13.73
$D_{\$t-1}$ [%]					2.11	3.11***	9.45***	2.39*	3.50	2.96***	0.83	-2.24**	-3.68***	0.89	0.81	8.10	1.39	3.82	2.66**	-2.52***	1.37	5.01**	6.87
					(1.36)	(1.07)	(3.16)	(1.33)	(6.02)	(0.92)	(0.71)	(1.00)	(1.22)	(1.69)	(0.71)	(6.48)	(1.80)	(2.95)	(1.28)	(0.59)	(1.89)	(2.30)	(5.68)
Φ_{t-1} [%]					16.39**	25.96***	5.20	9.46	-3.98	25.94***	24.46***	27.37***	25.66	4.15**	-12.07**	4.88**	2.53	11.50***	-0.74	29.73***	6.49	8.76	1.64
					(6.35)	(5.74)	(7.85)	(8.44)	(3.97)	(5.81)	(8.20)	(8.62)	(16.97)	(1.73)	(5.81)	(1.87)	(6.24)	(3.64)	(12.45)	(5.54)	(7.76)	(6.35)	(8.97)
$(y_{t-1} - y_{\$t-1})$ [%]					0.63	-2.08*	4.67	-0.79	1.86	3.21	-0.00	1.37	4.33*	1.14	0.08	-0.28	-0.27	6.60	-1.26	1.87	0.76	-0.69	-10.17*
					(1.76)	(1.15)	(3.81)	(1.57)	(1.47)	(2.65)	(1.96)	(1.27)	(2.51)	(1.65)	(1.03)	(1.68)	(2.47)	(4.06)	(1.43)	(1.26)	(2.45)	(1.67)	(4.56)
Controls	N	N	N	N	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
N	66	66	66	66	66	34	32	66	43	66	66	66	66	64	66	51	61	40	66	66	55	51	33
R^2	0.37	0.40	0.37	0.41	0.52	0.65	0.61	0.60	0.72	0.46	0.45	0.61	0.36	0.26	0.54	0.51	0.18	0.57	0.33	0.50	0.08	0.42	0.29

Notes: The table reports the OLS regression results for the contemporaneous relationship between the USD exchange rate and the USD asset share of non-U.S. banks, with specification: $\Delta s_{j,t} = \alpha_1 + \beta_1 \Delta D_{\$,j,t} + \beta_2 \Delta\Phi_{j,t} + \beta_3 \Delta\left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right) + \beta \boldsymbol{X}_{j,t} + \varepsilon_t$, where the dependent variable $\Delta s_{j,t}$ is the annualized quarterly change in the log nominal exchange rate of the USD vis-à-vis currency J, an increase of which means the USD appreciates against currency J between quarters t-1 and t. The variable $\Delta D_{\$,j,t}$ denotes quarterly changes in the USD asset share of non-U.S. banks from economy j between quarters t-1 and t. The variable $\Delta\Phi_{j,t}$ denotes quarterly changes in the U.S. Treasury premium vis-à-vis economy j between quarters t-1 and t. The variable $\Delta\left(y_{j,t}^{Govt} - y_{\$,t}^{Govt}\right)$ denotes the quarterly changes in the interest differential of economy j over the United States between quarters t-1 and t. Finally, $\boldsymbol{X}_{j,t}$ is a vector of control variables including $D_{\$,j,t-1}$, $\Phi_{j,t-1}$, and $\left(y_{j,t-1}^{Govt} - y_{\$,t-1}^{Govt}\right)$. In Columns (1) to (7), we perform regressions by using cross-sectional average series across 16 economies, within which Columns (1) to (5) uses the full sample, and Columns (6) and (7) use the 1st half of the sample (2001q2 – 2009q4) and the 2nd half of the sample (2010q1 – 2017q4), respectively. In Columns (1) – (7), all series are cross-sectional averages over 16 currencies and valued at quarter end. In Columns (8) – (23), we perform regressions by currencies. All specifications include a constant that is not reported in the table. The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are reported in parentheses. The lag length (m) is set to the integer part of $T^{1/4}$ (i.e., 2) as suggested by Greene (2003). *, **, and *** denote significance levels at 10%, 5%, and 1%. The full sample period is from 2001q2 to 2017q4.

TABLE 4—TWO-STAGE LEAST SQUARES: USD ASSET SHARE AND EXCHANGE RATES

	The first stage: Dependent variable = $\Delta D_{\$t}$ [%]				
Regressors	(1)	(2)	(3)	(4)	(5)
ΔSovereign $CDS^{sub}{}_t$ [bps]	0.01***	0.01***	0.01**	0.01***	0.01***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta\Phi^{sub}{}_t$ [%]	-0.15	-0.75**	-0.03	-0.65*	-0.73**
	(0.38)	(0.36)	(0.36)	(0.33)	(0.36)
$\Delta Leverage_t$ [%]	2.85***	2.54***	3.00***	2.65***	2.60***
	(0.39)	(0.40)	(0.36)	(0.34)	(0.29)
$\Delta\Phi_t$ [%]		0.45***		0.44***	0.33**
		(0.11)		(0.10)	(0.13)
$\Delta(y_t - y_{\$t})$ [%]			-0.10	-0.06	-0.05
			(0.14)	(0.13)	(0.12)
$D_{\$t-1}$ [%]					-0.09
					(0.06)
Φ_{t-1} [%]					-0.23*
					(0.12)
$(y_{t-1} - y_{\$t-1})$ [%]					0.01
					(0.03)
Stock-Yogo *F* Statistics	16.51**	23.85**	19.15**	24.79**	32.43***
N	66	66	66	66	65
R^2	0.19	0.26	0.21	0.27	0.33

	OLS regression	The second stage: Dependent variable = Δs_t [Annualized %]				
Regressors	(0)	(1)	(2)	(3)	(4)	(5)
$\Delta D_{\$t}$ [%] or Fitted $\Delta D_{\$t}$ [%]	28.82***	51.83***	51.15***	51.54***	50.23***	40.60***
	(5.57)	(12.27)	(11.82)	(12.76)	(12.29)	(8.70)
$\Delta\Phi_t$ [%]			4.23		4.81	16.39***
			(7.68)		(7.51)	(6.11)
$\Delta(y_t - y_{\$t})$ [%]				-0.56	-1.13	-1.87
				(4.24)	(4.36)	(4.18)
$D_{\$t-1}$ [%]						2.55*
						(1.49)
Φ_{t-1} [%]						15.82**
						(6.00)
$(y_{t-1} - y_{\$t-1})$ [%]						0.85
						(1.80)
N	66	66	66	66	66	65
Endogeneity: Durbin-Wu-Hausman *F* statistics	-	5.51**	6.16**	4.96**	5.42**	4.56**
Overidentification: Sargan χ^2 Statistics	-	4.01	2.51	4.17	2.58	1.18

Notes: The table reports the 2SLS regression results for the contemporaneous relationship between the USD exchange rate and the USD assets share of non-U.S. banks. The upper panel summarizes the regression results for the first stage: $\Delta\bar{D}_{\$,t} = \alpha_1^1 + \beta_1^1 \Delta\bar{\Phi}_t^{sub} + \beta_2^1 \Delta\overline{Sovereign\ CDS}_t^{sub} + \beta_3^1 \Delta\overline{Leverage}_t + \beta^1 \bar{\boldsymbol{X}}_t + \varepsilon_t^1$, where the variable $\Delta\bar{D}_{\$,t}$ denotes average quarterly changes in USD asset share of non-U.S. banks; $\Delta\bar{\Phi}_t^{sub}$ denotes the average quarterly changes in Treasury premium of substitute government bonds; $\Delta\overline{Sovereign\ CDS}_t^{sub}$ denotes the average quarterly changes in sovereign CDS spread of substitute government bonds; $\Delta\overline{Leverage}$ denotes the average quarterly changes of the log leverage ratio of non-U.S. banks, orthogonalized to U.S. bank leverage; Finally, $\boldsymbol{X}_t$ is a vector of control variables, including $\bar{D}_{\$,t-1}$, $\bar{\Phi}_{t-1}$, and $(\bar{y}_{t-1}^{Govt} - y_{\$,t-1}^{Govt})$. The lower panel summarizes the regression results for the second stage: $\Delta\bar{s}_t = \alpha_1^2 + \beta_1^2 \Delta\hat{\bar{D}}_{\$,t} + \beta_2^2 \Delta\bar{\Phi}_t + \beta_3^2 \Delta(\bar{y}_t^{Govt} - y_{\$,t}^{Govt}) + \beta^2 \bar{\boldsymbol{X}}_t + \varepsilon_t^2$, where the variable $\Delta\hat{\bar{D}}_{\$,t}$ denotes fitted value of average quarterly changes in USD assets share of non-U.S. banks obtained from the first stage. Other variables are defined the same way as those in Table 3. As a benchmark, the Column (0) of lower panel presents the OLS regression by using the original value of $\Delta D_{\$t}$, which is the result presented in Column (1) of Table 3, instead of fitted value from the first stage. The constants are not reported in the table for both stages. For testing for weak instrumental variables, F-statistics are reported. According to Stock and Yogo (2002), the critical value for the weak instrument test based on 2SLS bias of 5% significant level is 13.91 for parameters setting with 1 endogenous regressor, 3 instrumental variables, with the desired maximal bias of the IV estimator relative to OLS at 0.05. For endogeneity, the Durbin-Wu-Hausman F statistics are reported. For overidentifying restriction, the Sargan χ^2 statistic is reported. All series are cross-sectional averages over 16 currencies and valued at quarter end. The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are reported in parentheses. The lag length (m) is set to the integer part of $T^{1/4}$ (i.e., 2) as suggested by Greene (2003). *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is from 2001q2 to 2017q4.

TABLE 5—INSTRUMENTAL VARIABLE STRATEGY: VALIDATIONS

Regressors	Benchmark	The first stage: Dependent variable = $\Delta D_{\$t}$ [%]						
	(0)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
		Full Sample						
ΔSovereign CDS^{sub}_t [bps]	0.01***	0.01**	0.01***	0.01***	0.01***	0.01***	0.01***	0.01***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\Delta\Phi^{sub}_t$ [%]	-0.65*	-0.65*	-0.79**	-0.66*	-0.68**	-0.70**	-0.72*	-0.96**
	(0.33)	(0.34)	(0.37)	(0.33)	(0.33)	(0.33)	(0.37)	(0.43)
$\Delta Leverage_t$ [%]	2.65***	2.78***	2.49***	2.65***	2.62***	2.70***	2.66***	2.52***
	(0.34)	(0.45)	(0.35)	(0.34)	(0.39)	(0.39)	(0.34)	(0.58)
$\Delta\Phi_t$ [%]	0.44***	0.41***	0.51***	0.43***	0.45***	0.41***	0.44***	0.48***
	(0.10)	(0.10)	(0.10)	(0.10)	(0.12)	(0.10)	(0.10)	(0.12)
$\Delta(y_t\text{-}y_{\$t})$ [%]	-0.06	-0.08	-0.05	-0.06	-0.04	-0.09	-0.02	-0.00
	(0.13)	(0.12)	(0.12)	(0.13)	(0.17)	(0.12)	(0.13)	(0.14)
Δlog(General Govt Debt/GDP of USA)$_t$		1.36						-0.20
		(2.23)						(3.18)
Δlog(General Govt Debt/GDP of Substitute Economies)$_t$		-2.56						-0.63
		(2.93)						(3.56)
ΔUSD Funding Gap Ratio [%]			-0.02					-0.02
			(0.02)					(0.02)
FCI Global, IMF				-0.01				-0.02
				(0.02)				(0.04)
Consensus forecast growth of USA [%]					0.03			0.03
					(0.07)			(0.06)
Consensu forecast growth of Global [%]					-0.05			-0.06
					(0.07)			(0.10)
Δlog(VIX)						0.14		0.18
						(0.13)		(0.15)
Δ1y carry between Global Avg. and the U.S. [%]							-0.03	-0.02
							(0.04)	(0.04)
Stock-Yogo *F* Statistics	24.79**	20.99**	26.38**	24.76**	23.37**	20.91**	25.66**	22.79**
N	66	66	65	66	66	66	66	65
R^2	0.27	0.27	0.31	0.27	0.27	0.28	0.27	0.34

Regressors	Benchmark	The second stage: Dependent variable = Δs_t [Annualized %]						
	(0)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
		Full Sample						
$\Delta D_{\$t}$ [%] or Fitted $\Delta D_{\$t}$ [%]	50.23***	46.20***	49.28***	41.46***	49.48***	46.24***	49.19***	40.30***
	(12.29)	(11.30)	(11.65)	(10.43)	(10.80)	(9.60)	(12.20)	(6.94)
$\Delta\Phi_t$ [%]	4.81	5.65	4.14	10.78*	5.41	2.46	6.28	5.75
	(7.51)	(6.34)	(7.67)	(5.71)	(7.36)	(6.42)	(7.45)	(5.56)
$\Delta(y_t\text{-}y_{\$t})$ [%]	-1.13	-1.63	-1.87	-1.60	-1.95	-3.74	-3.49	-7.85
	(4.36)	(4.22)	(4.50)	(3.97)	(3.93)	(4.46)	(4.66)	(5.14)
Δlog(General Govt Debt/GDP of USA)$_t$		116.37						93.36
		(106.68)						(111.05)
Δlog(General Govt Debt/GDP of Substitute Economies)$_t$		47.93						64.93
		(149.36)						(132.36)
ΔUSD Funding Gap Ratio [%]			0.93*					0.78
			(0.50)					(0.55)
FCI Global, IMF				4.11***				0.57
				(0.83)				(1.79)
Consensus forecast growth of USA [%]					0.42			0.65
					(2.30)			(2.38)
Consensu forecast growth of Global [%]					-8.41***			-7.88*
					(3.16)			(4.25)
Δlog(VIX)						13.56*		12.91
						(7.44)		(8.25)
Δ1y carry between Global Avg. and the U.S. [%]							1.70	2.33*
							(1.44)	(1.17)
N	66	66	65	66	66	66	66	65
Endogeneity: Durbin-Wu-Hausman *F* statistics	5.42**	5.20**	5.40**	4.46**	6.59**	6.77**	5.14**	7.65***
Overidentification: Sargan χ^2 Statistics	2.58	2.52	2.6	1.65	3.13	1.90	2.86	1.82

Notes: The table reports the 2SLS regression results for the contemporaneous relationship between the USD exchange rate and the USD assets share of non-U.S. banks with additional controls to eliminate potential concerns in validity of the proposed IVs. The specification is built upon Column (4) of Table 4. The general government debt to GDP ratio on a consolidated basis for the U.S. and the other substitute economies is obtained from Haver Analytics. Following Barajas et al. (2020), we construct the USD funding gap ratio to proxy the capacity of non-U.S. banks to provide the USD denominated assets. The Financial Condition Index (FCI) is from IMF Monetary and Capital Market Department. VIX is the Chicago Board Options Exchange's CBOE Volatility Index. The 1-year carry between global average and the U.S. is measured as the average of 1-year forward premium between 16 economies against the USD. All other settings are as the same as Table 4. *, **, and *** denote significance levels at 10%, 5%, and 1%. The full sample period is from 2001q2 to 2017q4.

TABLE 6—USD ASSET SHARE, U.S. TREASURY PREMIUM, and CROSS-CURRENCY BASIS

	The Second Stage			
	(1)	(2)	(3)	(4)
Dependent variables	$\Delta\Phi^{Treasury}_t$		$\Delta\Phi^{Libor}_t$	
Sample periods	Full	Post-GFC	Full	Post-GFC
Fitted $\Delta D_{\$t}$ [%]	0.46***	0.61**	0.29	0.21**
	(0.15)	(0.26)	(0.18)	(0.11)
Controls	N	N	N	N
N	66	32	66	32
R^2	0.16	0.14	0.07	0.20

Notes: The table summarizes the second-stage results of the 2SLS regression of the U.S. Treasury premium on the USD asset share in Columns (1) to (2), and of the (-1)* cross-currency basis on USD asset share in Columns (3) and (4) with the specification: $\Delta\overline{\Phi}_t = \alpha_1^2 + \beta_1^2 \Delta\widehat{\overline{D}}_{\$,t} + \beta_2^2 \Delta\left(\bar{y}_t^{Govt} - y_{\$,t}^{Govt}\right) + \beta^2 \overline{\boldsymbol{X}}_t + \varepsilon_t^2$, where $\Delta\overline{\Phi}_t$ is the average quarterly changes in U.S. Treasury premium in Columns (1) and (2), or average quarterly changes in (-1)*cross-currency basis in Columns (3) and (4), respectively. The variable $\Delta\widehat{\overline{D}}_{\$,t}$ denotes the fitted value of average quarterly changes in USD asset share of non-U.S. banks obtained from the first stage. Other variables are defined as the same as those in Table 3. All specifications include a constant that is not reported in the table. All series are cross-sectional averages over 16 currencies and valued at quarter end. The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are reported in parentheses. *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is from 2001q2 to 2017q4.

TABLE 7—FORCASTING EXCHANGE RATES

	Dependent variable = Δs^h_t [Annualized %]				
	(1)	(2)	(3)	(4)	(5)
Regressors	*h* =1	*h* =4	*h* =8	*h* =12	*h* =20
$D_{\$t}$ [%]	-0.76	-2.79**	-3.17***	-2.68***	-2.19***
	(1.84)	(1.38)	(0.75)	(0.57)	(0.49)
Φ_t [%]	3.59	-5.64	-6.02*	-1.93	0.15
	(8.59)	(4.68)	(3.01)	(2.03)	(1.83)
$(y_t - y_{\$t})$ [%]	2.86	3.07**	1.98*	0.81	0.73
	(1.97)	(1.52)	(1.01)	(0.72)	(0.51)
N	66	63	59	55	47
R^2	0.07	0.25	0.40	0.45	0.68

Notes: This table reports OLS results for in-sample forecasts of annualized percentage changes in the U.S. dollar exchange rates against a basket of 16 currencies at the one-quarter, one-year, two-year, three-year and five-year horizons. The specification is: $\Delta s^h_{j,t} = \alpha_1 + \beta_1 D_{j,\$,t} + \beta_2 \Phi_{j,t} + \beta_3 \left(y^{Govt}_{j,t} - y^{Govt}_{\$,t}\right) + \varepsilon_{t+h}$, where $\Delta s^h_{j,t}$ is the annualized future change in the log of nominal exchange rate of the USD vis-à-vis currency J in economy j over the horizon $[t, t+h]$, an increase of which means that the USD appreciates against currency J over the period $[t, t+h]$. The variable $D_{j,\$,t}$ denotes the USD asset share by non-U.S. banks from economy j at end of quarter t. The variable $\Phi_{j,t}$ denotes the U.S. Treasury premium vis-à-vis economy j at end of quarter t. The variable $\left(y^{Govt}_{j,t} - y^{Govt}_{\$,t}\right)$ denotes the interest differential at end of quarter t. We run this regression using quarterly data as before but vary the horizons ahead from one-quarter ($h = 1$), one-year ($h = 4$), two-years ($h = 8$), three-years ($h = 12$), to five-years ($h = 20$). All series are cross-sectional averages over 16 currencies and valued at quarter end. All specifications include a constant (not reported). The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are reported in parentheses. The lag length (*m*) is set to the integer part of $T^{1/4}$ (i.e., 2) as suggested by Greene (2003). *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is 2001q2 to 2017q4.

TABLE 8—FORCASTING DOLLAR EXCHANGE RATES AGAINST INIDIVIDUAL CURRENCY

	Dependent variable = Δs^{20}_{jt} [Annualized %]															
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
	AUD	BRL	CAD	CHF	EUR	GBP	INR	JPY	KRW	MXN	MYR	NOK	SEK	SGD	TRY	ZAR
	Panel A: Individual USD Asset Share															
$D_{j\$t}$ [%]	-1.57***	-4.58***	0.75	-0.27	-1.49***	-0.97	1.92***	1.26*	-3.55***	-0.76***	2.30**	1.21**	-1.30***	2.21***	-1.54***	-1.89**
	(0.39)	(1.12)	(0.54)	(0.20)	(0.43)	(0.61)	(0.50)	(0.65)	(1.15)	(0.26)	(0.85)	(0.59)	(0.16)	(0.29)	(0.18)	(0.84)
N	47	43	47	47	47	47	45	47	32	43	21	47	47	36	47	14
R^2	0.22	0.58	0.05	0.08	0.29	0.11	0.53	0.11	0.31	0.27	0.32	0.08	0.74	0.63	0.77	0.33
	Panel B: Cross-Sectional Average USD Asset Share															
$D_{\$t}$ [%]	-3.02***	-4.35**	-2.76***	-1.45***	-2.88***	-1.94***	-2.11***	-1.69**	-1.47***	-0.65	-5.60***	-3.28***	-2.74***	-0.98**	-3.31***	-3.94**
	(0.60)	(1.66)	(0.55)	(0.30)	(0.33)	(0.38)	(0.44)	(0.78)	(0.37)	(0.42)	(1.00)	(0.51)	(0.31)	(0.39)	(0.97)	(0.66)
N	47	47	47	47	47	47	47	47	47	47	30	47	47	47	47	47
R^2	0.53	0.40	0.64	0.42	0.81	0.46	0.59	0.18	0.31	0.10	0.57	0.69	0.68	0.28	0.52	0.64

Notes: This table reports OLS regression results for in-sample forecasts of annualized percentage changes in the bilateral USD exchange rate at the five-year horizon. The univariate regression specification is: $\Delta s^{20}_{j,t} = \alpha_1 + \beta_1 D_{j,\$,t} + \varepsilon_{t+20}$, where $\Delta s^{20}_{j,t}$ is the annualized change in the log bilateral USD exchange rate against currency J in economy j over the horizon of $[t, t+20]$, an increase of which means USD appreciates against currency J. The variable $D_{j,\$,t}$ denotes USD asset share of non-U.S. banks from economy j at end of quarter t. In Panel A, the regressor is the USD asset share of non-U.S. banks from individual economy. In Panel B, the regressor is the cross-sectional average over the 16 currencies. All specifications include a constant (not reported). The Newey-West heteroskedasticity-and-autocorrelation-consistent asymptotic standard errors are reported in parentheses. The lag length (*m*) is set to the integer part of $T^{1/4}$ (i.e., 2) as suggested by Greene (2003). *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is 2001q2 to 2017q4.

TABLE 9—OUT-OF-SAMPLE FORECASTABILITY

	Dependent variable = Δs^{20}_{t} [Annualized %]																
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)
	ALL	AUD	BRL	CAD	CHF	EUR	GBP	INR	JPY	KRW	MXN	MYR	NOK	SEK	SGD	TRY	ZAR
	Panel A: Diebold-Mariano Test																
$MSE_r - MSE_u$	14.07	11.24	63.34	8.86	4.00	6.59	4.34	28.41	14.81	-3.34	19.55	8.73	21.45	6.82	2.36	98.55	72.43
OOS-T statistics	2.47***	3.96***	1.79***	2.40***	0.92*	2.40***	0.87*	4.42***	1.44*	-1.02	2.38***	2.20***	2.52***	2.56***	0.95*	4.33***	2.69**
	Panel B: Clark-West Test																
$MSE_r - (MSE_u - Adj.)$	20.81	19.91	113.69	13.74	13.12	14.16	14.42	55.52	35.49	2.42	30.83	14.07	31.57	9.54	9.11	164.02	118.84
C-W statistics	2.57***	5.88***	2.38***	3.07***	1.85***	3.35***	1.99***	5.14***	3.24***	0.60	3.17***	3.30***	2.60***	2.75***	1.89***	4.48***	3.01**

Notes: This table investigates the out-of-sample forecastability of the U.S. dollar exchange rates at the five-year horizon. We compare the forecasting performance of a model using the USD assets share as predictor against a benchmark random walk model. To estimate model parameters, we regress the annualized percentage changes in the USD exchange rate on the USD asset share to provide in-sample estimates in the period of $[T-P-K, T-P]$ with the specification: $\Delta s^{20}_{j,t} = \alpha_1 + \beta_1 D_{\$,t} + \varepsilon_{t+20}$, where Δs^{20}_{t} is the annualized percentage change in the USD exchange rate against currency J between quarters t and t+20, increases of which means that the USD appreciates against currency J over the horizon $[t, t+20]$. The variable $D_{\$,t}$ is the cross-sectional average of the USD assets share of non-U.S. banks over 16 economies at quarter t. The estimated coefficients and the realized observations of the USD asset share are used to forecast the exchange rate change in next period. We repeat this procedure P times using a rolling regression approach until we have an out-of-sample forecast for all quarters in the period between $T-P$ to T. We compute the mean-square-error (MSE) of our forecasting model and compare it against that of the random walk benchmark. Panel A reports the Diebold-Mariano difference in MSEs and the OOS-T statistic. Panel B reports the Clark-West adjusted difference in MSEs and the C-W statistic. The heteroskedasticity and autocorrelation consistent (HAC) standard errors are estimated to compute the OOS-T statistics and C-W statistics (Clark and West 2006). The corresponding 90%, 95%, and 99% critical values are 0.780, 1.111, and 1.784 (using $k_2 = 1$ and $\pi = 0.2$) obtained from Table 1 in McCracken (2007). In Column (1), the USD exchange rate against the cross-sectional average over 16 currencies is used. From Columns (2) – Column (17), the USD exchange rate against individual currency is used. All series are valued at quarter end. *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is 2001q2 to 2017q4.

TABLE 10—ROBUSTNESS OF OUT-OF-SAMPLE FORECASTABILITY

Name of Models	$MSE_{existing} - MSE_{augmented}$	OSS-T statistics
(1) UIP model	10.35	3.20***
(2) Monetary model with flexbile prices (Frankel-Bilson model)	10.75	3.61***
(3) Monetary model with sticky prices (Dornbusch-Frankel model)	7.48	2.27***
(4) Productivity differentials model (Balassa-Samuelson model)	6.73	1.96***
(5) Taylor rule model	4.40	4.35***
(6) Net foreign asset model (Gourichas-Rey model)	1.20	2.30***
(7) U.S. dollar liquidity model (Adrian-Etula-Shin model)	0.61	0.87*
(8) U.S. Treasury premium model (Jiang-Krishnamurthy-Lustig model)	0.69	0.87*
(9) U.S. foreign bond flow model (Lilley-Neiman-Maggiori-Schreger model)	6.60	2.47***

Notes: This table investigates the robustness of the out-of-sample forecasting performance of the USD asset share in predicting U.S. dollar exchange rates against a basket of 16 currencies at the 5-year horizon. We compare the performance of 9 existing exchange rate models against augmented versions including the USD asset share as an additional predictor. Each row represents one performance comparison between the augmented model and the existing model. We report the Diebold-Mariano difference in MSEs and the OOS-T statistics. The heteroskedasticity and autocorrelation consistent (HAC) standard errors are estimated to compute the OOS-T statistics (Clark and West 2006). The corresponding 90%, 95%, and 99% critical values are 0.780, 1.111, and 1.784 (using $k_2 = 1$ and $\pi = 0.2$) obtained from Table 1 in McCracken (2007). All series are the cross-sectional average over 16 currencies and valued at quarter end. *, **, and *** denote significance levels at 10%, 5%, and 1%. The sample period is 2001q2 to 2017q4.