Foreign Safe Asset Demand and the Dollar Exchange Rate*

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Abstract

We develop a theory that links the U.S. dollar's valuation in FX markets to foreign investors' demand for U.S. safe assets. When the convenience yield that foreign investors derive from holding U.S. safe assets increases, the U.S. dollar immediately appreciates, thus lowering the foreign investors' expected future return from owning U.S. safe assets. The foreign investors' convenience yield can be inferred from the wedge between the yield on safe U.S. Treasury bonds and currency-hedged foreign government bonds, which we call the U.S. Treasury basis. Consistent with the theory, we find that a widening of the U.S. Treasury basis coincides with an immediate appreciation and a subsequent depreciation of the U.S. dollar. Shocks to news about current and future convenience yields accounts for 54.2% of the quarterly innovations in the dollar. Our results lend empirical support to recent theories of exchange rate determination which impute a special role to the U.S. as the world's provider of safe assets and to the dollar, the world's reserve currency.

Keywords: Covered interest rate parity, exchange rates, safe asset demand, convenience yields.

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In the post-war era, the U.S. has been the world's most favored supplier of safe assets and investors pay a sizeable premium, the convenience yield, to own these assets (see Krishnamurthy and Vissing-Jorgensen, 2012, for example). Moreover, during episodes of global financial instability, there is a flight to the safety of U.S. Treasury bonds which increases their convenience yield. At the same time, the value of the dollar in foreign exchange markets rises. This paper develops and presents empirical evidence in support of a theory that connects these observations. Under our new convenience yield theory of exchange rates, the dollar's valuation depends on the convenience yields that foreign investors derive from the ownership of U.S. safe assets. Empirically, we show that news about current and future convenience yields accounts for 54.2% of the quarterly innovations in the dollar exchange rate. Our paper adds to a growing literature that seeks to understand the role of the U.S. as the world's safe asset supplier and its impact on the global economy (see Gourinchas and Rey, 2007*b*; Caballero, Farhi and Gourinchas, 2008; Caballero and Krishnamurthy, 2009; Maggiori, 2017; He, Krishnamurthy and Milbradt, 2018; Gopinath and Stein, 2018).

Our paper explores the response of the dollar exchange rate when foreign investors impute a higher convenience yield to U.S. safe assets, such as U.S. Treasurys, than U.S. investors. This being the case, in equilibrium, foreign investors should receive a lower return in their own currencies on holding U.S. safe assets than U.S. investors. To produce lower expected returns on U.S. safe assets in foreign currency, the dollar has to appreciate today and, going forward, depreciate in expectation to deliver a lower expected return to foreign investors than U.S. investors. We derive a novel expression for the dollar exchange rate as the expected value of all future interest rate differences and convenience yields less the value of all future currency risk premia, extending the work by Froot and Ramadorai (2005) to allow for convenience yields. Our theory predicts that a country's exchange rate will appreciate whenever foreign investors increase their valuation of the current and future convenience properties of that country's safe assets.

To develop a measure of the unobserved convenience yield on U.S. safe assets derived by foreign investors, we focus on U.S. Treasury bonds as the safest among the set of U.S. safe assets. U.S. Treasury bonds are known to offer liquidity and safety services to investors which results in lower equilibrium returns to investors from holding such bonds (see Krishnamurthy and Vissing-Jorgensen, 2012; Greenwood, Hanson and Stein, 2015). Covered interest rate parity cannot hold for Treasurys when investors derive convenience yields from cash positions in these securities, while foreign bonds do not, even in the absence of frictions. In our model, the foreign convenience yield is proportional to the Treasury basis, the difference in yields between the dollar yield on short-term U.S. Treasury bonds and short-term foreign government bonds, currency-hedged, into U.S. dollars. We measure this wedge using data on spot exchange rates, forward exchange rates, and pairs of government bond yields in a panel of countries that starts in 1988. We supplement our analysis with a dataset for the U.S/U.K. cross that begins in 1970.[1] The U.S. Treasury basis is generally negative and widens during global financial crises. These negative bases are pervasive even before the recent financial crisis because the U.S. Treasury, unlike the banks who operate in LIBOR markets, chooses not to expand the supply of Treasurys in response to a negative basis.

Exchange rates seem only weakly correlated with the macro-economic and financial variables that ought to drive exchange rate variation (see, e.g., Froot and Rogoff, 1995; Frankel and Rose, 1995, on the exchange rate disconnect puzzle). Our work helps to resolve the exchange rate disconnect puzzle. Using simple univariate regressions, we show that innovations in the U.S. Treasury basis account for 23.5% of the variation in the spot dollar exchange rate, with the right sign: a decrease in the basis coincides with an appreciation of the dollar. This is true even before the financial crisis. Moreover, a decrease in the basis today predicts a future depreciation of the dollar at longer horizons. There is no relation between foreign Treasury bases and the exchange rates of the corresponding currencies. For example, a widening of the U.K. Treasury basis does not lead to a significant appreciation of the pound against other currencies. Our result lends support to the proposition that the U.S. and the U.S. dollar occupy a unique position in the international monetary system.

Convenience yields enter as wedge into the foreign investors' Euler equation and the uncovered interest parity condition. Adopting a preference-free approach, Lustig and Verdelhan (2016) demonstrate that a large class of incomplete markets models without these wedges cannot simultaneously address the U.I.P. violations, the exchange rate disconnect and the exchange rate volatility puzzles, while Itskhoki and Mukhin (2017) argue that models with such a wedge are one way to solve the exchange rate disconnect puzzle. Real exchange rates do not co-vary with macroeconomic quantities in the right way (see Backus and Smith, 1993; Kollmann, 1995). Our work shows that convenience yields introduce a wedge between the real exchange rates and the difference in the log pricing kernels that can help to resolve this disconnect.

We further examine the sources of exchange rate variation by estimating the latent factors, i.e., the conve-

[1] Results for this dataset are reported in section D of the separate Appendix and are broadly consistent with the results reported in the main text.

nience yields and risk premia, driving exchange rates. We adapt Van Binsbergen and Koijen (2010)'s approach to modeling and estimating risk premia and expected dividend growth in stock markets to our foreign exchange setting. Our state space representation of the joint dynamics of the Treasury basis, the interest rate difference and the exchange rate imposes the pricing restriction implied by our new convenience-yield valuation equation for the exchange rate. Armed with the Kalman filter estimates, we compute a Campbell-Shiller-style decomposition of exchange rate innovations into a cash flow component which tracks interest rate differences, a discount rate component which tracks currency risk premia, and, finally, the foreign investors' convenience yield component. In Froot and Ramadorai (2005)'s decomposition, the latter would have been absorbed by the discount rate component. The convenience yield channel is quantitatively important: it accounts for 54.2% of the variance of quarterly real exchange rate innovations, compared to only 3% for the cash flow component.

We find that a one standard deviation positive shock to the convenience yield drives a 4% appreciation in the dollar over the next quarters and a widening of the Treasury basis by 32 bps. Subsequently, there is a gradual reversal over the next two to three years as the high basis leads to a positive excess return on owning the U.S. dollar. Finally, our estimates imply that over 90% of the foreign convenience yield can be attributed to the exposure to dollar safe assets, rather than only U.S. Treasurys, implying that dollar-Libor deposits as well as other safe dollar-denominated assets also carry a convenience yield and are substitutes for U.S. Treasurys. For example, we construct a dollar basis for KfW bonds and show that this basis largely tracks the U.S. Treasury basis.

The paper proceeds as follows. Section 2 sets out the stylized facts regarding the U.S. Treasury basis. Section 3 lays out the convenience yield theory of exchange rates. Section 4 takes the theory to data. Section 5 estimates the model while imposing the cross-equations restrictions on exchange rates implied by the model. Section 6 concludes. The figures and tables are printed at the end of the paper. The proofs and the state space representation are in the appendix. The separate online appendix provides further derivations of the theory, additional empirical evidence, and details our data sources.

1 Related Literature

Our results lend empirical support to theories of the U.S. as the provider of world safe assets. There is ample empirical evidence that non-U.S. borrowers tilt the denomination of their borrowings (loans, deposits, bonds) especially towards the U.S. dollar (see Shin, 2012; Bräuning and Ivashina, 2017; Bruno and Shin, 2017, on bank borrowing, and corporate bond borrowing respectively). Moreover, foreign investors tilt their portfolio towards owning U.S. dollar-denominated corporate bonds when they invest in bonds denominated in foreign currencies (see Maggiori, Neiman and Schreger, 2017). The quantity evidence on the dollar bias in credit markets is silent on whether demand or supply factors are the main drivers. Our evidence from sovereign bond markets supports a demand-based explanation. The Treasury dollar basis is typically negative and reductions in the basis appreciate the dollar, suggesting that foreign investors' special demand for dollar-denominated assets lowers their expected returns.

There is a separate literature on the special role of the U.S. dollar and U.S. asset markets in the world economy. Gourinchas and Rey (2007*a*); Gourinchas, Rey and Govillot (2011); Maggiori (2017) focus on the "exorbitant privilege" of the U.S. that drives low rates of return on U.S. dollar assets. In their analysis, the low return stems from the role of the U.S. in international risk sharing, while Gopinath (2015) provides evidence on the dominant role of the dollar as an invoicing currency. Lustig, Roussanov and Verdelhan (2014) present evidence that a global dollar factor drives currency returns around the world. Our results underscore that there is something special about the dollar but does not directly speak to the evidence of this literature.

Our empirical approach is directly related to three other recent papers. First, Du, Im and Schreger (2018) also study the Treasury basis, but for a different purpose. They note that the U.S. Treasury basis is negative for short-maturity bonds, suggesting that short-maturity bonds carry a convenience yield. They delve into the term-structure of the basis, noting that the basis for long-maturity bonds has been positive recently. We use the basis to infer a convenience yield, but our main interest is in showing that the basis has explanatory power for the dollar exchange rate.[2]

Second, Valchev (2016) shows that the quantity of U.S. Treasury bonds outstanding helps to explain the return on the dollar. Valchev (2016) builds an open-economy model to relate the quantity of U.S. Treasury bonds to the convenience yield on Treasury bonds and the failure of uncovered interest parity. We show that the existence of a foreign convenience yield for U.S. Treasury bonds causes both uncovered interest parity and covered interest parity to fail. Moreover, we show that variation in the convenience yields as measured by the dollar basis explains a sizeable portion of the variation in the dollar exchange rate.

[2] An abridged version of the theory in this paper as well as results similar to that presented in Table 3 are published in Jiang, Krishnamurthy and Lustig (2018).

Third, there is a recent literature that explores the failure of LIBOR covered interest rate parity (see Ivashina, Scharfstein and Stein, 2015; Du, Tepper and Verdelhan, 2017). A common conclusion from this literature is the LIBOR-based CIP fails in part because of financial constraints faced by banks. Our results reinforce this conclusion, and we add to it by showing that LIBOR CIP fails when there is *both* foreign demand for dollar-LIBOR assets and financial constraints faced by banks. In this situation, the LIBOR basis will also reflect the foreign demand for dollar-denominated safe assets and will explain movements in the dollar exchange rate. Our empirical evidence is consistent with these points.

Finally, Froot and Thaler (1990); Gourinchas and Tornell (2004); Bacchetta and van Wincoop (2005) have argued that expectational errors are behind the failure of uncovered interest rate parity in currency markets. We also find evidence for expectational errors when examining the relation between convenience yields and the exchange rate. Our Kalman filter model allows for stickiness in expectations, while nesting the rational expectations model as a special case. The model with rational expectations is soundly rejected by the data because it cannot accommodate the time-series momentum of the dollar exchange rate in response to convenience yield shocks.

2 The U.S. Treasury Basis: Stylized Facts

Our paper relates movements in the value of the dollar exchange rate to the demand for dollar safe assets. The key metric for this demand for dollar safe assets is the U.S. Treasury basis. This section defines the Treasury basis and presents some stylized facts on the movement of the basis.

We define the U.S. Treasury basis as the difference between the yield on a cash position in U.S. Treasurys $y_t^{\$}$ and the synthetic dollar yield constructed from a cash position in a foreign government bond, which earns a yield y_t^* in foreign currency, that is hedged back into dollars:

$$x_t^{Treas} \equiv y_t^{\$} + (f_t^1 - s_t) - y_t^*, \tag{1}$$

where s_t denotes the log of the nominal exchange rate in units of foreign currency per dollar and f_t^1 denotes the log of the forward exchange rate. x_t^{Treas} measures the violation of covered interest rate parity (C.I.P.) constructed from U.S. Treasury and foreign government bond yields. A negative U.S. Treasury basis means that the U.S. Treasurys are expensive relative to their foreign counterpart. We also construct the LIBOR basis (x_t^{LIBOR}) using LIBOR rates. There is a recent literature examining the failure of the LIBOR C.I.P. condition (see Du, Tepper and Verdelhan (2017)). Our Treasury basis measure is closely related to the LIBOR C.I.P. deviation. That deviation is constructed using LIBOR rates for home and foreign countries while our basis measure is the same deviation but constructed using government bond yields for home and foreign countries. We discuss the relation between the Treasury basis and the LIBOR basis fully in Section 4.2.

We use two datasets, a panel of countries that spans 1988-2017 and a longer single time series from 1970 to 2016 for the United States/United Kingdom pair. The shorter panel is based on quarterly data from 10 developed economies. The countries are Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, United States, and United Kingdom. The sample starts in 1988Q1 and ends in 2017Q2. However, the panel is unbalanced, with data for only a few countries at the start of the sample. In order to ensure results from Treasury basis and results from Libor basis are comparable, we only include the country/quarter observations if both Treasury basis and Libor basis are available. Because New Zealand's 12-month Treasury yield is available from 1987 whereas its 12-month Libor rate is available from 1996, and Sweden's 12-month Treasury yield is available from 1984 whereas its 12-month Libor rate is available from 1991, we leave out some observations in which Treasury basis is available but Libor basis is not. We have confirmed our main empirical results are robust in the sample that contains these additional observations of Treasury basis.

The data comprises the bilateral exchange rates with respect to the U.S. dollar, 12-month bilateral forward foreign exchange contract prices, and 12-month government bond yields and LIBOR rates in all countries. We use actual rather than fitted yields for government bonds whenever possible. The main exception is the 2001:9-2008:5 period when the U.S. stopped issuing 12-month bills. We convert the daily data to quarterly frequencies using end-of-quarter observations on the same day for bond yields, interest rates, forward rates and exchange rates. There are some quarters for which all of the data are not available on the last day of the quarter, in which case we find a date earlier in the quarter, but as close to the end-of-quarter as possible, when all data are available. The Data Appendix in the separate online appendix contains information about data sources.

We construct the Treasury and LIBOR basis using the 12-month yields and forwards for each currency following (1). In each quarter, we construct the mean basis across all the countries in the panel for that quarter. Because our data is an unbalanced panel, we construct country-level changes in the basis first, and then take the cross-country average to arrive at the change in the basis. We denote the cross-sectional mean basis in the panel as $\overline{x}_t^{Treas}$. Similarly, we use $\overline{y}_t^* - y_t^{\$}$ to denote the cross-sectional average of yield differences, and $\overline{s}_t$ denotes the equally weighted cross-sectional average of the log of bilateral exchange rates against the dollar. For each of these cross-sectional averages, we employ the same set of countries that are in the sample at time t.

Figure 1 plots these series. The dotted line is the mean LIBOR basis of the U.S. dollar against the basket of currencies. The pre-crisis spikes in the average LIBOR basis are driven by idiosyncracies of LIBOR rates in Sweden (currency crisis) in 1992 and Japan in 1995 (note the difference between the mean and median LIBOR basis in 1992 and 1995). The LIBOR basis is close to zero for most of the sample and turns negative and volatile beginning in 2007. These stylized facts about the LIBOR basis are known from the work of Du, Tepper and Verdelhan (2017). The solid line is the mean Treasury basis. Unlike the LIBOR basis, the Treasury basis has always been negative and volatile. Table 1 reports the time-series moments of the Treasury basis, the Libor basis, the 12M Treasury yield difference and the 12M forward discount. The average mean Treasury basis is -25 bps per annum, which means that foreign investors are willing to give up 25 bps per annum more for holding currency-hedged U.S. Treasurys than their own bonds. The standard deviation of the mean Treasury basis is 24 bps per quarter. In contrast, the average LIBOR basis is -7 bps.

When LIBOR C.I.P. holds, the Treasury basis is simply the difference between the U.S. Treasury-LIBOR spread and its foreign counterpart:

$$x_t^{Treas} = \left(y_t^{\$} - y_t^{\$,Libor}\right) - \left(y_t^{*} - y_t^{*,Libor}\right). \tag{2}$$

Before the financial crisis, when the LIBOR basis was close to zero (-4 bps), the Treasury basis (-27 bps) is mostly due to this differential in the Treasury-LIBOR spreads. The U.S. LIBOR-Treasury spread is 23 bps larger than its foreign counterpart. During and after the crisis, this U.S. LIBOR-Treasury spread is only 7 bps per annum higher than the foreign one, while the average LIBOR basis increases to -13 bps per annum. Over the entire sample, the Treasury and LIBOR basis have a correlation of 0.36. This correlation is largely driven by the post-crisis relation where the correlation 0.56. Finally, the Treasury basis is negatively correlated (-0.27) with the U.S.-foreign Treasury yield difference and the forward discount.

Table 2 provides some statistics on the covariates of the Treasury basis. In the first column, we regress the basis on the OIS-T-bill spread which is a measure of the liquidity premium on Treasury bonds. Note that the basis is negative on average (see Figure 1). There is little relation between the basis and OIS-Tbill. The second column instead uses the spread between LIBOR and OIS. This spread is strongly negatively related to the basis and the R^2 of the regression is 69.5%. When the LIBOR-OIS spread rises, as in a flight-to-quality, the basis goes more negative. Note that OIS data is only available since 2001. Column (3) reports the correlation with the LIBOR-Tbill spread which we can construct to the start of our sample in 1988. There is a strong negative relation between the spread and the basis, and we learn from columns (1) and (2) that the relation is likely due to the LIBOR-OIS component of this spread (note also that the coefficient on LIBOR-OIS is quite similar to the coefficient on LIBOR-T-bill). Column (4) includes the spread between U.S. interest rates and the mean foreign interest rate. When U.S. rates are high relative to foreign rates, the basis is more negative. We have run specifications where we include both U.S. and foreign interest rates, and subject to the caveat that these rates do move together, the correlation seems to be driven by the U.S. interest rate and not the foreign rate. Column (5) and (6) include both the LIBOR spread and the U.S. to world interest rate differential. The explanatory power for the basis is largely driven by the LIBOR spread as one can see when comparing the R^2 in columns (5) and (6) to those in columns (3) and (4).

Section D of the separate appendix consider the U.S./U.K. Treasury basis over a longer sample and finds similar results.

3 A Theory of Spot Exchange Rates, Forward Exchange Rates and Convenience Yields on Bonds

There are two countries, foreign ($*$) and the U.S. ($), each with its own currency. We use S_t to denote the nominal exchange rate in units of foreign currency per dollar, so that an increase in S_t corresponds to an appreciation of the U.S. dollar. There are domestic (foreign) nominal bonds denominated in dollars (foreign currency). We derive bond and exchange rate pricing conditions that must be satisfied in asset market equilibrium.

For expositional purposes, we develop our basic results in a stylized model. First, we focus on the pricing of U.S. Treasury bonds as the asset that produces convenience yields. Second, for now, we assume that only U.S. bonds produce convenience yields.

As we will make clear, our theory is about the pricing of all U.S. dollar-denominated safe assets, not just U.S. Treasury bonds. On the other hand, our empirical work is specifically about the measured convenience yields on U.S. Treasury bonds. As a result, the theoretical expressions we derive for U.S. Treasury bonds are the relevant ones to guide our empirical work.

3.1 Convenience yields and exchange rates

We use y_t^* to denote the nominal yield on a one-period risk-free zero-coupon bond in foreign currency. Likewise, $y_t^\$$ denotes the nominal yield on a one-period risk-free zero-coupon Treasury bond in dollars. The stochastic discount factor (SDF) of the foreign investor is denoted M_t^*, while that of the U.S. investor is denoted $M_t^\$$.

Foreign investors price foreign bonds denominated in foreign currency, and the foreign investor's Euler equation is given by:

$$\mathbb{E}_t\left(M_{t+1}^* e^{y_t^*}\right) = 1 \tag{3}$$

Foreign investors can also invest in U.S. Treasurys. To do so, they convert local currency to U.S. dollars to receive $\frac{1}{S_t}$ dollars, invest in U.S. Treasurys, and then convert the proceeds back to local currency at date $t+1$ at S_{t+1}. Then,

$$\mathbb{E}_t\left(M_{t+1}^* \frac{S_{t+1}}{S_t} e^{y_t^\$}\right) = e^{-\lambda_t^*}, \qquad \lambda_t^* \geq 0. \tag{4}$$

The expression on the left side of the equation is standard. On the right side, we allow foreign investors in U.S. Treasurys to derive a convenience yield, λ_t^*, on their Treasury bond holdings. This λ_t^* is asset-specific and hence cannot be folded into the stochastic discount factor.

If the convenience yield rises, lowering the right side of equation (4), the required return on the investment in U.S. Treasury bonds (the left side of the equation) falls; either the expected rate of dollar appreciation declines or the yield $y_t^\$$ declines, or both.

Next, we use these pricing conditions to derive an expression linking the exchange rate and the convenience yield. We assume that $m_t^* = \log\ M_t^*$ and $\Delta s_{t+1} = \log \frac{S_{t+1}}{S_t}$ are conditionally normal. Then, (3) can be rewritten as,

$$\mathbb{E}_t\left[m_{t+1}^*\right] + \frac{1}{2} var_t\left[m_{t+1}^*\right] + y_t^* = 0, \tag{5}$$

and (4) as,

$$\mathbb{E}_t\left[m^*_{t+1}\right]+\frac{1}{2}var_t\left[m^*_{t+1}\right]+\mathbb{E}_t[\Delta s_{t+1}]+\frac{1}{2}var_t[\Delta s_{t+1}]+y^\$_t+\lambda^*_t-RP^*_t=0. \tag{6}$$

Here $RP^*_t=-cov_t\left(m^*_{t+1},\Delta s_{t+1}\right)$ is the risk premium the foreign investor requires for the exchange rate risk when investing in U.S. bonds. We combine these two expressions to find:

Lemma 1. *The expected return in levels on a long position in dollars earned by a foreign investor is decreasing in the convenience yield:*

$$\mathbb{E}_t[\Delta s_{t+1}]+\left(y^\$_t-y^*_t\right)+\frac{1}{2}var_t[\Delta s_{t+1}]=RP^*_t-\lambda^*_t \tag{7}$$

The left hand side is the excess return to a foreign investor from investing in the U.S. bond relative to the foreign bond. This is the return on the reverse carry trade, given that U.S. yields are typically lower than foreign yields. On the right hand side, the first term is the familiar currency risk premium demanded by a foreign investor going long U.S. Treasurys in dollars. The second term is the convenience yield attached by foreign investors to U.S. Treasurys: A positive convenience yield lowers the return on the reverse carry trade, i.e., the return to investing in U.S. Treasury bonds. Even in the absence of priced currency risk, $RP^*_t=0$, U.I.P. fails when the convenience yield is greater than zero.

3.2 U.S. demand for foreign bonds

Since U.S. investors have access to foreign bond markets, there is another pair of Euler equations to consider. An increase in the foreign convenience yield imputed to U.S. Treasurys implies an expected deprecation of the dollar. For a U.S. investor, buying foreign bonds when the dollar is expected to depreciate produces a high carry return. The U.S. investor's Euler equation when investing in the foreign bond is:

$$\mathbb{E}_t\left(M^\$_{t+1}\frac{S_t}{S_{t+1}}e^{y^*_t}\right)=1. \tag{8}$$

We also assume that U.S. investors derive a convenience yield when investing in U.S. Treasurys:

$$\mathbb{E}_t\left(M^\$_{t+1}e^{y^\$_t}\right)=e^{-\lambda^\$_t}. \quad \lambda^\$_t\geq 0. \tag{9}$$

$\lambda_t^{\$}$ is asset-specific. An increase in the U.S. investor's convenience yield lowers U.S. Treasury bond yields, holding the SDF fixed: $y_t^{\$} = \rho_t^{\$} - \lambda_t^{\$}$, where $\rho_t^{\$} = -\log \mathbb{E}_t \left(M_{t+1}^{\$} \right)$. We assume log-normality and rewrite these equations to derive an expression for the carry trade return,

$$\left(y_t^* - y_t^{\$} \right) - \mathbb{E}_t[\Delta s_{t+1}] + \frac{1}{2} var_t[\Delta s_{t+1}] = RP_t^{\$} + \lambda_t^{\$}. \tag{10}$$

where, $RP_t^{\$} = -cov_t \left(m_{t+1}^{\$}, -\Delta s_{t+1} \right)$ is the risk premium the U.S. investor requires for the exchange rate risk when investing in foreign bonds (i.e. the risk premium attached to the dollar appreciating).

Finally, we combine (7) and (10) to derive a cross-country restriction on the convenience yields imputed to Treasurys and the currency risk premia,

$$\lambda_t^* - \lambda_t^{\$} = RP_t^{\$} + RP_t^* - var_t[\Delta s_{t+1}]. \tag{11}$$

All else equal, an increase in λ_t^* has to be accompanied by a proportional increase in the risk premium U.S. investors ($RP_t^{\$}$) demand on foreign bonds to enforce the U.S. investor's Euler equation for foreign bonds. In an incomplete markets setting, the increase in the risk premium is a natural equilibrium outcome given that U.S. investors would increase their exposure to foreign exchange risk via the foreign bond carry trade in response to the expected depreciation of the dollar.

Thus far, we have only considered the Euler equations for risk-free assets. This raises the question of what happens when we enrich the menu of traded assets. We discuss this in section B.1 of the separate online appendix. An important theoretical result from the analysis is that λ_t^* will be equal to $\lambda_t^{\$}$ and convenience yields will not impact exchange rates if markets are complete. That is, our theory requires markets to be incomplete. However, beyond positing the Euler equations of this section, it is not necessary to specify the nature of the incompleteness to derive a relation between the convenience yield and the exchange rate. We proceed in this minimalist fashion.

3.3 Exchange rates and convenience yields

By forward iteration on eqn. (7), the level of exchange rates can be stated as a function of the interest rate differences, the currency risk premia and the future convenience yields (see Froot and Ramadorai, 2005, for a version without convenience yields). Campbell and Clarida (1987); Clarida and Gali (1994) developed an early version of this decomposition that imposed U.I.P.

Lemma 2. *The level of the nominal exchange can be written as:*

$$s_t = \mathbb{E}_t \sum_{\tau=0}^{\infty} \lambda^*_{t+\tau} + \mathbb{E}_t \sum_{\tau=0}^{\infty} (y^{\$}_{t+\tau} - y^*_{t+\tau}) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP^*_{t+\tau} - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \mathbb{E}_t [\lim_{\tau \to \infty} \; s_{t+\tau}]. \tag{12}$$

The term $\bar{s} = \mathbb{E}_t[\lim_{\tau \to \infty} \; s_{t+\tau}]$ *is constant under the assumption that the nominal exchange rate is stationary.*

The exchange rate level is determined by yield differences, the convenience yields, and the currency risk premia. This is an extension of Froot and Ramadorai (2005)'s expression for the level of exchange rates. The first term involves the sum of expected convenience yields $\lambda^*_{t+\tau}$ earned by *foreign investors* on their holdings of U.S. Treasurys. The second term involves the sum of bond yield differences. Note that the convenience yield earned by *U.S. investors* on their holdings of U.S. Treasurys lowers the U.S. Treasury yield $y^{\$}_{t+\tau}$ and hence lowers the second term. This expression implies that an increase in the expected future convenience yields earned by foreigners relative to those earned by U.S. investors should cause the dollar to appreciate today.

Alternatively, we can rewrite this equation as the sum of the convenience yield differentials, the fundamental yield differences, stripped of the convenience yields, and the risk premia:

$$s_t = \mathbb{E}_t \sum_{\tau=0}^{\infty} (\lambda^*_{t+\tau} - \lambda^{\$}_{t+\tau}) + \mathbb{E}_t \sum_{\tau=0}^{\infty} (\rho^{\$}_{t+\tau} - \rho^*_{t+\tau}) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP^*_{t+\tau} - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \mathbb{E}_t [\lim_{\tau \to \infty} \; s_{t+\tau}]. \tag{13}$$

where $\rho^{\$}_t = -\log \mathbb{E}_t \left(M^{\$}_{t+1} \right)$ is the fundamental (no convenience effect) bond yield in dollars, and likewise for foreign. Expression (13) clarifies that the exchange rate responds only to the difference in perceived convenience yields on U.S. Treasurys by foreigners and by domestic investors.

There is empirical support for the proposition that the even the nominal dollar exchange rate is stationary in our sample of developed countries. Over the last 30 years, which is our data sample, inflation has been highly correlated and similar across these developed countries, so that the nominal exchange rate is also plausibly stationary. However, when inflation rates are high and volatile across countries, nominal exchange rates are unlikely to be stationary. Next, we derive expressions for the real exchange rate, which is likely to be stationary stationary even if inflation rates are high.

We denote the log of the foreign and domestic price levels as p_t^* and $p_t^\$$, respectively. The real exchange rate is,

$$q_t = s_t + p_t^\$ - p_t^*. \tag{14}$$

We substitute the real exchange rate expression, (14), into the earlier expressions for nominal exchange rates and rewrite to find the following result.

Lemma 3. *The level of the real exchange rate can be written as:*

$$q_t = \mathbb{E}_t \sum_{\tau=0}^{\infty} \lambda_{t+\tau}^* + \mathbb{E}_t \sum_{\tau=0}^{\infty} (r_{t+\tau}^\$ - r_{t+\tau}^*) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP_{t+\tau}^* - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \bar{q}. \tag{15}$$

where, $\bar{q} = \mathbb{E}_t[\lim_{\tau \to \infty} \; q_{t+\tau}]$ *is constant under the assumption that the real exchange rate is stationary. The terms* $r_t^\$$ *and* r_t^* *are the real interest rates, i.e.,* $y_t^\$ - \mathbb{E}_t[\Delta p_{t+1}^\$]$ *is the real dollar interest rate.*

The first component measures the impact of variation in the convenience yield earned by foreign investors from holding U.S. Treasurys on the real exchange rate. The second component measures yield differences, which the effects of convenience yields earned by U.S. investors. The last component measures risk premia. In section 5 of this paper, we estimate each of these components.

3.4 Cash Treasurys, Synthetic Treasurys, and the Treasury basis

The key measure in our theory is λ_t^*, the convenience yield derived by foreign investors. This object can be inferred from the Treasury basis. To do so, we consider the foreign investor's Euler equation for an investment in a foreign government bond that is swapped into dollars via the forward market. This investment has the investor owning a bundle of safe foreign government bond and a forward position. Together, these produce a synthetic dollar asset, but one that is not as safe and liquid as the cash position in U.S. Treasurys because it involves some bank counterparty risk, and a bond and forward that are not as liquid as U.S. Treasury bonds. Thus, we posit that it provides a smaller convenience yield to U.S. Treasurys:

$$\mathbb{E}_t \left[M_{t+1}^* \frac{S_{t+1}}{F_t^1} e^{y_t^*} \right] = e^{-\beta_t^{*,h} \lambda_t^*},$$

where F_t^1 denotes the one-period forward exchange rate, in foreign currency per dollar, and $\beta_t^{*,h}\lambda_t^*$, with $0 < \beta_t^{*,h} < 1$, denotes the convenience yield on the bond+forward investment. We can use this equation along with the foreign investor's Euler equation for the U.S Treasury bond, (4), to find an expression for the unobserved U.S. Treasury convenience yield.

Lemma 4. *λ_t^*, the foreign convenience yield on U.S. Treasury bonds, is proportional to the Treasury basis:*

$$x_t^{Treas} = -(1-\beta_t^*)\lambda_t^*. \tag{16}$$

This lemma is the key to our empirical work as it provides an avenue to testing our theory linking λ_t^* and the dollar exchange rate.[3,]

3.5 Summary

We arrive at four key implications of our theory relating the Treasury basis to the dollar exchange rate.

Proposition 1. Treasury basis and the dollar

1. *The level of the nominal exchange can be written as:*

$$s_t = -\mathbb{E}_t \sum_{\tau=0}^{\infty} \frac{x_{t+\tau}^{Treas}}{1-\beta_t^*} + \mathbb{E}_t \sum_{\tau=0}^{\infty} (y_{t+\tau}^{\$} - y_{t+\tau}^*) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP_{t+\tau}^* - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \bar{s}. \tag{17}$$

2. *The level of the real exchange can be written as:*

$$q_t = -\mathbb{E}_t \sum_{\tau=0}^{\infty} \frac{x_{t+\tau}^{Treas}}{1-\beta_t^*} + \mathbb{E}_t \sum_{\tau=0}^{\infty} (r_{t+\tau}^{\$} - r_{t+\tau}^*) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP_{t+\tau}^* - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \bar{q}. \tag{18}$$

   where, $\bar{q} = \mathbb{E}_t[\lim_{\tau\to\infty} q_{t+\tau}]$ is constant under the assumption that the real exchange rate is stationary. The terms $r_t^{\$}$ and r_t^* are the real interest rates, i.e., $y_t^{\$} - \mathbb{E}_t[\Delta p_{t+1}^{\$}]$ is the real dollar interest rate.

3. *The expected log excess return to a foreign investor of a long position in Treasury bonds is increasing in*

[3] The observation that Treasury-based C.I.P. violations may be driven by convenience yields was pointed out by Adrien Verdelhan in a discussion at the Macro Finance Society (2017).

the risk premium and the Treasury basis:

$$\mathbb{E}_t[\Delta s_{t+1}] + \left(y_t^{\$} - y_t^*\right) = RP_t^* - \frac{1}{2}var_t[\Delta s_{t+1}] + \frac{x_t^{Treas}}{1-\beta_t^*} \tag{19}$$

4. *The expected log return to a foreign investor of going long the dollar via the forward contract is:*

$$\mathbb{E}_t[\Delta s_{t+1}] - (f_t^1 - s_t) = RP_t^* - \frac{1}{2}var_t[\Delta s_{t+1}] + \frac{\beta_t^*}{1-\beta_t^*}x_t^{Treas}. \tag{20}$$

3.6 Convenience yields on foreign bonds

Thus far we have assumed that foreign government bonds generate no convenience utility for its holders. This assumption allows us to clearly exposit how the convenience yield affects exchange rate determination. In the separate online appendix, we consider the realistic case in which foreign bonds also carry a convenience yield. The notation is more cumbersome, but the economics follows through. We show that all of the prior results continue to hold with the twist that λ_t^* should be interpreted as the convenience yield foreigners derive from holding U.S. Treasurys in excess of the convenience yields they derive from holding their own bonds, and $\lambda_t^{\$}$ should be interpreted as the convenience yield U.S. investors derive from U.S. Treasurys in excess of the yield derived from the foreign bonds. Additionally, x_t^{Treas} should be interpreted as the convenience yield foreigners derive from holding U.S. Treasurys relative to U.S. LIBOR assets, relative to the same object in foreign bonds. Section B.2 of the separate online appendix provides the details of the derivations.

4 Joint Dynamics of the Dollar Exchange Rate, the Treasury Basis, and the Convenience Yield

Next, we explore the empirical implication of our theory. We begin by showing that innovations to the Treasury basis covary with innovations in the dollar exchange rate. This provides support for result 1 in Proposition 1. We also show that the basis predicts future returns to a foreign investor going long Treasury bonds relative to foreign bonds, consistent with result 3 of the proposition. We then show that our results are more broadly about dollar safe assets, relate our results to the violation of LIBOR-based covered interest parity, and show that our results are specific to the dollar and do not extend to other currencies.

Before turning to the results, we discuss the choice of bond and forward maturity in measuring the Treasury basis, which is an important issue for the empirical strategy. We report results obtained using the 12-month maturity for all instruments and assets. Longer-maturity Treasury bonds carry considerable interest rate risk and may not satisfy the safe asset demand of foreign investors. In this case, their prices will not reflect a convenience yield. In fact, Du, Im and Schreger (2018) document that convenience yields when measured from long-maturity Treasury bonds have been negative recently, indicating that the safe-asset demand effects are not contained in these prices. The results using the 3-month basis are reported in section C.3 of the separate appendix. These are broadly consistent with the 12-month results but uniformly weaker, likely because the 3-month basis is a noisy measure of the long-term expectation term that drives exchange rates under our theory.

4.1 Variation in the Treasury Basis and the Dollar

We construct quarterly AR(1) innovations in the basis by regressing $\overline{x}_t^{Treas} - \overline{x}_{t-1}^{Treas}$ on $\overline{x}_{t-1}^{Treas}$ and $y_{t-1}^{\$} - \overline{y}_{t-1}^{*}$ and computing the residual, $\Delta\overline{x}_t^{Treas}$. We then regress the contemporaneous quarterly change in the spot exchange rate, $\Delta\overline{s}_t \equiv \overline{s}_t - \overline{s}_{t-1}$, on this innovation. Note that we have verified the robustness of the results reported here to the case where the innovation $\Delta\overline{x}_t^{Treas}$ is the simple change in $\overline{x}_t^{Treas}$ rather than the AR(1) innovation. The results are reported in the Separate Online Appendix.

Table 3 reports the results. From columns (1), (3), (5), (6) and (8), we see that the innovation in the Treasury basis strongly correlates with changes in the exchange rate. In the context of the well-known exchange rate disconnect puzzle (Froot and Rogoff, 1995; Frankel and Rose, 1995), the R^2s are quite high. Our regressors account for 16.1% to 42.4% of the variation in the dollar's rate of appreciation. The sign is negative as predicted by Proposition 1. The result is also stable across the pre-crisis and post-crisis sample. From column (1), we see that a 10 bps decrease in the basis (or an increase in the foreign convenience yield) below its mean coincides with a 0.96% appreciation of the U.S. dollar.

To provide a further sense of magnitudes, note that the basis is mean reverting with an AR(1) coefficient of 0.53. A 10 basis point increase in the basis today implies that next quarter's basis will be about 5 basis points, and the following quarter will be 2.5 basis points, etc. Substituting these numbers into (17) and dividing by 4 to convert to quarterly values, the sum of these future increases is $\frac{10}{4} \times \frac{1}{1-0.53} = 5.3$. From (17), to rationalize the 0.96% appreciation we need a value of $1-\beta_t^*$ of $\frac{5.3}{96} = 0.055$, suggesting that much of the foreign convenience yield derives from dollar exposure. In section 5, we will estimate β_t using the Kalman filter.

Column (3) of Table 3 includes the contemporaneous and the lagged innovation to the basis. This specification increases the R^2 to 23.5%. The explanatory power of the lag is somewhat surprising and is certainly not consistent with our model as it indicates that there is a delayed adjustment of the exchange rate to shocks to the basis. On the other hand, time-series momentum has been shown to be a common phenomena in many asset markets, including currency markets (see Moskowitz, Ooi and Pedersen, 2012), although there is no commonly agreed upon explanation for such phenomena. We will investigate this momentum phenomena more formally using a model of sticky expectations in Section 5.

Column (4) of the table includes the innovation in the interest rate differential, $y^{\$} - \overline{y}^*$, constructed analogous to the basis innovation. We see that increases in this interest rate spread has significant explanatory power in our sample. A rise in the U.S. rate relative to foreign appreciates the currency, which is what textbook models of exchange rate determination will predict (and is what equation (12) predicts). Note that a decrease in the convenience yields earned by U.S. investors will increase the U.S. Treasury yield $y^{\$}$, and cause the dollar to appreciate. We include this covariate in column (5) along with the basis innovation. The R^2 rises to 42.4% and the coefficient estimates and standard errors are nearly unchanged. This is because the basis innovation and interest rate innovation are nearly uncorrelated in this sample (note: the levels are negatively correlated).

In the Separate Appendix, we split our sample to the dates in which the Treasury basis is above or equal to the 25th percentile of its distribution, and the dates in which it is below. We find that the association between the Treasury basis and the U.S. dollar's exchange rate continues to hold strongly in the subsample in which the Treasury basis does not take extremely negative values, so that the phenomena we uncover is not only a crisis phenomena.

4.2 LIBOR and Treasury Bases

Columns (2), (7), and (9) of Table 3 shows that the LIBOR basis has explanatory power in the post-crisis sample. This result has been documented in prior work by Avdjiev et al. (2016). We note that the LIBOR basis has no explanatory power in the pre-crisis sample, and moreover has less explanatory power for the dollar than the Treasury basis. Furthermore, Du, Tepper and Verdelhan (2017) document that the LIBOR basis was near zero pre-crisis and has often been significantly different than zero post-crisis. They show that the movements in the LIBOR basis are closely connected to frictions in financial intermediation that prevent arbitrage activities. We next discuss these results and connect them to our safe asset theory.[4]

Suppose that foreign investors place a convenience yield on both dollar Treasury bonds and other dollar safe assets, including bank deposits paying LIBOR. Krishnamurthy and Vissing-Jorgensen (2012) present evidence that there is a convenience yield on both U.S. Treasury bonds and other near-riskless private bonds such as U.S. bank deposits. Moreover, they show that some investors view near-riskless private bonds as partial substitutes for Treasury bonds.

In this case, we would expect that an increase in foreign demand for Treasurys would drive down the foreign return to holding Treasurys, that is induce a widening of the Treasury basis, and drive down the foreign return to holding dollar LIBOR bank deposits. In particular, consider the LIBOR basis, which is the spread between dollar LIBOR deposits and a foreign LIBOR deposit swapped into dollars,

$$x_t^{LIBOR} \equiv y_t^{\$,LIBOR} + \left(f_t^1 - s_t - y_t^{*,LIBOR}\right). \tag{21}$$

Ceteris paribus an increase in foreign demand for dollar safe assets will drive down $y_t^{\$,LIBOR}$ and widen the LIBOR basis. However, note that such a widening of the LIBOR basis presents a riskless profit opportunity for a bank that funds itself in both dollars and foreign currency. In particular, faced with a widening LIBOR basis, a bank can increase its supply of dollar deposits by one dollar, swap the one dollar into foreign currency so that its currency risk remains unchanged and strictly increase its profits by x_t^{LIBOR}.

In the pre-crisis period, consistent with the analysis of Du, Tepper and Verdelhan (2017), banks are active on this margin and hence the LIBOR basis is zero and the LIBOR basis does not reflect foreign safe asset demand. Effectively, quantities adjust to any shifts in safe-asset demand rather than prices. In the post-crisis period, as Du, Tepper and Verdelhan (2017) emphasize, regulatory constraints on banks limit the capacity of banks to conduct the arbitrage. In this case, the LIBOR basis opens-up; prices adjust because quantities cannot. Then, the LIBOR basis reflects both safe asset demand and banks' regulatory constraints. In our explanation, the LIBOR basis becomes non-zero because of both "demand" – a desire from one set of agents to overpay for dollar deposits – and "supply" – a limited capacity of other agents to supply these dollar deposits.

[4] See Section B.3 in the separate online Appendix for a formalization of the economic argument we make in this section.

Other recent papers similarly cite both a demand factor and limited supply factor as driving the LIBOR basis (see Ivashina, Scharfstein and Stein, 2015).

Our analysis explains why the LIBOR basis comoves with the Treasury basis in the post-crisis period, as is evident from Figure 1 and why the LIBOR basis explains movements in the dollar exchange rate post-crisis. In short, when bank regulatory constraints bind on their arbitrage activities, the LIBOR basis reflects movements in λ_t^*. This raises the question of why the Treasury basis persists, given that the Treasury is unconstrained and could issue more Treasurys, similar to what banks do in LIBOR markets prior to the crisis. The answer must be that the Treasury, unlike unconstrained banks, chooses not to exploit this basis, because it has other objectives in managing government debt.

4.3 The Treasury Basis and Dollar Safe Asset Demand

Our theory posits that a specific form of capital flows consisting of flows into safe dollar assets drives the value of the U.S. dollar. This section discusses how our evidence supports this interpretation.

First, note that we construct the basis from the safest asset, the U.S. Treasury bond, and document a relation between this basis and the dollar. Second, we have shown that the LIBOR basis also helps explain movements in the dollar post-crisis, consistent with the broad dollar safe asset demand theory. Third, we compute a KfW bond basis in section C.6 of the separate Online Appendix. KfW is a German issuer whose bonds are backed by the German government, so that they are near default free. KfW issues bonds in different currencies allowing us to compute the basis for the bonds of the same issuer, i.e., holding safety fixed, in different currencies. We compute the basis for KfW bonds using one-year yields on these bonds for Australia, Euro, UK, and Switzerland against the U.S.. The KfW and the Treasury bases have roughly the same magnitude and track each other closely. This evidence shows that foreign investors' demand is for all safe assets denominated in U.S. dollars.

Lastly, we perform a placebo test of dollar safe asset demand. We repeat the univariate regression of Table 3, column (1), but using other non-U.S. countries as the base country. In Table 4, we use a different base country, and we calculate the equally weighted cross-sectional average of exchange rates and Treasury basis of other non-U.S. countries against this base country's currency. We report the coefficient of the regression of nominal exchange rate movement on Treasury basis innovation. For other countries the regression coefficients are largely statistically insignificant and/or the R^2 are considerably lower than the U.S. regressions. That is, the negative association between the exchange rate movement and Treasury basis is a phenomenon that is particularly strong for the U.S. where we posit that these safe asset demand effects should be most pronounced.

4.4 Predictability of Exchange Rates and Excess Returns

We turn to result 3 of Proposition 1, which can be read as a forecasting regression. A more negative x_t (high λ_t^*) today means that today's dollar exchange rate appreciates, which induces an expected depreciation over the next period. We define the annualized log excess return as $rx_{t\to t+k} = \frac{4}{k}\left(\Delta s_{t\to t+k}\right) + \left(y_t^{\$} - \overline{y}_t^*\right)$. Note that the LHS of equation (19) is akin to the return on the reverse currency carry trade. It involves going long the U.S. Treasury bond, funded by borrowing at the rate of the foreign government bond. The carry trade return has a risk premium term (RP), and following the literature, a proxy for this risk premium is the rate differential across the countries. Thus we include the mean Treasury yield differential at each date as a control in our regression. Note that the U.S. Treasury yield also captures variation in the convenience yields earned by U.S. investors. Additionally as we have shown in Table 3, there is a slow adjustment to basis shocks, as given by the lag of $\Delta\overline{x}_t^{Treas}$, which we also include in our regression.

We project future excess returns $rx_{t\to t+k}$ on the average Treasury basis, $\overline{x}^{Treas}$, the nominal Treasury yield difference $(y^{\$} - \overline{y}^*)$ and the change in the average Treasury basis from $t-1$ to t, as well as the lagged change in the Treasury basis (from $t-2$ to $t-1$):

$$rx_{t\to t+k} = \alpha^k + \beta_x^k \overline{x}_t^{Treas} + \beta_y^k (y_t^{\$} - \overline{y}_t^*) + \beta_L^k \Delta\overline{x}_t^{Treas} + \gamma_L^k \Delta\overline{x}_{t-1}^{Treas} + \epsilon_{t+k}^k$$

Our theory suggests that the coefficient β_x should be positive. We run this regression using quarterly data, but compute the returns on the LHS as 3-months, one-year, two-year, and three-year returns. Because there is overlap in the observations, we compute heteroskedasticity and autocorrelation adjusted standard errors.

Table 5 reports the results obtained when forecasting the annualized excess returns on a long position in the dollar. We project future excess returns $rx_{t\to t+k}$ on the average Treasury basis, $\overline{x}^{Treas}$, the nominal Treasury yield difference $(y^{\$} - \overline{y}^*)$ and the change in the average Treasury basis, as well as the lagged change in the Treasury basis. Panel A reports the regression results for the entire sample.

The slope coefficient on the average basis β_x^k varies from -16.76 at the 3-month horizon to 8.12 at the

3-year horizon. A one-standard-deviation negative shock to the basis of 20 bps increases the expected excess returns by 3.35% per annum over the first 3 months, as the dollar continues to appreciate for another quarter. However, this basis-induced momentum effect is short-lived and the slope coefficient switches signs over longer holding periods. The long-horizon estimates are an accurate reflection of the basis effect after stripping away the momentum effect, as we discuss below. These effects are economically significant. A one-standard-deviation basis shock of 20 bps. raises the expected excess return by 1.62% per annum over the next three years. These regressors jointly explain about 12% of the variation in excess returns at the 3-year horizon. The basis is also not a persistent regressor. Hence, there is no mechanical relation between the forecasting horizon and the R^2.

From equation (10) we see that the value of $\frac{1}{1-\beta^*}$ is equal to β_x for the 1-year horizon, but the 1-year β_x for 1-year is small and imprecisely estimated, likely because of the momentum effect we have found. A lower bound for $1-\beta^*$ is $\frac{1}{\beta_x}$ on the 2- and 3-year horizon regressions. This is a lower bound because a shock to the basis gradually reverses over time (we explore this formally in the next section), so that the returns in the 2nd and 3rd year are responding to a smaller value of the basis. This gives a lower bound for $1-\beta^*$ of 0.12. Panel B and C of Table 5 report the regression results for the pre-and post-crisis sample. The momentum effect is only present prior the crisis. In the post-crisis sample, the slope coefficients on the basis are all positive. At the 3-year horizon, the coefficient is 13.35: A one-standard-deviation basis shock raises the expected excess return by 2.67% per annum over the next three years. There is much more predictability after the crisis. In the post-crisis sample, these regressors jointly explain about 35% of the joint variation in excess returns at the 3-year horizon.

Table 6 provides more detail for the 3-year forecasting results. Panel A reports the 3-year results. Panel B excludes the first four quarters from the cumulative excess return to remove the momentum effect. At the 3-year horizon, the slope coefficient in a univariate regression of returns on the basis is not statistically significantly different from zero. However, when we exclude the first 4 quarters from the left-hand-side return, the slope coefficient in the univariate regression increases from 1.62 to 6.54. The variation in the Treasury basis explains 7% of the variation in the returns at the 3-year horizon.

The return predictability is mostly driven by the exchange rate component of returns. Table A.8 and A.9 in section C.5 of the Separate Appendix report predictability results for exchange rate changes rather than excess returns. After excluding the first year, there is solid statistical evidence that the average Treasury basis forecasts changes in exchange rates: the slope coefficient estimate is 11.24, implying that the dollar appreciates by 2.24% per annum over the next 3 years following a one-standard-deviation widening of average Treasury basis. When we add the other regressors, the R^2 increases to 15% over the entire sample. The slope coefficient on the Treasury basis increases to 13.01. A one-standard-deviation positive basis shock raises the expected excess return by 2.60% per annum over the next three years. The coefficient estimates reported in Panel B after cleansing the cumulative returns of the momentum effect are remarkably stable across subsamples.

In section D of the separate appendix, we construct a longer sample for the U.S.-U.K. Treasury basis (the sample starts 1970Q1 and ends in 2016Q2), and we run the same battery of statistical tests. The results are broadly in line with the results obtained on the shorter sample for the G-10 currencies.

5 Structural Estimation of the Convenience Yield Model

In this section, we investigate the dynamic relation between the convenience yield, risk premia, interest rates, and the exchange rates through the lens of a structural model. The model allows us to infer the unobserved risk premia and convenience yields earned by foreign investors driving the exchange rate dynamics and quantify the relative importance of the convenience yield earned by foreign investors, the currency risk premia, and interest rates in explaining exchange rate behavior.

5.1 Exchange Rate Model with Sticky Expectations

To accommodate the evidence of time-series momentum in the dollar exchange rate, we analyze a version of the model in which foreign exchange investors have sticky expectations: we posit that these investors do not update their expectations each period, but, when they do, they use the right model, as in Mankiw and Reis (2002). We assume that in any given period, a fraction $(1-\varpi)$ of investors update their information set each period. When they update, they use rational expectations. We use $\mathbb{F}_t$ to denote the cross-sectional average of the sticky information forecasts. Reis (2006) shows that the cross-sectional average forecast of a variable x_t h periods from now is simply given by: $\mathbb{F}_t x_{t+h} = (1-\varpi)\sum_{j=0}^{\infty}\varpi^j \mathbb{E}_{t-j}x_{t+h}$.

We posit the following autoregressive processes for the convenience yield earned by foreign investors, the real interest rate difference $i_t = r_t^{\$} - r_t^*$ and the (negative of the) risk premium $rp_t = -(RP_t^* - Var_t(\Delta q_{t+1}))$:

$$\lambda_{t+1}^* = \gamma_0 + \gamma_1 \lambda_t^* + \varepsilon_{t+1}^{\lambda}, \tag{22}$$

$$i_{t+1} = \psi_0 + \psi_1 i_t + \varepsilon_{t+1}^{i}, \tag{23}$$

$$rp_{t+1} = \delta_0 + \delta_1 rp_t + \varepsilon_{t+1}^{rp}. \tag{24}$$

Next we assume that the equilibrium exchange rate reflects the cross-sectional average across all investors of their forecasts of the convenience yield earned by foreign investors, interest rate differences, and risk premium components of equation (18). We show in the appendix that:

Proposition 2. *The log of the real exchange rate can be stated as a function of current and lagged fundamentals:*

$$q_t = \bar{q} + (1 - \varpi) \sum_{j=0}^{\infty} \left(\frac{\varpi^j \gamma_1^j}{1 - \gamma_1} (\lambda_{t-j}^* - \theta_\lambda) + \frac{\varpi^j \psi_1^j}{1 - \psi_1} (i_{t-j} - \theta_i) + \frac{\varpi^j \delta_1^j}{1 - \delta_1} (rp_{t-j} - \theta_{rp}) \right). \tag{25}$$

The first term measures the impact of the convenience yields earned by foreign investors on the real exchange rate. The second term measures the interest rate differences. The U.S. Treasury yield includes the effect of the U.S. investors' convenience yields. The last term comprises the currency risk premia.

To understand this expression, it is helpful to consider the case of rational expectations. As $\varpi \to 0$, the expression simplifies to $q_t = \bar{q} + \frac{1}{1-\gamma_1}(\lambda_t^* - \theta_\lambda) + \frac{1}{1-\psi_1}(i_t - \theta_i) + \frac{1}{1-\delta_1}(rp_t - \theta_{rp})$, which is the equivalent of eqn. (12). The terms here correspond to the sum of future convenience yields earned by foreign investors, the interest rate differentials, and the risk premia, evaluated under the assumed AR(1) structure. When investors have sticky expectations ($\varpi > 0$), the log of the real exchange rate adjust slowly to new information about convenience yields, interest rates and risk premia, as captured by the lagged terms in the sum in (25). The inertia in the response to basis shocks is essential to match the time-series dynamics of the real exchange rate.

5.2 State Space Representation

The convenience yield earned by foreign investors, λ_t^*, the currency risk premium, rp_t, and the fraction of convenience foreign investors derive from the synthetic dollar Treasury position relative to the cash Treasury,

β_t, are latent factors driving exchange rates. Given the linear structure of the model, we can use the Kalman filter to learn about these latent factors, while imposing the present-value relation for the real exchange rate. Methodologically, we adapt Van Binsbergen and Koijen (2010)'s approach to learning about expected returns and cash flow growth in stock markets to currency markets.

To develop a state-space representation of this model, we define three new state variables which summarize the average investor's expectations about future convenience yields, interest rate differences and risk premia, respectively $x_t^\lambda, y_t^i, z_t^{rp}$, so that,

$$q_t = \bar{q} + (1-\varpi)(x_t^\lambda + y_t^i + z_t^{rp}).$$

Given this definition, the state variables have an autoregressive structure:

$$x_t^\lambda = \varpi\gamma_1 x_{t-1}^\lambda + \frac{1}{1-\gamma_1}(\lambda_t^* - \theta_\lambda), \quad (26)$$

$$y_t^i = \varpi\psi_1 y_{t-1}^i + \frac{1}{1-\psi_1}(i_t - \theta_i), \quad (27)$$

$$z_t^{rp} = \varpi\delta_1 z_{t-1}^{rp} + \frac{1}{1-\delta_1}(rp_t - \theta_{rp}). \quad (28)$$

We observe the Treasury basis and the log of the real exchange rate:

$$x_{t+1}^{Treasury} = -(1-\beta_{t+1}^*)\lambda_{t+1}^* + \varepsilon_{t+1}^x, \quad (29)$$

$$q_{t+1} = \bar{q} + (1-\varpi)(x_{t+1}^\lambda + y_{t+1}^i + z_{t+1}^{rp}), \quad (30)$$

as well as the interest rate differential, i_{t+1}. Note that we allow for measurement error in the Treasury basis. Furthermore, we allow β_t^* to vary over time and write its dynamics as

$$\beta_t^* = \kappa_0 + \kappa_1\beta_{t-1}^* + \varepsilon_t^\beta.$$

The Kalman filter estimates the unobserved state variables from the observable variables given our assumed dynamics for λ_t^*, i_t, rp_t, and β_t^*. In matrix notation, the model with sticky expectations has the following succinct state space representation:

$$\begin{aligned}\boldsymbol{Y}_t &= \boldsymbol{Z}_t + h(\boldsymbol{X}_t) + \boldsymbol{e}_t, \boldsymbol{e}_t \sim N(\boldsymbol{0}, \boldsymbol{H}_t),\\ \boldsymbol{X}_{t+1} &= \boldsymbol{A}_t + \boldsymbol{F}\boldsymbol{X}_t + \boldsymbol{\Gamma}\boldsymbol{\eta}_{t+1}, \boldsymbol{\eta}_{t+1} \sim N(\boldsymbol{0}, \boldsymbol{Q}_t),\end{aligned}$$

where $\boldsymbol{Y}_t$ is a 3-dimensional observation vector and $\boldsymbol{X}_t$ is an 11-dimensional state vector:

$$\boldsymbol{Y}_t' = \begin{bmatrix} x_t^{Treas} & i_t & \Delta q_t \end{bmatrix}, \boldsymbol{X}_t' = \begin{bmatrix} x_{t-1}^{\lambda} & y_{t-1}^{i} & x_{t-2}^{\lambda} & y_{t-2}^{i} & \lambda_{t-1} & \beta_{t-1}^{*} & rp_{t-1} & \varepsilon_t^{\lambda} & \varepsilon_t^{\beta} & \varepsilon_t^{rp} & \varepsilon_t^{i} \end{bmatrix}.$$

The other matrices are defined in the appendix.

5.3 Identification and Estimation

To fully identify all parameters, we need to impose two restrictions on the variance-covariance matrix of the residuals in the transition equations. Our identification assumptions are that innovations to the risk premium are uncorrelated with innovations to the convenience yield and innovations to β_t^*. The appendix discusses identification in detail. Given these identification assumptions, we search over the parameter space to maximize the log likelihood of the residual, defined as

$$\mathcal{L} = \sum_t -\frac{1}{2}\log|\boldsymbol{S}_t| - \frac{1}{2}\boldsymbol{\delta}_t'\boldsymbol{S}_t^{-1}\boldsymbol{\delta}_t, \tag{31}$$

$$\boldsymbol{\delta}_t = \boldsymbol{Y}_t - h(\boldsymbol{A}_{t-1} + \boldsymbol{F}\boldsymbol{X}_{t-1}), \tag{32}$$

$$\boldsymbol{S}_t = var(\boldsymbol{e}_t) + \boldsymbol{H}_t\boldsymbol{P}_{t|t-1}\boldsymbol{H}_t', \tag{33}$$

where $\boldsymbol{P}_{t|t-1}$ is the *a priori* error covariance of the state variable, and $\boldsymbol{H}_t$ is the first-order derivative of $h(\cdot)$ at time t. In the estimation, we use the log of the real exchange rate, the 12-M Treasury dollar basis and the real interest rate spreads for an equal-weighted basket of 9 countries against the dollar.

Table 7 reports the parameter estimates. Panel A reports the restrictions on the coefficients. Panel B reports the ML estimates. The estimate for the stickiness parameter ϖ is 0.42, consistent with quantitatively significant stickiness of expectations: only 58% of FX investors adjust their expectations each quarter. The λ innovation is positively correlated with the β innovation (0.54). Since the Treasury basis $x_t^{Treas} = (1-\beta_t^*)\lambda_t^*$, the positive correlation implies that the dollar typically appreciates more than what is suggested by the widening of the Treasury basis alone, because an increase in λ would typically be offset by an increase in β.

Our estimates impute a negligible amount of measurement error to the Treasury basis; σ^x is close to zero. According to our estimates, the convenience yield and the risk premium are the most persistent state variables. The quarterly (annual) persistence of the convenience yield and the risk premium is 0.90 (0.66), while the quarterly (annual) persistence of the real interest difference is 0.84 (0.50). The real exchange rate inherits the persistence of these state variables, but the persistence is amplified by the stickiness of expectations.

Finally, in Panel C, we report the R^2 values for the Treasury basis: $R^2 = 1 - \frac{\hat{var}(x_t^{Treas} - (1-\beta_t^*)\lambda_t^*)}{\hat{var}(x_t^{Treas})}$, which estimates the fraction of variation in the observable variable x_t^{Treas} that is accounted for by the latent variables. The latent factors account for 81% of the variation in the U.S. Treasury basis.

We use the likelihood ratio (LR) test to test the null that investors have rational expectations: $\varpi = 0$. We use $\mathcal{L}_1$ ($\mathcal{L}_0$) to denote the log-likelihood that corresponds to the unconstrained (constrained) model. The likelihood ratio test statistic is then given by $LR = 2(\mathcal{L}_1 - \mathcal{L}_0)$, which is asymptotically chi-squared distributed with degrees of freedom equal to the number of constrained parameters. Table 8 reports the test results. The null hypotheses that investors have rational expectations is rejected at the 1% level. Expectations about future convenience yields, future risk premia, and future interest rates seem sticky. For example, when the basis widens and the convenience yield rises, the dollar exchange rate continues to appreciate for a few quarters. The rational expectations model cannot accommodate this feature of the data, leading to a rejection of the null.

The time-varying β_t^* introduces a wedge between Treasury basis variation and the variation in the convenience yield. We test the null hypothesis that the beta is constant: $\kappa_1 = \sigma^\beta = \rho_{\lambda,\beta} = \rho_{\beta,i} = 0$. The null hypotheses is rejected at the 1% level, consistent with the notion that the perceived substitutability of U.S Treasurys and synthetic Treasurys is time-varying.

Figure 2 plots the filtered series for x_t^λ, y_t^i, λ_t, β_t^*, the risk premium rp_t, the filtered and observed Treasury basis, the interest rate differences and changes in the log real exchange rate. λ_t is a near mirror of the observed basis, x_t, and is highest in 1993 around the European currency crises and 2008 around the global financial crisis. The β varies between 0.83 and 1.05. Economically, it does not make sense for β to exceed one, but we have kept to a linear structure and not imposed any restrictions on β.

The moments of these series are reported in Table 9. On average, the quarterly convenience yield λ is 0.90%, while the average β^* is 0.93, consistent with an average U.S. Treasury basis (ignoring the covariance) of -6 bps per quarter or -24 bps per annum. The average β^* of 0.93 suggest that more than 90% of the convenience yield that foreign investors derive from a cash position in U.S. Treasurys is derived from the exposure to a safe dollar assets rather than only Treasurys. This surprisingly large estimate is identified in the data from the large response of the dollar exchange rate to a shock to the the Treasury basis in the data. We note that this result is consistent with our evidence from the LIBOR basis and the KfW basis which non-Treasury safe dollar assets.

5.4 Exchange Rate Accounting

In our model, the real exchange rate innovation is given by:

$$(\mathbb{E}_t - \mathbb{E}_{t-1})q_t \quad = \quad (1-\varpi)\left(\frac{1}{1-\gamma_1}\varepsilon_t^{\lambda} + \frac{1}{1-\psi_1}\varepsilon_t^{i} + \frac{1}{1-\delta_1}\varepsilon_t^{rp}\right). \tag{34}$$

The effect of an innovation on the real exchange rate is amplified by the persistence of the underlying shocks (γ_1, ψ_1, and δ_1), but is dampened by the stickiness of expectations, ϖ. We define the following objects: news about current and future convenience yields earned by *foreign investors*, interest rates and risk premia:

$$N_t^{\lambda} = \frac{1-\varpi}{1-\gamma_1}\varepsilon_t^{\lambda}, \qquad N_t^{i} = \frac{1-\varpi}{1-\psi_1}\varepsilon_t^{i}, \qquad N_t^{rp} = \frac{1-\varpi}{1-\delta_1}\varepsilon_t^{rp}. \tag{35}$$

The second term includes the effects of news about convenience yields earned by *U.S. investors* through the effect on U.S. interest rates. The variance of the innovation to the real exchange rate can be decomposed into the variance of each of these components, as well as the covariance terms:

$$\begin{aligned} var((\mathbb{E}_t - \mathbb{E}_{t-1})q_t) \quad &= \quad var(N_t^{\lambda}) + var(N_t^{i}) + var(N_t^{rp}) \\ &+ \quad cov(N_t^{\lambda}, N_t^{i}) + cov(N_t^{\lambda}, N_t^{rp}) + cov(N_t^{i}, N_t^{rp}). \end{aligned} \tag{36}$$

Similarly, the variance of the real exchange rate level can be decomposed as

$$\begin{aligned} var(q_t) &= var(L_t^{\lambda}) + var(L_t^{i}) + var(L_t^{rp}) \\ &+ 2cov(L_t^{\lambda}, L_t^{i}) + 2cov(L_t^{\lambda}, L_t^{rp}) + 2cov(L_t^{i}, L_t^{rp}), \end{aligned} \tag{37}$$

where $L_t^{\lambda} = (1-\omega)x_t^{\lambda}, L_t^{i} = (1-\omega)y_t^{i}, L_t^{rp} = (1-\omega)z_t^{rp}$.

Table 10 reports the results. Panel A (B) considers real exchange rate innovations (levels). Our estimates impute a quantitatively important role to shocks to safe asset demand and supply when accounting for variation in the dollar exchange rate. News about future convenience yields earned by foreign investors N_t^{λ} accounts for 54.2% of the variation in the innovation to the dollar's real exchange rate, while news about future risk premia accounts for 39.9%, and interest rate news only accounts for 2.7%. The latter partly reflects the lower persistence of interest rates.

Figure 3 plots the real exchange rate movement and the convenience yield news N_t^{λ}. According to our estimates, news about current and future convenience yields on U.S. dollar denominated safe assets that accrue to foreign investors is the main driver of variation in the dollar exchange rate at quarterly frequencies.

In Table 11, we examine how well the filtered convenience yield innovation (ε_t^{λ}) explains concurrent real exchange rate variation. Since the Kalman filter efficiently uses information contained in all observed variables in the state space system, the filtered convenience yield innovation explains a larger fraction of real exchange rate variation than the innovation to U.S. Treasury basis. Still, it is worth noting that the magnitude of this improvement: the R^2 increases from 16% to 73% when we use the convenience yield innovations as regressors. The slope coefficients using the innovations are stable across subsamples, while the results using the Treasury basis innovations are not. In the model with sticky expectations, we expect the lagged λ_t^* shock to affect current exchange rate movement. This is confirmed in Column (3), (6) and (9). These results are consistent with the variance decomposition results.

5.5 Impulse Responses

In the Appendix, we explicitly compute the impulse responses to innovations.

Corollary 1. *The impulse response function of the real exchange rate q_t to innovations $(\epsilon^{\lambda}, \epsilon^{i}, \epsilon_{rp})$ j periods* *after impact is given by:*

$$\begin{aligned}
\varphi_{\lambda,j} &= \frac{(1-\varpi)}{1-\gamma_1}\left[\frac{(1-\varpi^j)}{1-\varpi}\gamma_1^j + \left(\varpi^j\gamma_1^j - 1\right)\right] \textit{ for } j = 0,1,2,\ldots. \\
\varphi_{i,j} &= \frac{(1-\varpi)}{1-\psi_1}\left[\frac{(1-\varpi^j)}{1-\varpi}\psi_1^j + \left(\varpi^j\psi_1^j - 1\right)\right] \textit{ for } j = 0,1,2,\ldots. \\
\varphi_{rp,j} &= \frac{(1-\varpi)}{1-\delta_1}\left[\frac{(1-\varpi^j)}{1-\varpi}\delta_1^j + \left(\varpi^j\delta_1^j - 1\right)\right] \textit{ for } j = 0,1,2,\ldots.
\end{aligned}$$

We plot the response to a one-standard deviation foreign investors' convenience yield shock, denoted ε_t^λ, in Figure 4. The innovation to the convenience yield ε_t^λ increases λ_t^* in period $t = 1$ (top-left panel) and also raises the state variable x_t^λ, which keeps track of the average investor's sticky expectations, shown in the top-right panel. As a result, the real exchange rate (bottom-left panel) continues to appreciate in the next quarter, leading to return momentum in the short run. The real exchange rate inherits the dynamics of x_t^λ. A one standard deviation shock to convenience yields causes a 4.5% appreciation after 3 quarters, while the quarterly Treasury basis widens by 32 bps. After 3 quarters, the real exchange rate continues to depreciate, and the realized excess returns (bottom-left panel) are negative. The dotted lines plot the rational expectations $\varpi = 0$ case, holding all other parameters fixed. In the case of rational expectations $\varpi = 0$, the real exchange rate jumps upon impact and then gradually depreciates. In this case, there is no momentum in realized returns.

Figure 5 reports the impulse response to a shock to the interest rate differential. The dynamics of each variable pictured is qualitatively similar to that of Figure 4. The main difference is quantitative: the exchange rate appreciates by about 1% in response to the one-standard deviation increase in U.S. rates, consistent with U.I.P. for these innovations, while it moves 4.5% in the case of the convenience yield shock.

The impulse response in Figure 5 shows that U.I.P is restored in U.S. data once you allow for convenience yield shocks. Textbook models predict that the exchange rate should appreciate immediately in response to an unexpected increase in the home interest rate, and then depreciate after that, but U.I.P. is soundly rejected in the data (see Hansen and Hodrick, 1980; Fama, 1984): an increase in the home interest rate above that country's average leads to a subsequent appreciation of the home currency. Recently, Engel (2016); Valchev (2016); Dahlquist and Penasse (2016) show that an increase in the short-term interest rate forecasts a short-horizon appreciation in the dollar, which is inconsistent with standard models, and a long-horizon depreciation, consistent with theoretical models. Engel (2016) describes an exchange rate model with a role for liquidity services provided by short-term debt that could explain the dynamics of U.I.P. deviations without directly measuring these convenience yields. The bottom-left panel of Figure 5 shows that U.I.P. is essentially restored once we control for convenience yields, as suggested by the theory: in eqn. (7), U.I.P. holds if $RP = 0$ only when we control for variation in λ_t^*.

Finally, Section E of the separate online appendix documents quantitatively similar impulse responses obtained from a VAR, in which structural innovations to the convenience yield are identified using a Cholesky decomposition.

5.6 Return Predictability

To help us understand the return predictability produced by this model, we derive an expression for the excess return expected by a rational investor who continuously updates her expectations:

Proposition 3. *In the model with sticky expectations, the log excess return expected by a rational foreign investor on a long position in U.S. Treasury bonds relative to the foreign bond is:*

$$\begin{aligned}\mathbb{E}_t rx_{t+1} = \mathbb{E}_t[\Delta q_{t+1}] + i_t &= i_t + (1-\varpi)\left(\frac{\gamma_1}{1-\gamma_1}(\lambda_t^* - \theta_\lambda) + \frac{\psi_1}{1-\psi_1}(i_t - \theta_i) + \frac{\delta_1}{1-\delta_1}(rp_t - \theta_{rp})\right) \\ &+ (1-\varpi)\left((\varpi\gamma_1 - 1)x_t^\lambda + (\varpi\psi_1 - 1)y_t^i + (\varpi\delta_1 - 1)z_t^{rp}\right).\end{aligned}$$

When all agents have rational expectations $\varpi = 0$, this collapses to the standard expression:

$$\mathbb{E}_t[\Delta q_{t+1}] + i_t = RP_t^* - \lambda_t.$$

An increase in λ decreases the expected excess return one-for-one. However, in the sticky expectations case, returns will be predictable by all three state variables. Indeed, if ϖ is large enough, an increase in λ initially increases expected excess returns.

Corollary 2. *At longer horizons, the rate of appreciation expected by a rational investor is given by*

$$\begin{aligned}\mathbb{E}_t(q_{t+k} - q_t) &= (1-\varpi^k)\left(\frac{\gamma_1^k}{1-\gamma_1}(\lambda_t^* - \theta_\lambda) + \frac{\psi_1^k}{1-\psi_1}(i_t - \theta_i) + \frac{\delta_1^k}{1-\delta_1}(rp_t - \theta_{rp})\right) \\ &+ (1-\varpi)\left((\varpi^k\gamma_1^k - 1)x_t^\lambda + (\varpi^k\psi_1^k - 1)y_t^i + (\varpi^k\delta_1^k - 1)z_t^{rp}\right).\end{aligned} \quad (38)$$

Next, we examine whether the filtered convenience yield variables x_t^λ and λ_t^* also predict the real excess return of the dollar. There are two opposing forces of the convenience yield in the model with sticky expectations. Since our estimates imply $\varpi\gamma_1 - 1 < 0$ and $\gamma_1 > 0$, we expect that a higher x_t^λ predicts a lower expected return on the dollar while a higher λ_t^* predicts a higher expected return. When $\varpi = 0$, there is only a single predictor λ_t and the slope coefficient on λ_t is -1 for the next period's excess return. The sticky expectations model predicts different effects of x_t^λ and λ_t^*. In the next 3 months to 1 year, a shock that increases λ_t^* predicts a higher return on the dollar because of the time-series momentum induced by sticky expectations. In the next 1 year to 2 years, a higher x_t^λ predicts a lower return on the dollar because the convenience yield on the U.S. Treasury will eventually fall. In this way, our model allows us to disentangle the short-run momentum and the long-run predictability effects. As can be verified from equation (2), for intermediate horizons, where $\gamma_1^j \approx \left(1 - \varpi^j\gamma_1^j\right)$, there is no return predictability induced by the convenience yield: The momentum and the convenience yield effect offset each other.

Table 12 reports the loadings of the optimal forecast on the latent variables. A higher λ_t^* predicts a higher return on the dollar, while a higher x_t^λ predicts a lower return on the dollar. As the forecast horizon j goes up, the loadings on the spot variables λ_t^*, i_t, and rp_t decline whereas the loadings on the stock variables x_t^λ, y_t^i, z_t^{rp} increase. Therefore, a positive ε_t^λ shock will first increase the expected return on the dollar and then decrease the expected return on the dollar.

We check how much of the variation in realized returns on long positions in U.S. Treasurys and short positions in foreign bonds is accounted for by the model's conditional forecast $\mathbb{E}_t(q_{t+k} - q_t)$ (conditional on the shock and filtered state variables at time t). Panel A in Table 13 reports the results. At horizons of 2 to 3 years, the Kalman filtered forecast $\mathbb{E}_t(q_{t+k} - q_t)$ explains between 12 and 13% of the variation in the future excess returns. As a result, the structural model matches the predictability generated by the linear regressions with the U.S. Treasury basis, innovations to the U.S. Treasury basis and the interest rate difference. Panel B controls for the U.S. Treasury basis, innovation to the U.S. Treasury basis and the real interest rate difference. However, the Treasury basis has some incremental forecasting power, even when controlling for the model's optimal forecast, indicating that the model fit can be improved.

5.7 Out-of-Sample Predictability

Since our parameter estimates are obtained from the entire sample, the filtered variables may contain look-ahead bias. To examine whether the model-based conditional expectation predicts the exchange rate movement out of sample, we split our sample into two halves: We use the first half of the sample (1988Q3—2002Q2) to estimate the parameters, and use the second half (2002Q3—2016Q2) to examine return predictability.

Using these parameter estimates we run the one-sided Kalman filter and construct the conditional expectations of exchange rate movements. Then, we repeat the return predictablity regressions in Table 13.

Table 14 reports the predictability results. Panel A is the in-sample prediction for the period from 1988Q3 – 2002Q2. These results are in line with our full sample result. Panel B contains the out-of-sample test. The expectation of $E_t(q_{t+k} - q_t)$ is constructed using parameter estimates from the first half of the sample along with time-t observables that are the input into the Kalman filter. We note that an expected dollar appreciation indeed predicts a dollar appreciation rejecting the null hypothesis that the exchange rate movement is unpredictable.

Next, we consider the regression of exchange rate changes from t to $t+1$ on the Kalman-filtered convenience yield innovations at $t+1$. For this exercise, we estimate the Kalman filter from the first-half of the sample, and use these estimates to extract the convenience yield shocks ε_t^{λ}. We re-run the contemporaneous regression of the dollar's nominal exchange rate movement on the innovation to the Treasury basis and the convenience yield shock, and report the result in Table 15. Consistent with Table 11, the convenience yield shock and its lagged value both explain the exchange rate movement. However, note that the R^2 in column (6) is 25%, while in our baseline full-sample estimate (column (9) of Table 11), the R^2 is 77%.

6 Conclusion

We present a theory of exchange rates which departs from existing theories by imputing a central role to international flows in Treasury debt and related dollar safe asset markets in exchange rate determination.

In our theory, the spot exchange rate of a safe asset currency will reflect the cumulative value of all future convenience yields that are earned by foreign investors on safe assets denominated in that currency. The empirical evidence strongly supports the theory. Our results shed light on two important topics in international finance. First, we help to resolve the exchange rate disconnect puzzle by demonstrating that shocks to the demand for dollar-denominated safe assets drive a sizeable portion of the variation in the dollar exchange rate. Second, we provide strong empirical support for recent theories regarding safe assets and the central role of the U.S. in the international monetary system.

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Table 1: Summary Statistics of Cross-sectional Mean Basis and Interest Rate Difference

Table reports summary statistics in percentage points for the 12-M Treasury dollar basis $\overline{x}^{Treas}$, the Libor dollar basis $\overline{x}^{Libor}$, the 12M yield spread $y^{\$} - \overline{y}^{*}$, and the 12M forward discount $\overline{f-s}$ in logs. Table reports time-series averages, time-series standard deviations and correlations. Numbers reported are time-series moments of the cross-sectional means of the unbalanced Panel. The countries are Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, United States, and United Kingdom. The sample starts in 1988Q1 and ends in 2017Q2. For each of these cross-sectional averages, we employ the same set of countries that are in the sample at time t.

	$\overline{x}^{Treas}$	$\overline{x}^{Libor}$	$y^{\$} - \overline{y}^{*}$	$\overline{f-s}$
Panel A: 1988Q1–2017Q2				
mean	-0.25	-0.07	-0.45	-0.20
stdev	0.24	0.18	1.87	1.95
skew	-1.33	-3.08	-0.61	-0.29
$\overline{x}^{Treas}$	1.00	0.36	-0.27	-0.39
$\overline{x}^{Libor}$	0.36	1.00	0.47	0.40
$y^{U.S.} - \overline{y}^{*}$	-0.27	0.47	1.00	0.99
Panel B: 1988Q1–2007Q4				
mean	-0.27	-0.04	-0.33	-0.05
stdev	0.26	0.16	2.25	2.34
skew	-0.93	-4.62	-0.69	-0.42
$\overline{x}^{Treas}$	1.00	0.34	-0.31	-0.41
$\overline{x}^{Libor}$	0.34	1.00	0.54	0.48
$y^{U.S.} - \overline{y}^{*}$	-0.31	0.54	1.00	0.99
Panel C: 2008Q1–2017Q2				
mean	-0.20	-0.13	-0.64	-0.44
stdev	0.21	0.20	0.78	0.80
skew	-2.51	-1.78	0.16	0.09
$\overline{x}^{Treas}$	1.00	0.56	0.00	-0.25
$\overline{x}^{Libor}$	0.56	1.00	0.51	0.35
$y^{U.S.} - \overline{y}^{*}$	0.00	0.51	1.00	0.97

Table 2: The Treasury Basis and Interest Rate Spreads

We regress the quarterly average Treasury basis, $\overline{x}^{Treas}$, on a number of U.S. money market spreads and the U.S. to foreign government bond interest rate differential. The spreads and interest rate differential are constructed as the quarterly average of the indicated series. Data is from 1988Q1 to 2017Q2 for the regressions with 118 observations and 2001Q4 to 2017Q2 for the regressions with 63 observations. OLS standard errors in parentheses.

	(1)	(2)	(3)	(4)	(5)	(6)
U.S. 6-month OIS−T-bill	0.13					
	(0.18)					
U.S. 6-month LIBOR−OIS		-0.40			-0.44	
		(0.034)			(0.027)	
U.S. 6-month LIBOR−T-bill			-0.47			-0.43
			(0.045)			(0.044)
$y^{\$} - \overline{y}^{*}$				-0.047	-0.07	-0.029
				(0.01)	(0.01)	(0.007)
R^2	1%	69.5	48	16.7	82.9	54
N	63	63	118	118	63	118

Table 3: Average Treasury Basis and the USD Spot Nominal Exchange Rate

The dependent variable is the quarterly change in the log of the spot USD exchange rate against a basket. The independent variables are the innovation in the average Treasury basis, $\Delta\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation, the innovation in the LIBOR basis, and the innovation in the U.S.-to-foreign Treasury yield differential. Data is quarterly. OLS standard errors in parentheses.

	1988Q1−2017Q2					1988Q1−2007Q4		2008Q1−2017Q2	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta\overline{x}^{Treas}$	-9.62		-9.70		-9.30	-8.52		-12.93	
	(2.05)		(1.95)		(1.70)	(2.58)		(3.18)	
$\Delta\overline{x}^{LIBOR}$		-2.48					2.79		-10.05
		(3.05)					(4.19)		(3.98)
Lag $\Delta\overline{x}^{Treas}$			-6.65		-6.21				
			(1.94)		(1.69)				
$\Delta(y^{\$} - \overline{y}^{*})$				3.80	3.59				
				(0.70)	(0.59)				
R^2	16.1	0.6	23.5	20.4	42.4	12.2	0.6	32.1	15.4
N	117	117	116	117	116	80	80	37	37

Table 4: Explain exchange rate movement using Treasury basis innovation.

We regress the exchange rate movement on concurrent Treasury basis innovation. A higher exchange rate means a stronger base currency. For each non-U.S. country, we exclude the U.S. when we calculate its average Treasury basis and average exchange rate movement against other non-U.S. countries.

	Dependent variable:									
	Base Country's Exchange Rate Movement									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	U.S.	AUS	CAN	DEU	JPN	NZL	NOR	SWE	SWI	UK
$\Delta\bar{x}^{Treas}$	−9.62	−0.38	2.49	−5.88	3.19	−3.94	0.01	−0.97	1.76	2.64
	(2.05)	(3.38)	(1.61)	(3.87)	(4.83)	(1.80)	(0.95)	(0.85)	(1.48)	(2.32)
Constant	0.002	0.001	0.0002	0.0000	−0.002	0.0001	−0.001	−0.003	0.01	−0.004
	(0.004)	(0.005)	(0.004)	(0.003)	(0.01)	(0.01)	(0.003)	(0.003)	(0.003)	(0.004)
Observations	117	69	93	78	87	69	108	104	108	78
R^2	0.16	0.0002	0.03	0.03	0.01	0.07	0.0000	0.01	0.01	0.02

Table 5: Forecasting Currency Excess Returns in Panel Data

The dependent variable is the annualized nominal excess return (in logs) $rx^{fx}_{t\to t+k}$ on a long position in U.S. Treasuries and a short position (equal-weighted) in all foreign bonds over k quarters. The independent variables are the average Treasury basis, $\overline{x}^{Treas}$, the nominal Treasury yield difference $(y^{\$} - \overline{y}^{*})$, the change in the average Treasury basis $\Delta\overline{x}^{Treas}_{t}$ from $t-1$ to t, and a lag of the change in the average Treasury basis $\Delta\overline{x}^{Treas}_{t-1}$. Data is quarterly from 1988Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	$\overline{x}^{Treas}$	$y^{\$} - \overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2
			Panel A: 1988-2017		
3 months	-16.76	-0.17	-7.96	2.80	0.08
	(13.15)	(1.36)	(10.29)	(10.25)	
1 year	-0.63	0.28	-7.46	-2.93	0.04
	(8.72)	(0.96)	(5.64)	(3.93)	
2 years	4.72	0.35	-7.65	-5.24	0.05
	(5.07)	(0.58)	(3.27)	(2.48)	
3 years	8.12	0.49	-8.45	-5.22	0.12
	(3.99)	(0.39)	(2.74)	(2.07)	
			Panel B: 1988-2007		
3 months	-30.66	-0.37	5.79	14.55	0.13
	(15.04)	(1.47)	(9.51)	(10.52)	
1 year	-10.23	0.16	-5.55	-3.70	0.13
	(9.63)	(0.92)	(7.00)	(5.21)	
2 years	-2.33	0.18	-4.41	-4.31	0.06
	(5.51)	(0.56)	(4.12)	(3.04)	
3 years	5.14	0.42	-8.11	-5.50	0.09
	(4.83)	(0.37)	(3.50)	(2.51)	
			Panel C: 2007-2017		
3 months	17.94	-3.58	-32.05	-29.41	0.21
	(10.92)	(3.04)	(11.78)	(10.40)	
1 year	18.56	-2.41	-6.44	-3.43	0.22
	(7.11)	(2.57)	(6.19)	(3.27)	
2 years	20.71	0.35	-13.35	-9.06	0.46
	(3.47)	(0.82)	(2.21)	(2.22)	
3 years	13.87	0.24	-7.72	-5.73	0.33
	(3.21)	(0.76)	(2.10)	(1.58)	

Table 6: Forecasting 3-year Currency Excess Returns in Panel Data

In Panel A (B), the dependent variable is the annualized nominal excess return (in logs) $rx^{fx}_{t\to t+12}$ ($rx^{fx}_{t+4\to t+12}$) on a long position in U.S. Treasuries and a short position (equal-weighted) in all foreign bonds over k quarters. The independent variables are the Treasury basis, x^{Treas}, the nominal Treasury yield difference ($y^{\$} - \overline{y}^{*}$), the change in the Treasury basis $\Delta\overline{x}^{Treas}_{t}$ from $t-1$ to t, and a lag of the change in the Treasury basis $\Delta\overline{x}^{Treas}_{t-1}$. Data is quarterly from 1988Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	$\overline{x}^{Treas}$	$y^{\$} - \overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2	$\overline{x}^{Treas}$	$y^{\$} - \overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2
	Panel A: $rx^{fx}_{t\to t+12}$					*Panel B:* $rx^{fx}_{t+4\to t+12}$				
1988-2017	1.62				0.01	6.54				0.07
	(2.47)					(2.82)				
	2.26	0.30			0.02	7.12	0.27			0.07
	(2.46)	(0.33)				(2.63)	(0.33)			
	5.34	0.40	-5.38		0.07	9.89	0.45	-5.69		0.10
	(3.22)	(0.35)	(1.82)			(3.60)	(0.32)	(2.37)		
	8.12	0.49	-8.45	-5.22	0.12	13.01	0.61	-9.42	-6.75	0.15
	(3.99)	(0.39)	(2.74)	(2.07)		(4.60)	(0.34)	(3.70)	(2.84)	
1988-2007	-1.17				0.00	6.53				0.06
	(2.96)					(3.11)				
	-0.49	0.25			0.01	7.21	0.25			0.07
	(2.93)	(0.31)				(2.98)	(0.32)			
	2.30	0.33	-4.48		0.04	9.97	0.41	-5.38		0.09
	(3.98)	(0.33)	(2.22)			(4.25)	(0.33)	(2.90)		
	5.14	0.42	-8.11	-5.50	0.09	12.82	0.55	-9.38	-6.40	0.13
	(4.83)	(0.37)	(3.50)	(2.51)		(5.55)	(0.35)	(4.83)	(3.55)	
2007-2017	8.17				0.20	5.73				0.06
	(2.80)					(3.09)				
	8.17	-0.17			0.20	5.70	1.13			0.09
	(2.82)	(0.62)				(3.23)	(1.30)			
	10.67	0.03	-5.65		0.27	8.50	1.34	-6.33		0.14
	(2.88)	(0.71)	(2.21)			(3.22)	(1.33)	(2.56)		
	13.87	0.24	-7.72	-5.73	0.33	12.89	1.63	-9.16	-7.84	0.22
	(3.21)	(0.76)	(2.10)	(1.58)		(3.59)	(1.33)	(2.29)	(2.23)	

Table 7: Parameter Estimates

This table reports Maximum Likelihood parameter estimates for the exchange rate model with sticky expectations. The model is estimated on a sample of quarterly data between 1988 and 2016 on U.S. Treasury basis, the dollar real exchange rate and the real interest rate spread. We use equal-weighted averages for a basket of 9 countries. Panel A presents the parameter constraints imposed for identification. Panel B presents the coefficient estimates (with bootstrapped standard errors in parentheses). Standard errors are bootstrapped from 200 samples of the original sample size generated using the estimated model. Panel C reports the R^2 value for the observable U.S. Treasury basis.

Panel A: Coefficient Restrictions		
$\rho_{\lambda,rp}$	0.0000	
$\rho_{\beta,rp}$	0.0000	
δ_0	0.0000	
Panel B: Parameter Estimates		
ψ_1	0.8412	(0.0635)
ψ_0	-0.0004	(0.0003)
δ_1	0.9039	(0.0621)
γ_1	0.9013	(0.0630)
γ_0	0.0003	(0.0011)
κ_1	0.8765	(0.1741)
κ_0	0.1072	(0.1520)
σ^x	0.0000	(0.0001)
σ^λ	0.0062	(0.0048)
σ^β	0.0380	(0.0266)
σ^{rp}	0.0051	(0.0067)
σ^i	0.0022	(0.0002)
$\rho_{\lambda,\beta}$	0.5439	(0.2819)
$\rho_{\lambda,i}$	0.1398	(0.1118)
$\rho_{\beta,i}$	-0.0062	(0.2574)
$\rho_{rp,i}$	-0.0063	(0.1388)
c	-0.0033	(0.0108)
ω	0.4168	(0.0936)
Panel C: Goodness of Fit		
Treasury Basis	81.01%	

Table 8: Testing Parameter Restrictions

This table tests parameter restrictions for the exchange rate model with sticky expectations. The model in the second column imposes rational expectations: $\varpi = 0$. The model in the third column imposes a constant beta: $\kappa_1 = \sigma^\beta = \rho_{\lambda,\beta} = \rho_{\beta,i} = 0$. Panel A reports the LR test and the p-values. Panel B reports the parameter estimates for restricted coefficients. The model is estimated on a sample of quarterly data between 1988 and 2016 on U.S. Treasury basis, the dollar real exchange rate and the real interest rate spread. We use equal-weighted averages for a basket of 9 countries.

Panel A: Test Statistics			
Statistics	Benchmark	Model 1	Model 2
$\mathcal{L}$	1733.1473	1721.295116	1714.8674
Likelihood Ratio		23.7044	36.5598
p Value		0.0000	0.0000
Panel B: Restricted Parameter Estimates			
Parameter	Benchmark	Model 1	Model 2
σ^β	0.0380		0.0000
$\rho_{\lambda,\beta}$	0.5439		0.0000
$\rho_{\beta,i}$	-0.0062		0.0000
ϖ	0.4168	0.0000	

Table 9: Moments of Latent State Variables and Observables.

Table reports the mean, the standard deviation, and the autocorrelation of the filtered state variables and the observables in model with sticky expectations. The model is estimated on a sample of quarterly data between 1988 and 2016 on U.S. Treasury basis, the dollar real exchange rate and the real interest rate spread. We use equal-weighted averages for a basket of 9 countries.

Variable	Mean	Std Dev	Autocorr
Panel A: State Variables			
x_t^λ (%)	10.34	10.37	0.84
y_t^i (%)	-0.11	4.16	0.93
λ_t (%)	0.90	0.72	0.70
β_t^*	0.93	0.05	0.81
rp_t (%)	0.55	0.99	0.88
Panel B: Observables			
x_t^{Treas} (%)	-0.05	0.05	0.56
i_t (%)	-0.28	0.45	0.87
Δq_t (%)	0.27	4.71	0.00
q_t	0.17	0.13	0.93

Table 10: Variance Decomposition of Real Exchange Rate

Variance Decomposition of Real Exchange Rate in model with sticky expectations. Panel A reports the variance decomposition of quarterly real exchange rate innovations into news about current and future convenience yields earned by foreign investors N_t^λ, interest rates N_t^i, and risk premia N_t^{rp}. Panel B reports the variance decomposition of quarterly real exchange rate in levels into convenience yields L_t^λ, interest rates L_t^i, and risk premia L_t^{rp}. The model with sticky expectations is estimated on a sample of quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Equal-weighted averages for a basket of 9 countries.

			Panel A: RER Innovations			
	$var(N_t^\lambda)$	$var(N_t^i)$	$var(N_t^{rp})$	$cov(N_t^\lambda, N_t^i)$	$cov(N_t^\lambda, N_t^{rp})$	$cov(N_t^i, N_t^{rp})$
Value (bps.)	13.25	0.65	9.75	0.82	0.00	-0.03
Percentage (%)	54.20	2.68	39.88	3.37	0.00	-0.13
			Panel B: RER Levels			
	$var(L_t^\lambda)$	$var(L_t^i)$	$var(L_t^{rp})$	$2cov(L_t^\lambda, L_t^i)$	$2cov(L_t^\lambda, L_t^{rp})$	$2cov(L_t^i, L_t^{rp})$
Value (bps.)	166.41	4.69	126.19	3.77	0.00	-0.15
Percentage (%)	55.30	1.56	41.94	1.25	0.00	-0.05

Table 11: Accounting for Real Exchange Rate Variation

Time series regression of Δq_t on regressors $\Delta\overline{x}_t^{Treas}$, and the filtered convenience yield innovations: ε_t^λ and lagged ε_t^λ. Standard errors (in parentheses) are HAC consistent. The model with sticky expectations is estimated on quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries.

	Full Sample			Before 2008			From 2008		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta\overline{x}_t^{Treas}$	-9.28			-8.20			-12.49		
	(2.66)			(3.39)			(3.35)		
ε_t^λ		7.91	8.26		8.35	8.78		6.97	7.06
		(0.44)	(0.56)		(0.41)	(0.52)		(1.01)	(1.18)
lag ε_t^λ			2.38			2.79			1.56
			(0.51)			(0.61)			(0.57)
Constant	0.003	-0.004	-0.01	0.001	-0.01	-0.01	0.01	0.001	0.001
	(0.004)	(0.003)	(0.003)	(0.01)	(0.003)	(0.003)	(0.01)	(0.004)	(0.003)
Observations	112	112	111	78	78	77	34	34	33
R^2	0.16	0.67	0.73	0.12	0.66	0.74	0.35	0.73	0.77

Table 12: Loadings of Optimal Forecast on the Latent Variables.

Loadings of optimal forecast $\mathbb{E}_t(q_{t+k} - q_t)$ on latent variables. The model with sticky expectations is estimated on a sample of quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries.

horizon j	λ_t^*	i_t	rp_t	x_t^λ	y_t^i	z_t^{rp}
1 quarter	5.33	3.09	5.48	-0.36	-0.38	-0.36
4 quarters	6.49	3.06	6.74	-0.57	-0.57	-0.57
8 quarters	4.41	1.58	4.63	-0.58	-0.58	-0.58
12 quarters	2.91	0.79	3.09	-0.58	-0.58	-0.58

Table 13: In-Sample Predictability of Real Exchange Rates

Panel A reports the regression of the dollar's future real exchange rate change on $\mathbb{E}_t(q_{t+k} - q_t)$ produced by the Kalman filter. Panel B controls for the Treasury basis, the innovation in the Treasury basis, and the interest rate differential. Both dependent and explanatory variables are annualized. Standard errors are HAC consistent. The model with sticky expectations is estimated on quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries.

	3 months	1 year	2 years	3 years
	(1)	(2)	(3)	(4)
		Panel A		
$\mathbb{E}_t(q_{t+1} - q_t) * 4$	−0.04			
	(0.17)			
$\mathbb{E}_t(q_{t+4} - q_t)$		0.19		
		(0.28)		
$\mathbb{E}_t(q_{t+8} - q_t)/2$			0.80	
			(0.33)	
$\mathbb{E}_t(q_{t+12} - q_t)/3$				0.80
				(0.29)
Observations	111	108	104	100
R^2	0.001	0.01	0.15	0.18
		Panel B		
$\mathbb{E}_t(q_{t+1} - q_t) * 4$	−0.12			
	(0.18)			
$\mathbb{E}_t(q_{t+4} - q_t)$		0.14		
		(0.32)		
$\mathbb{E}_t(q_{t+8} - q_t)/2$			0.80	
			(0.37)	
$\mathbb{E}_t(q_{t+12} - q_t)/3$				0.81
				(0.36)
$\overline{x}^{Treas}$	−51.74	16.25	15.53	26.36
	(63.55)	(41.82)	(22.34)	(14.10)
$\Delta\overline{x}^{Treas}$	−8.77	−11.27	−6.19	−7.47
	(17.24)	(6.73)	(2.44)	(1.93)
$r^{\$} - r^*$	−0.82	−0.32	0.11	0.36
	(1.36)	(1.03)	(0.79)	(0.57)
Observations	111	108	104	100
R^2	0.06	0.05	0.17	0.22

Table 14: Out-of-Sample Predictability of Real Exchange Rates

We report regressions of the dollar's future real exchange rate change on $\mathbb{E}_t(q_{t+k} - q_t)$ produced by the Kalman filter. Panel A reports in-sample results for 1988Q3—2002Q2. Panel B reports out-of-sample results for 2002Q3—2016Q2. Both dependent and explanatory variables are annualized. We estimate the model with sticky expectations using only the first half of the sample: 1988Q3—2002Q2. Data is based on equal-weighted averages for a basket of 9 countries.

Panel A: 1988Q3—2002Q2				
	(1)	(2)	(3)	(4)
$\mathbb{E}_t(q_{t+1} - q_t) * 4$	0.05			
	(0.13)			
$\mathbb{E}_t(q_{t+4} - q_t)$		0.27		
		(0.26)		
$\mathbb{E}_t(q_{t+8} - q_t)/2$			1.12	
			(0.30)	
$\mathbb{E}_t(q_{t+12} - q_t)/3$				1.11
				(0.34)
Observations	56	53	49	45
R^2	0.002	0.02	0.26	0.28

Panel B: 2002Q3—2016Q2				
	(1)	(2)	(3)	(4)
$\mathbb{E}_t(q_{t+1} - q_t) * 4$	0.21			
	(0.12)			
$\mathbb{E}_t(q_{t+4} - q_t)$		0.41		
		(0.21)		
$\mathbb{E}_t(q_{t+8} - q_t)/2$			0.59	
			(0.20)	
$\mathbb{E}_t(q_{t+12} - q_t)/3$				0.50
				(0.22)
Observations	55	52	48	44
R^2	0.07	0.14	0.33	0.33

Table 15: Accounting for Real Exchange Rate Variation, Out of Sample

Time series regression of Δq_t on regressors $\Delta \overline{x}_t^{Treas}$, and the filtered convenience yield innovations: ε_t^{λ} and lagged ε_t^{λ}. Standard errors (in parentheses) are HAC consistent. We estimate the model with sticky expectations using only the first half of the sample: 1988Q3—2002Q2. Data is based on equal-weighted averages for a basket of 9 countries.

	1988Q3—2002Q2			2002Q3—2016Q2		
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta \overline{x}_t^{Treas}$	-7.62			-12.98		
	(3.65)			(3.00)		
ε_t^{λ}		6.71	7.11		14.15	15.22
		(1.82)	(1.81)		(3.03)	(3.45)
lag ε_t^{λ}			3.83			8.31
			(2.13)			(3.25)
Observations	56	56	55	56	56	55
R^2	0.11	0.22	0.30	0.28	0.17	0.25

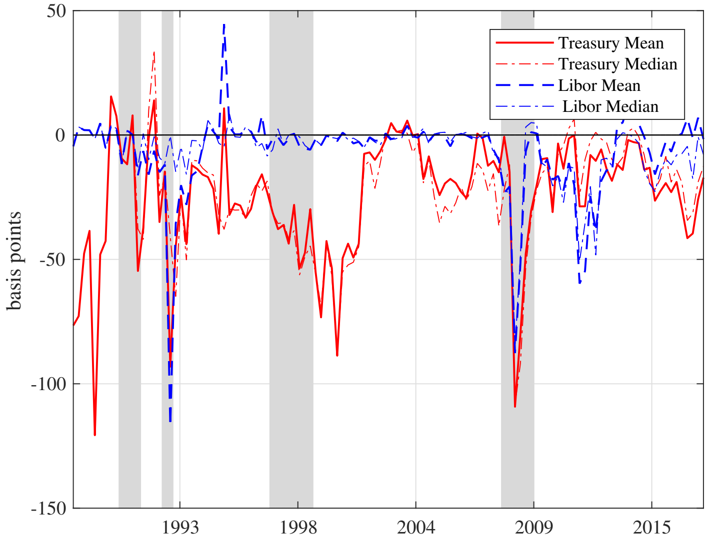

Figure 1: U.S. LIBOR and Treasury Bases

U.S. LIBOR and Treasury basis in basis points from 1988Q1 to 2017Q2. The maturity is one year. We plot the cross-sectional mean and median for each of the bases.

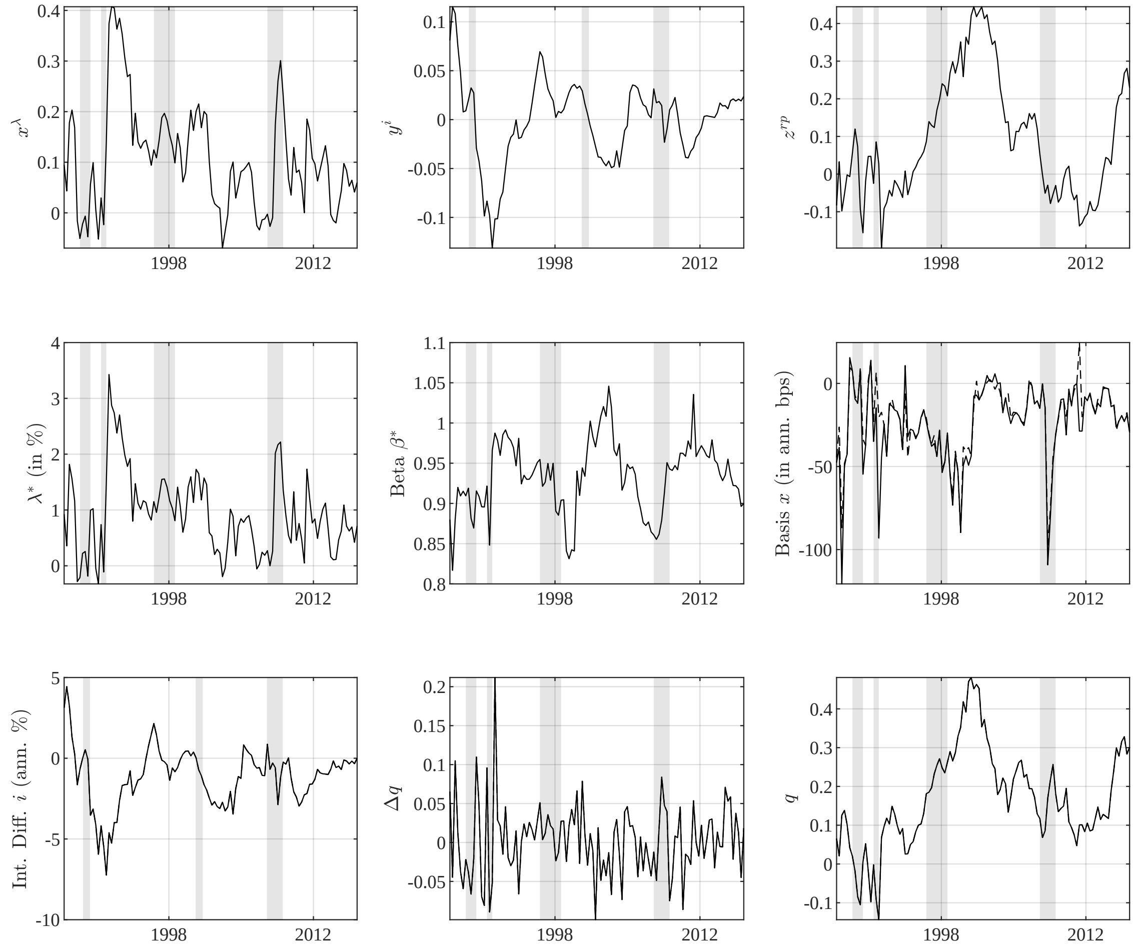

Figure 2: Filtered Latent Variables

Filtered Series for Latent Variables. The exchange rate model with sticky expectations is estimated on a sample of quarterly data between 1988 and 2016 on Treasury basis, real exchange rates and real interest rates. Equal-weighted averages for a basket of 9 countries.

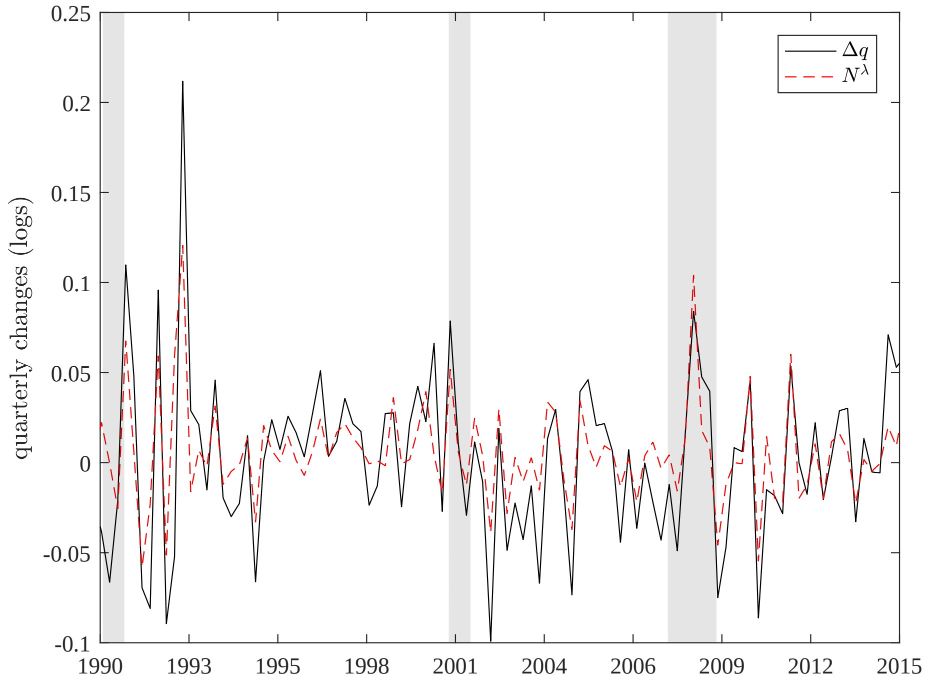

Figure 3: Real Exchange Rate Innovation and News about Convenience Yields

We plots news about current and future convenience yields N_t^{λ} and real exchange rate changes Δq_t. The model with sticky expectations is estimated on quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries.

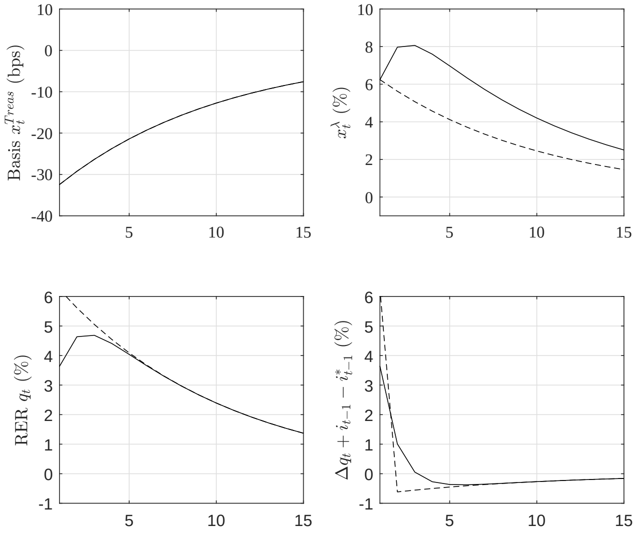

Figure 4: Impulse Response to Convenience Yield Shock

The solid line (dotted line) figure plots the impulse responses for the exchange rate model with sticky expectations (rational expectations) to a one-standard deviation shock to ε_t^λ. The model is estimated from quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries. The rational expectations impulse responses are computed by setting $\varpi = 0$.

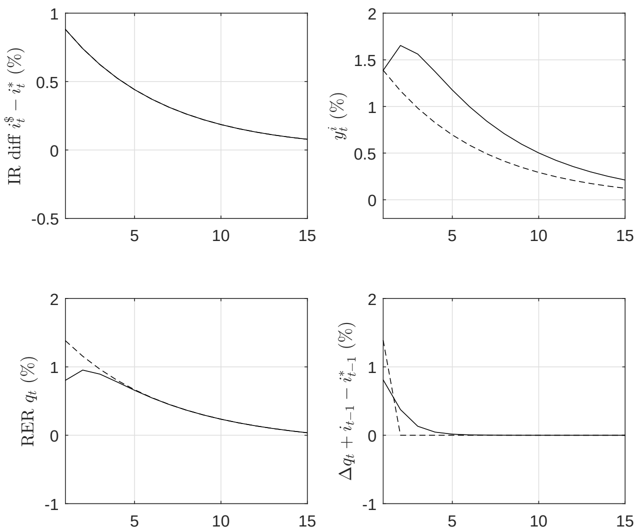

Figure 5: Impulse Response to Interest Rate Shock

The solid line (dotted line) figure plots the impulse responses for the exchange rate model with sticky expectations (rational expectations) to a shock to ϵ_t^i. The model with sticky expectations is estimated on a sample of quarterly data between 1988 and 2016 on the U.S. Treasury basis, the real exchange rate and the real interest rate difference. Data is based on equal-weighted averages for a basket of 9 countries. The rational expectations impulse responses are computed by setting $\varpi = 0$.

Appendix

A Proofs

- Proof of Proposition 2:

Proof. We consider an environment with a continuum of currencies/investor pairs. For a foreign investor who had last updated k periods ago, his valuation of the real exchange rate can be written as:

$$q_t(k) = \mathbb{E}_{t-k}\sum_{\tau=0}^{\infty}\lambda^*_{t+\tau} + \mathbb{E}_{t-k}\sum_{\tau=0}^{\infty}(r^{\$}_{t+\tau} - r^*_{t+\tau}) - \mathbb{E}_{t-k}\sum_{\tau=0}^{\infty}\left(RP^*_{t+\tau} - \frac{1}{2}Var[\Delta s_{t+\tau+1}]\right) + \bar{q}. \tag{39}$$

where, $\bar{q} = \mathbb{E}_t[\lim_{\tau\to\infty}\ q_{t+\tau}]$ is constant under the assumption that the real exchange rate is stationary. The terms $r^{\$}_t$ and r^*_t are the real interest rates, i.e., $y^{\$}_t - \mathbb{E}_t[\Delta p^{\$}_{t+1}]$ is the real dollar interest rate.

We posit that the real dollar exchange rate is equal to the average valuation of the real dollar exchange rate across investors. In the case of information stickiness, the convenience yield component is given by:

$$\mathbb{F}_t\left[\sum_{k=0}^{\infty}(\lambda^*_{t+k} - \theta_\lambda)\right] = \sum_{k=0}^{\infty}(1-\varpi)\sum_{j=0}^{\infty}\varpi^j\gamma_1^{j+k}(\lambda^*_{t-j} - \theta_\lambda),$$

which can be simplified as

$$\mathbb{F}_t\left[\sum_{k=1}^{\infty}(\lambda^*_{t+k} - \theta_\lambda)\right] = \sum_{j=0}^{\infty}(\varpi)^j(1-\varpi)\frac{\gamma_1^j}{1-\gamma_1}(\lambda^*_{t-j} - \theta_\lambda).$$

By the same token, the aggregate cash flow and risk premium components are given by:

$$\mathbb{F}_t\left[\sum_{k=1}^{\infty}(i_{t+k} - \theta_i)\right] = \sum_{j=0}^{\infty}(\varpi)^j(1-\varpi)\frac{\psi_1^j}{1-\psi_1}(i_{t-j} - \theta_i),$$

$$\mathbb{F}_t\left[\sum_{k=1}^{\infty}(rp_{t+k} - \theta_{rp})\right] = \sum_{j=0}^{\infty}(\varpi)^j(1-\varpi)\frac{\delta_1^j}{1-\delta_1}(rp_{t-j} - \theta_{rp}).$$

As a result, we end up with the following expression for the log of the real exchange rate:

$$q_t = \bar{q} + (1-\varpi)\sum_{j=0}^{\infty}\left(\frac{\varpi^j\gamma_1^j}{1-\gamma_1}(\lambda^*_{t-j} - \theta_\lambda) + \frac{\varpi^j\psi_1^j}{1-\psi_1}(i_{t-j} - \theta_i) + \frac{\varpi^j\delta_1^j}{1-\delta_1}(rp_{t-j} - \theta_{rp})\right).$$

□

- Proof of Proposition 3:

Proof. Armed with this expression, we can compute the (rationally) expected change in the real exchange given by:

$$\begin{aligned}
\mathbb{E}_t(q_{t+1}-q_t) &= \bar{q} + (1-\varpi)\mathbb{E}_t \sum_{j=0}^{\infty}\left(\frac{\varpi^j\gamma_1^j}{1-\gamma_1}(\lambda^*_{t+1-j}-\theta_\lambda) + \frac{\varpi^j\psi_1^j}{1-\psi_1}(i_{t+1-j}-\theta_i) + \frac{\varpi^j\delta_1^j}{1-\delta_1}(rp_{t+1-j}-\theta_{rp})\right) \\
&- \left[\bar{q} + (1-\varpi)\sum_{j=0}^{\infty}\left(\frac{\varpi^j\gamma_1^j}{1-\gamma_1}(\lambda^*_{t-j}-\theta_\lambda) + \frac{\varpi^j\psi_1^j}{1-\psi_1}(i_{t-j}-\theta_i) + \frac{\varpi^j\delta_1^j}{1-\delta_1}(rp_{t-j}-\theta_{rp})\right)\right].
\end{aligned}$$

This can be simplified as follows:

$$\begin{aligned}
\mathbb{E}_t(q_{t+1}-q_t) &= (1-\varpi)\left(\frac{\gamma_1}{1-\gamma_1}(\lambda^*_t-\theta_\lambda) + \frac{\psi_1}{1-\psi_1}(i_t-\theta_i) + \frac{\delta_1}{1-\delta_1}(rp_t-\theta_{rp})\right) \\
&+ (1-\varpi)\sum_{j=1}^{\infty}\left(\frac{\varpi^j\gamma_1^j-\varpi^{j-1}\gamma_1^{j-1}}{1-\gamma_1}(\lambda^*_{t+1-j}-\theta_\lambda) + \frac{\varpi^j\psi_1^j-\varpi^{j-1}\psi_1^{j-1}}{1-\psi_1}(i_{t+1-j}-\theta_i) + \frac{\varpi^j\delta_1^j-\varpi^{j-1}\delta_1^{j-1}}{1-\delta_1}(rp_{t+1-j}-\theta_{rp})\right),
\end{aligned}$$

or, equivalently,

$$\begin{aligned}
\mathbb{E}_t(q_{t+1}-q_t) &= (1-\varpi)\left(\frac{\gamma_1}{1-\gamma_1}(\lambda^*_t-\theta_\lambda) + \frac{\psi_1}{1-\psi_1}(i_t-\theta_i) + \frac{\delta_1}{1-\delta_1}(rp_t-\theta_{rp})\right) \\
&+ (1-\varpi)\sum_{j=0}^{\infty}\left((\varpi^j\gamma_1^j)\frac{\varpi\gamma_1-1}{1-\gamma_1}(\lambda^*_{t-j}-\theta_\lambda) + (\varpi^j\psi_1^j)\frac{\varpi\psi_1-1}{1-\psi_1}(i_{t-j}-\theta_i) + (\varpi^j\delta_1^j)\frac{\varpi\delta_1-1}{1-\delta_1}(rp_{t-j}-\theta_{rp})\right).
\end{aligned}$$

Using the expression for the new state variables, $x_t^\lambda = \varpi\gamma_1 x_{t-1}^\lambda + \frac{1}{1-\gamma_1}(\lambda_t^*-\theta_\lambda)$, $y_t^i = \varpi\psi_1 y_{t-1}^i + \frac{1}{1-\psi_1}(i_t-\theta_i)$, $z_t^{rp} = \varpi\delta_1 z_{t-1}^{rp} + \frac{1}{1-\delta_1}(rp_t-\theta_{rp})$, we obtain the following expression for the (rational) expected rate of appreciation:

$$\begin{aligned}
\mathbb{E}_t(q_{t+1}-q_t) &= (1-\varpi)\left(\frac{\gamma_1}{1-\gamma_1}(\lambda^*_t-\theta_\lambda) + \frac{\psi_1}{1-\psi_1}(i_t-\theta_i) + \frac{\delta_1}{1-\delta_1}(rp_t-\theta_{rp})\right) \\
&+ (1-\varpi)\left((\varpi\gamma_1-1)x_t^\lambda + (\varpi\psi_1-1)y_t^i + (\varpi\delta_1-1)z_t^{rp}\right)
\end{aligned}$$

□

- Proof of Corollary 2:

Proof. We can compute the (rationally) expected change in the real exchange given by:

$$\begin{aligned}
\mathbb{E}_t(q_{t+k}-q_t) &= \bar{q} + (1-\varpi)\mathbb{E}_t \sum_{j=0}^{\infty}\left(\frac{\varpi^j\gamma_1^j}{1-\gamma_1}(\lambda^*_{t+k-j}-\theta_\lambda) + \frac{\varpi^j\psi_1^j}{1-\psi_1}(i_{t+k-j}-\theta_i) + \frac{\varpi^j\delta_1^j}{1-\delta_1}(rp_{t+k-j}-\theta_{rp})\right) \\
&- \left[\bar{q} + (1-\varpi)\sum_{j=0}^{\infty}\left(\frac{\varpi^j\gamma_1^j}{1-\gamma_1}(\lambda^*_{t-j}-\theta_\lambda) + \frac{\varpi^j\psi_1^j}{1-\psi_1}(i_{t-j}-\theta_i) + \frac{\varpi^j\delta_1^j}{1-\delta_1}(rp_{t-j}-\theta_{rp})\right)\right].
\end{aligned}$$

This can be simplified as follows:

$$
\begin{aligned}
\mathbb{E}_t(q_{t+k}-q_t) &= (1-\varpi)\sum_{j=0}^{k-1}\left(\frac{\varpi^j\gamma_1^k}{1-\gamma_1}(\lambda_t^*-\theta_\lambda)+\frac{\varpi^j\psi_1^k}{1-\psi_1}(i_t-\theta_i)+\frac{\varpi^j\delta_1^k}{1-\delta_1}(rp_t-\theta_{rp})\right)\\
&+ (1-\varpi)\sum_{j=k}^{\infty}\left(\frac{\varpi^j\gamma_1^j-\varpi^{j-k}\gamma_1^{j-k}}{1-\gamma_1}(\lambda_{t+k-j}^*-\theta_\lambda)+\frac{\varpi^j\psi_1^j-\varpi^{j-k}\psi_1^{j-k}}{1-\psi_1}(i_{t+k-j}-\theta_i)+\frac{\varpi^j\delta_1^j-\varpi^{j-k}\delta_1^{j-k}}{1-\delta_1}(rp_{t+k-j}-\theta_{rp})\right),
\end{aligned}
$$

or, equivalently,

$$
\begin{aligned}
\mathbb{E}_t(q_{t+k}-q_t) &= (1-\varpi)\sum_{j=0}^{k-1}\varpi^j\left(\frac{\gamma_1^k}{1-\gamma_1}(\lambda_t^*-\theta_\lambda)+\frac{\psi_1^k}{1-\psi_1}(i_t-\theta_i)+\frac{\delta_1^k}{1-\delta_1}(rp_t-\theta_{rp})\right)\\
&+ (1-\varpi)\sum_{j=k}^{\infty}\left(\varpi^k\gamma_1^k\frac{\varpi^{j-k}\gamma_1^{j-k}-1}{1-\gamma_1}(\lambda_{t+k-j}^*-\theta_\lambda)+\varpi^k\psi_1^k\frac{\varpi^{j-k}\psi_1^{j-k}-1}{1-\psi_1}(i_{t+k-j}-\theta_i)+\varpi^k\delta_1^k\frac{\varpi^{j-k}\delta_1^{j-k}-1}{1-\delta_1}(rp_{t+k-j}-\theta_{rp})\right)
\end{aligned}
$$

Using the expression for the new state variables, $x_t^\lambda = \varpi\gamma_1 x_{t-1}^\lambda + \frac{1}{1-\gamma_1}(\lambda_t^*-\theta_\lambda)$, $y_t^i = \varpi\psi_1 y_{t-1}^i + \frac{1}{1-\psi_1}(i_t-\theta_i)$, $z_t^{rp} = \varpi\delta_1 z_{t-1}^{rp} + \frac{1}{1-\delta_1}(rp_t-\theta_{rp})$,

we obtain the following expression for the (rational) expected rate of appreciation:

$$
\begin{aligned}
\mathbb{E}_t(q_{t+k}-q_t) &= (1-\varpi^k)\left(\frac{\gamma_1^k}{1-\gamma_1}(\lambda_t^*-\theta_\lambda)+\frac{\psi_1^k}{1-\psi_1}(i_t-\theta_i)+\frac{\delta_1^k}{1-\delta_1}(rp_t-\theta_{rp})\right)\\
&+ (1-\varpi)\left((\varpi^k\gamma_1^k-1)x_t^\lambda+(\varpi^k\psi_1^k-1)y_t^i+(\varpi^k\delta_1^k-1)z_t^{rp}\right)
\end{aligned}
$$

□

B Kalman Filter

The change in the log real exchange rate is given by:

$$
\begin{aligned}
q_{t+1}-q_t &= (1-\varpi)(x_{t+1}^\lambda-x_t^\lambda+y_{t+1}^i-y_t^i+z_{t+1}^{rp}-z_t^{rp})\\
&= (1-\varpi)\left((\varpi\gamma_1-1)x_t^\lambda+\frac{\lambda_{t+1}-\frac{\gamma_0}{1-\gamma_1}}{1-\gamma_1}+(\varpi\psi_1-1)y_t^i+\frac{i_{t+1}-\frac{\psi_0}{1-\psi_1}}{1-\psi_1}+(\varpi\delta_1-1)z_t^{rp}+\frac{rp_{t+1}-\frac{\delta_0}{1-\delta_1}}{1-\delta_1}\right).
\end{aligned}
$$

Following Van Binsbergen and Koijen (2010), we substitute out the risk premium state variable z_t^{rp} in the deterministic observation equation for the real exchange rate, to get rid of one transition equation. As a result, we obtain:

$$
\begin{aligned}
(q_{t+1}-q_t) &= \varpi\delta_1(q_t-q_{t-1})+c_1+(1-\varpi)\left((\varpi\gamma_1-1)(x_t^\lambda-\varpi\delta_1x_{t-1}^\lambda)+\frac{\lambda_{t+1}-\varpi\delta_1\lambda_t}{1-\gamma_1}\right.\\
&+ \left.(\varpi\psi_1-1)(y_t^i-\varpi\delta_1y_{t-1}^i)+\frac{i_{t+1}-\varpi\delta_1i_t}{1-\psi_1}+(\varpi\delta_1-1)(z_t^{rp}-\varpi\delta_1z_{t-1}^{rp})+\frac{rp_{t+1}-\varpi\delta_1rp_t}{1-\delta_1}\right)\\
&= \varpi\delta_1(q_t-q_{t-1})+c+(1-\varpi)\left((\varpi\gamma_1-1)(x_t^\lambda-\varpi\delta_1x_{t-1}^\lambda)+\frac{(\gamma_1-\varpi\delta_1)\lambda_t+\varepsilon_{t+1}^\lambda}{1-\gamma_1}\right.\\
&+ \left.(\varpi\psi_1-1)(y_t^i-\varpi\delta_1y_{t-1}^i)+\frac{(\psi_1-\varpi\delta_1)i_t+\varepsilon_{t+1}^i}{1-\psi_1}+\frac{(\delta_1-1)rp_t+\varepsilon_{t+1}^{rp}}{1-\delta_1}\right).
\end{aligned}
$$

To summarize, the *observation equations* are given by:

$$
\begin{aligned}
x_t^{Treas} &= (\kappa_0 + \kappa_1 \beta_{t-1}^* + \varepsilon_t^\beta - 1)(\gamma_0 + \gamma_1 \lambda_{t-1} + \varepsilon_t^\lambda) + \varepsilon_t^x, \\
i_t &= \psi_0 + \psi_1 i_{t-1} + \varepsilon_t^i, \\
\Delta q_t &= \varpi \delta_1 \Delta q_{t-1} + c + (1-\varpi)\Bigg((\varpi\gamma_1 - 1)(x_{t-1}^\lambda - \varpi\delta_1 x_{t-2}^\lambda) + \frac{(\gamma_1 - \varpi\delta_1)\lambda_{t-1} + \varepsilon_t^\lambda}{1-\gamma_1} \\
&+ (\varpi\psi_1 - 1)(y_{t-1}^i - \varpi\delta_1 y_{t-2}^i) + \frac{(\psi_1 - \varpi\delta_1) i_{t-1} + \varepsilon_t^i}{1-\psi_1} - rp_{t-1} + \frac{\varepsilon_t^{rp}}{1-\delta_1}\Bigg),
\end{aligned}
$$

and the *transition equations* are given by:

$$
\begin{aligned}
x_t^\lambda &= \varpi\gamma_1 x_{t-1}^\lambda + \frac{1}{1-\gamma_1}\left(\gamma_0 + \gamma_1\lambda_{t-1} + \varepsilon_t^\lambda - \frac{\gamma_0}{1-\gamma_1}\right), \\
y_t^i &= \varpi\psi_1 y_{t-1}^i + \frac{1}{1-\psi_1}\left(i_t - \frac{\psi_0}{1-\psi_1}\right), \\
\lambda_t &= \gamma_0 + \gamma_1\lambda_{t-1} + \varepsilon_t^\lambda, \\
\beta_t^* &= \kappa_0 + \kappa_1\beta_{t-1}^* + \varepsilon_t^\beta, \\
rp_t &= 0 + \delta_1 rp_{t-1} + \varepsilon_t^{rp}.
\end{aligned}
$$

This system has the following state space representation. For $t = 1, \ldots, T$,

$$
\begin{aligned}
\boldsymbol{Y}_t &= \boldsymbol{Z}_t + h(\boldsymbol{X}_t) + \boldsymbol{e}_t, \boldsymbol{e}_t \sim N(\boldsymbol{0}, \boldsymbol{H}_t), \\
\boldsymbol{X}_{t+1} &= \boldsymbol{A}_t + \boldsymbol{F}\boldsymbol{X}_t + \boldsymbol{\Gamma}\boldsymbol{\eta}_{t+1}, \boldsymbol{\eta}_{t+1} \sim N(\boldsymbol{0}, \boldsymbol{Q}_t),
\end{aligned}
$$

where $\boldsymbol{Y}_t$ is a 3-dimensional observation vector and $\boldsymbol{X}_t$ is a 9-dimensional state vector:

$$
\boldsymbol{Y}_t' = \begin{bmatrix} x_t^{Treas} & i_t & \Delta q_t \end{bmatrix}, \boldsymbol{X}_t' = \begin{bmatrix} x_{t-1}^\lambda & y_{t-1}^i & x_{t-2}^\lambda & y_{t-2}^i & \lambda_{t-1} & \beta_{t-1} & rp_{t-1} & \varepsilon_t^\lambda & \varepsilon_t^\beta & \varepsilon_t^{rp} & \varepsilon_t^i \end{bmatrix}
$$

The other matrices are given by:

$$
\begin{aligned}
\boldsymbol{Z}_t &= \begin{bmatrix} 0 \\ \psi_0 + \psi_1 i_{t-1} \\ c + \varpi\delta_1\Delta q_{t-1} + (1-\varpi)\frac{(\psi_1 - \varpi\delta_1)}{1-\psi_1} i_{t-1} \end{bmatrix}, \boldsymbol{e}_t = \begin{bmatrix} \varepsilon_t^x \\ 0 \\ 0 \end{bmatrix}, \\
h(\boldsymbol{X}_t) &= \begin{bmatrix} (\kappa_0 + \kappa_1\beta_{t-1}^* + \varepsilon_t^\beta - 1)(\gamma_0 + \gamma_1\lambda_{t-1} + \varepsilon_t^\lambda) \\ \varepsilon_t^i \\ (1-\varpi)\Big((\varpi\gamma_1 - 1)(x_{t-1}^\lambda - \varpi\delta_1 x_{t-2}^\lambda) + \frac{(\gamma_1 - \varpi\delta_1)\lambda_{t-1} + \varepsilon_t^\lambda}{1-\gamma_1} \\ + (\varpi\psi_1 - 1)(y_{t-1}^i - \varpi\delta_1 y_{t-2}^i) + \frac{\varepsilon_t^i}{1-\psi_1} - rp_{t-1} + \frac{\varepsilon_t^{rp}}{1-\delta_1}\Big) \end{bmatrix},
\end{aligned}
$$

$$
\boldsymbol{A}_t = \begin{bmatrix} \frac{-\gamma_0\gamma_1}{(1-\gamma_1)^2} \\ \frac{1}{1-\psi_1}\left(i_t - \frac{\psi_0}{1-\psi_1}\right) \\ 0 \\ 0 \\ \gamma_0 \\ \kappa_0 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \boldsymbol{F} = \begin{bmatrix} \varpi\gamma_1 & 0 & 0 & 0 & \frac{\gamma_1}{1-\gamma_1} & 0 & 0 & \frac{1}{1-\gamma_1} & 0 & 0 & 0 \\ 0 & \varpi\psi_1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \gamma_1 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & \kappa_1 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & \delta_1 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix},
$$

$$
\boldsymbol{\Gamma} = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}, \boldsymbol{\eta}_{t+1} = \begin{bmatrix} \varepsilon_{t+1}^{\lambda} \\ \varepsilon_{t+1}^{\beta} \\ \varepsilon_{t+1}^{rp} \\ \varepsilon_{t+1}^{i} \end{bmatrix}.
$$

The observation noise ε_{t-1}^{x} has covariance $(\sigma^x)^2$. The structural shocks $\boldsymbol{\eta}_{t+1}$ have the following covariance matrix

$$
var(\boldsymbol{\eta}_{t+1}) = \begin{bmatrix} (\sigma^\lambda)^2 & \sigma^\lambda\sigma^\beta\rho_{\lambda,\beta} & \sigma^\lambda\sigma^{rp}\rho_{\lambda,rp} & \sigma^\lambda\sigma^i\rho_{\lambda,i} \\ \sigma^\lambda\sigma^\beta\rho_{\lambda,\beta} & (\sigma^\beta)^2 & \sigma^\beta\sigma^{rp}\rho_{\beta,rp} & \sigma^\beta\sigma^i\rho_{\beta,i} \\ \sigma^\lambda\sigma^{rp}\rho_{\lambda,rp} & \sigma^\beta\sigma^{rp}\rho_{\beta,rp} & (\sigma^{rp})^2 & \sigma^{rp}\sigma^i\rho_{rp,i} \\ \sigma^\lambda\sigma^i\rho_{\lambda,i} & \sigma^\beta\sigma^i\rho_{\beta,i} & \sigma^{rp}\sigma^i\rho_{rp,i} & (\sigma^i)^2 \end{bmatrix}
$$

B.1 Identification

In this system, the risk premium term and the convenience yield term are not identified without additional restrictions. The Kalman filter maximizes the likelihood of the innovations to the observable variables condition on the information in the previous period:

$$\begin{aligned}(\mathbb{E}_t - \mathbb{E}_{t-1})x_t^{Treas} &= b_0 + b_{1,t-1}\varepsilon_t^{\beta} + b_{2,t-1}\varepsilon_t^{\lambda} + \varepsilon_t^{x}, \\ (\mathbb{E}_t - \mathbb{E}_{t-1})i_t &= \varepsilon_t^{i}, \\ (\mathbb{E}_t - \mathbb{E}_{t-1})\Delta q_t &= c_0 + c_1\varepsilon_t^{\lambda} + c_2\varepsilon_t^{i} + c_3\varepsilon_t^{rp}.\end{aligned}$$

Let us suppose we have found a set of latent variables $(\varepsilon_t^{\lambda}, \varepsilon_t^{i}, \varepsilon_t^{rp}, \varepsilon_t^{\beta})$ that satisfy these equations. Consider the following perturbation $\tilde{\varepsilon}$. We can choose innovations: $\tilde{\varepsilon}_t^{\lambda} = \varepsilon_t^{\lambda} + \tilde{\varepsilon}$, $\tilde{\varepsilon}_t^{\beta} = \varepsilon_t^{\beta} - \frac{b_{2,t-1}}{b_{1,t-1}}\tilde{\varepsilon}$, $\tilde{\varepsilon}_t^{rp} = \varepsilon_t^{rp} - \frac{c_1}{c_3}\tilde{\varepsilon}$, and obtain observationally equivalent results. However, different $\tilde{\varepsilon}$ will give rise to a different covariance matrix of the structural shocks. So, we can take an a priori stance on one of the correlation parameters in order to identify the structural shocks. Similarly, we can construct $\tilde{\varepsilon}_t^{x} = \varepsilon_t^{x} + \tilde{\varepsilon}$, $\tilde{\varepsilon}_t^{\beta} = \varepsilon_t^{\beta} - \frac{1}{b_{1,t-1}}\tilde{\varepsilon}$, and all observable variables will be equivalent. Our identification assumptions are $\rho_{\lambda,rp} = 0$ and $\rho_{\beta,rp} = 0$.

B.2 Variance-Covariance Matrix

We can also decompose the variance in the level of the real exchange rate. Starting with

$$\begin{aligned}x_t^{\lambda} &= \omega\gamma_1 x_{t-1}^{\lambda} + \frac{1}{1-\gamma_1}\left(\gamma_1\tilde{\lambda}_{t-1} + \varepsilon_t^{\lambda}\right), \\ y_t^{i} &= \omega\psi_1 y_{t-1}^{i} + \frac{1}{1-\psi_1}\left(\psi_1\tilde{i}_{t-1} + \varepsilon_t^{i}\right), \\ z_t^{rp} &= \omega\delta_1 z_{t-1}^{rp} + \frac{1}{1-\delta_1}\left(\delta_1\tilde{rp}_{t-1} + \varepsilon_t^{rp}\right), \\ \tilde{\lambda}_t &= \gamma_1\tilde{\lambda}_{t-1} + \varepsilon_t^{\lambda}, \\ \tilde{i}_t &= \psi_1\tilde{i}_{t-1} + \varepsilon_t^{i}, \\ \tilde{rp}_t &= \delta_1\tilde{rp}_{t-1} + \varepsilon_t^{rp},\end{aligned}$$

we can express this system as a VAR of $\tilde{X}_t = [x_t^{\lambda}, y_t^{i}, z_t^{rp}, \tilde{\lambda}_t, \tilde{i}_t, \tilde{rp}_t]'$:

$$\tilde{X}_t = \tilde{\Psi}\tilde{X}_{t-1} + \tilde{\Gamma}\varepsilon.$$

Then the covariance of $\tilde{X}_t$ can be solved from

$$\tilde{X}_t\tilde{X}_t' = \tilde{\Psi}\tilde{X}_t\tilde{X}_t'\tilde{\Psi}' + \tilde{\Gamma}cov(\varepsilon)\tilde{\Gamma}'$$

Separate Online Appendix

A Data Sources

We start by discussing the Panel Dataset. For the FX data, before December 1996, we use the Barclays Bank source from Datastream. After December 1996, we use World Markets Reuters (WMR) from Datastream. The Datastream codes for the spot rates and 12M forward rates are: *BBGBPSP, BBGBPYF, BBAUDSP, BBAUDYF, BBCADSP, BBCADYF, BBDEMSP, BBDEMYF, BBJPYSP, BBJPYYF, BNZDSP, BBNZDYF, BBNOKSP, BBNOKYF, BBSEKSP, BBSEKYF, BBCHFSP, BBCHFYF, AUSTDOL, UKAUDYF, CNDOLLR, UKCADYF, DMARKER, UKDEMYF, JAPAYEN, UKJPYYF, NZDOLLR, UKNZDYF, NORKRON, UKNOKYF, SWEKRON, UKSEKYF, SWISSFR, UKCHFYF, UKDOLLR, UKUSDYF.*

For the Government Bond Yields (see Table A.2), most country-maturities pairs only use one source, except if there are gaps. If there are gaps, we use all the data from the first source wherever available, as indicated in the Table, and then fill in any gaps

Table A.1: Country Composition of Unbalanced Panel

Country	Maturity	Source	Ranges			
'Australia'	12	All	199912 - 201707			
'Canada'	12	All	199312 - 201707			
'Germany'	12	All	199707 - 201707			
'Japan'	12	All	199504 - 201707			
'New Zealand'	12	All	199603 - 200905	201006 - 201212	201310 - 201412	201606 - 201707
'Norway'	12	All	199001 - 199611	199701 - 201707		
'Sweden'	12	All	199103 - 199611	199701 - 201304	201306 - 201707	
'Switzerland'	12	All	198801 - 201707			
'United Kingdom'	12	All	199707 - 201707			
'United States'	12	All	198801 - 201707			

Table A.2: Sources for Government Bond Yields

Country	Maturity	Months		Mnemonic
'Australia'	12	'Bloomberg'		GTAUD1Y Govt
'Canada'	12	'Bank of Canada (Datastream)'		CNTBB1Y
'Germany'	12	'Bloomberg'		GTDEM1Y Govt
'Japan'	12	'Bloomberg'		GTJPY1Y Govt
'New Zealand'	12	'Bloomberg'	1	GTNZD1Y Govt
'New Zealand'	12	'Reserve Bank of New Zealand (Datastream)'	2	NZGBY1Y
'Norway'	12	'Oslo Bors'		ST3X
'Sweden'	12	'Sveriges Riksbank (from Cristina)'	2	
'Sweden'	12	'Sveriges Riksbank (website)'	1	
'Switzerland'	12	'Swiss National Bank'		
'United Kingdom'	12	'Bloomberg'		GTGBP1Y Govt
'United States'	12	'Bloomberg'	1	GB12 Govt
'United States'	12	'FRED'	2	

The numbers indicate which source takes precedence.

for some year month using the second data source (indicated by '2'). For LIBORs (see Table A.3), we use the BBA-ICE LIBOR when available. Coverage is good for Germany, Japan, Switzerland, UK, and U.S.. For other countries, we then use other interbank survey rates (BBSW, CDOR, NIBOR, STIBOR) to fill in any gaps. We then use deposit rates (Bank Bill, NKD, SKD) for any remaining gaps.

Table A.3: Sources for LIBOR

Country	Maturity	Source	Mnemonic
'Australia'	12	'Bank Bill (Bloomberg)'	ADBB12M Curncy
'Australia'	12	'Bank Bill Swap (Bloomberg)'	BBSW1Y Index/BBSW1MD Index
'Canada'	12	'CDOR (Bloomberg)'	CDOR12 Index
'Australia'	12	'BBA-ICE LIBOR (Datastream)'	BBAUD12
'New Zealand'	12	'Bank Bill (Bloomberg)'	NDBB12M Curncy
'Canada'	12	'BBA-ICE LIBOR (Datastream)'	BBCAD12
'Germany'	12	'BBA-ICE LIBOR (Datastream)'	BBDEM12
'Japan'	12	'BBA-ICE LIBOR (Datastream)'	BBJPY12
'New Zealand'	12	'BBA-ICE LIBOR (Datastream)'	BBNZD12
'Norway'	12	'NIBOR (Bloomberg)'	NIBOR12M Index
'Norway'	12	'Norwegian Krone Deposit (Bloomberg)'	NKDR1 Curncy
'Sweden'	12	'BBA-ICE LIBOR (Datastream)'	BBSEK12
'Sweden'	12	'STIBOR (Bloomberg)'	STIB1Y Index
'Sweden'	12	'Swedish Krona Deposit (Bloomberg)'	SKDR1 Curncy
'Switzerland'	12	'BBA-ICE LIBOR (Datastream)'	BBCHF12
'United Kingdom'	12	'BBA-ICE LIBOR (Datastream)'	BBGBP12
'United States'	12	'BBA-ICE LIBOR (Datastream)'	BBUSD12

B Theory of Convenience Yields and Exchange Rates

This section explores three issues: (i) what happens in a complete markets environment to exchange rates when investors derive convenience yields, (ii) what if foreign bonds also deliver convenience yields to their owners, and (iii) an analysis of the role of the banking sector in LIBOR markets.

B.1 Convenience Yields in Complete Markets

We follow the approach of Backus, Foresi and Telmer (2001). Consider the Euler equations (3) and (8) for the U.S. and foreign investor when investing in the foreign bond. To satisfy these Euler equations, we conjecture an exchange rate process that satisfies,

$$M_t^{\$} \frac{S_t}{S_{t+1}} = M_t^*.$$

This guess, as can easily be verified, satisfies the Euler equations. If financial markets are complete, then this is the unique exchange rate process that is consistent with the absence of arbitrage opportunities. Using lower case letters to denote logs, and log-linearizing this expression, we find:

$$\Delta s_{t+1} = m_t^{\$} - m_t^*. \tag{40}$$

Next consider the pair of Euler equations, (4) and (9), which apply to investments in the U.S. bond that gives a convenience yield. We conjecture an exchange rate process that satisfies,

$$M_t^* e^{\lambda_t^*} \frac{S_{t+1}}{S_t} = M_t^\$ e^{\lambda_t^\$}.$$

Log-linearizing this expression, we find:

$$\Delta s_{t+1} = \left(m_t^\$ - m_t^*\right) + \left(\lambda_t^\$ - \lambda_t^*\right) \tag{41}$$

It is evident that (40) and (41) cannot both be satisfied in an equilibrium unless $\lambda_t^* = \lambda_t^\$$. But note that in this case, convenience yields have no impact on exchange rates.

How is equilibrium restored when $\lambda_t^* \neq \lambda_t^\$$? The answer is that one of the Euler equations must be an inequality. There are many ways this may happen. Portfolio choices could be at a corner. For example, if foreign investors assign a positive convenience yield to their own foreign bonds, while U.S. investor do not, then the U.S. investor Euler equation does not apply to foreign bonds. Alternatively, if foreign convenience demand for U.S. bonds is so high that U.S. investors do not own U.S. bonds, then the U.S. Euler equation does not apply to U.S. bonds. Another possibility are forms of market segmentation. Suppose that some U.S. investors derive convenience value from U.S. bonds, but these same investors do not own foreign bonds. Other U.S. investors do not derive convenience value from U.S. bonds, and these investors do own foreign bonds. In these cases as well, one of the Euler equations we have posited is an inequality. Alternatively, we could consider scenarios in which the U.S. Euler equation for foreign bonds does not hold for all investors. Suppose that the Euler equations for the U.S. investor in foreign bonds apply to a financial intermediary that is subject to financing frictions as in intermediary asset pricing models. Then, the Lagrange multiplier on this constraint will enter the Euler equation, so that a binding constraint can also restore equilibrium. But note that even with such frictions, our equations linking foreign convenience yield valuations and the exchange rate remain valid.

B.2 Convenience Yields on Foreign Bonds

This section allows for a convenience yield on foreign bonds. We use $\lambda_t^{i,j}$ to denote the convenience yield of investors in country j for bonds issued by the government in country i. Similarly, $\beta^{i,j}\lambda_t^{i,j}$ is the convenience yield of investors in country j for LIBOR deposits in i's currency.

Foreign investors price foreign bonds denominated in foreign currency, and the foreign investor's Euler equation is given by:

$$\mathbb{E}_t\left(M_{t+1}^* e^{y_t^*}\right) = e^{-\lambda^{*,*}}. \tag{42}$$

Foreign investors can also invest in U.S. Treasurys. To do so, they convert local currency to U.S. dollars to receive $\frac{1}{S_t}$ dollars, invest in U.S. Treasurys, and then convert the proceeds back to local currency at date $t+1$ at S_{t+1}. Then,

$$\mathbb{E}_t\left(M_{t+1}^* \frac{S_{t+1}}{S_t} e^{y_t^\$}\right) = e^{-\lambda_t^{\$,*}}, \quad \lambda_t^* \geq 0. \tag{43}$$

Next, we use these pricing conditions to derive an expression linking the exchange rate and the convenience yield. We assume that $m_t^* = \log\ M_t^*$ and $\Delta s_{t+1} = \log \frac{S_{t+1}}{S_t}$ are conditionally normal. Then, (42) can be rewritten as,

$$\mathbb{E}_t\left(m_{t+1}^*\right) + \frac{1}{2}Var_t\left(m_{t+1}^*\right) + y_t^* + \lambda_t^{*,*} = 0, \tag{44}$$

and (43) as,

$$\mathbb{E}_t\left(m_{t+1}^*\right) + \frac{1}{2}Var_t\left(m_{t+1}^*\right) + \mathbb{E}_t[\Delta s_{t+1}] + \frac{1}{2}var_t[\Delta s_{t+1}] + y_t^{\$} + \lambda_t^{\$,*} - RP_t^* = 0. \tag{45}$$

Here $RP_t^* = -cov_t\left(m_{t+1}^*, \Delta s_{t+1}\right)$ is the risk premium the foreign investor requires for the exchange rate risk when investing in U.S. bonds. We combine these two expressions to find that the expected return in levels on a long position in dollars earned by a foreign investor is given by:

$$\mathbb{E}_t[\Delta s_{t+1}] + \left(y_t^{\$} - y_t^*\right) + \frac{1}{2}var_t[\Delta s_{t+1}] = RP_t^* - \lambda_t^{\$,*} + \lambda_t^{*,*}. \tag{46}$$

The U.S. investor's Euler equation when investing in the foreign bond is:

$$\mathbb{E}_t\left(M_{t+1}^{\$}\frac{S_t}{S_{t+1}}e^{y_t^*}\right) = e^{-\lambda_t^{*,\$}}. \tag{47}$$

We also assume that U.S. investors derive a convenience yield when investing in U.S. Treasurys:

$$\mathbb{E}_t\left(M_{t+1}^{\$}e^{y_t^{\$}}\right) = e^{-\lambda_t^{\$,\$}}. \quad \lambda_t^{\$,\$} \geq 0. \tag{48}$$

$\lambda_t^{\$}$ is asset-specific. An increase in the U.S. investor's convenience yield lowers U.S. Treasury bond yields, holding the SDF fixed: $y_t^{\$} = \rho_t^{\$} - \lambda_t^{\$,\$}$, where $\rho_t^{\$} = -\log \mathbb{E}_t\left(M_{t+1}^{\$}\right)$.

We assume log-normality and rewrite these equations to derive an expression for the carry trade return,

$$\left(y_t^* - y_t^{\$}\right) - \mathbb{E}_t[\Delta s_{t+1}] + \frac{1}{2}var_t[\Delta s_{t+1}] = RP_t^{\$} - \lambda_t^{*,\$} + \lambda_t^{\$,\$}. \tag{49}$$

where, $RP_t^{\$} = -cov_t\left(m_{t+1}^{\$}, -\Delta s_{t+1}\right)$ is the risk premium the U.S. investor requires for the exchange rate risk when investing in foreign bonds (i.e. the risk premium attached to the dollar appreciating).

Finally, we combine (46) and (49) to derive a cross-country restriction on the convenience yields imputed to Treasurys and the currency risk premia,

$$(\lambda_t^{\$,*} - \lambda_t^{*,*}) - (\lambda_t^{\$,\$} - \lambda_t^{*,\$}) = RP_t^{\$} + RP_t^* - var_t[\Delta s_{t+1}]. \tag{50}$$

By forward iteration on eqn. (46), the level of exchange rates can be stated as a function of the interest rate differences, the currency risk premia and the future convenience yields (see Froot and Ramadorai, 2005, for a version without convenience yields):

$$s_t = \mathbb{E}_t\sum_{\tau=0}^{\infty}(\lambda_{t+\tau}^{\$,*} - \lambda_{t+\tau}^{*,*}) + \mathbb{E}_t\sum_{\tau=0}^{\infty}(y_{t+\tau}^{\$} - y_{t+\tau}^*) - \mathbb{E}_t\sum_{\tau=0}^{\infty}\left(RP_{t+j}^* - \frac{1}{2}Var_{t+j}[\Delta s_{t+j+1}]\right) + \bar{s}. \tag{51}$$

The term $\bar{s} = \mathbb{E}_t[\lim_{j\to\infty}\ s_{t+j}]$ which is constant under the assumption that the nominal exchange rate is stationary.

Last, we construct the basis measure:

$$\begin{aligned} x_t^{Treas} &\equiv y_t^{\$} + (f_t^1 - s_t) - y_t^* \\ &= (y_t^{\$} - y_t^{\$,Libor}) - (y_t^* - y_t^{*,Libor}) \\ &= -(1 - \beta^{\$,*})\lambda_t^{\$,*} + (1 - \beta^{*,*})\lambda_t^{*,*} \end{aligned}$$

The basis reflects the difference between the relative yields of dollar government bonds and LIBOR deposits, and foreign government bonds and foreign deposits.

Proposition 4. *Under the assumption that $\beta^{\$,*} = \beta^{*,*}$, the level of the nominal exchange can be written as:*

$$s_t = -\mathbb{E}_t \sum_{\tau=0}^{\infty} \frac{x_t^{Treas}}{1 - \beta_t^{\$,*}} + \mathbb{E}_t \sum_{\tau=0}^{\infty} (y_{t+\tau}^{\$} - y_{t+\tau}^*) - \mathbb{E}_t \sum_{\tau=0}^{\infty} \left(RP_{t+\tau}^* - \frac{1}{2} Var_{t+\tau}[\Delta s_{t+\tau+1}] \right) + \bar{s}. \tag{52}$$

The assumption that the beta's for U.S. and foreign LIBOR are the same means that the relative safety of these assets are the same regardless of currency. To a first order, the assumption seems plausible to us.

B.3 Convenience yields on LIBOR deposits, the LIBOR basis, and the Treasury basis

In U.S. data, Krishnamurthy and Vissing-Jorgensen (2012) observe that there is a convenience yield on both Treasury bonds and other near-riskless private bonds such as bank deposits. They moreover show that some investors view near-riskless private bonds as partial substitutes for Treasury bonds. This section introduces dollar LIBOR deposits which also offer convenience yields to foreign investors, but less so than U.S. Treasurys. That is, as noted earlier, our theory posits that investors receive convenience utility from U.S. safe assets, a set that includes both U.S. Treasurys and bank deposits. We first show how to understand the LIBOR basis and safe asset demand in this case, and then offer another way to understand the Treasury basis.

Foreign and domestic investors have access to U.S. LIBOR markets, and they satisfy the following Euler equations:

$$\mathbb{E}_t \left(M_{t+1}^{\$} e^{y_t^{\$,Libor}} \right) = e^{-\beta_t^{\$} \lambda_t^{\$}}$$

$$\mathbb{E}_t \left(M_{t+1}^* \frac{S_{t+1}}{S_t} e^{y_t^{\$,Libor}} \right) = e^{-\beta_t^* \lambda_t^*}$$

where $\beta_t^{\$} \lambda_t^{\$}$ ($\beta_t^* \lambda_t^*$) denotes the convenience yield from a cash position in dollars derived by U.S. (foreign) investors, and these βs are less than one.

Banks issue foreign and dollar deposits that pay LIBOR at rates $y_t^{\$,Libor}$ and $y_t^{*,Libor}$. The dollar deposits offer a convenience yield to investors but not to the banks, so that banks will wish to issue these deposits in equilibrium. Consider a given bank that has a mix of deposits in both currencies in (dollar-equivalent) amounts $(\bar{\theta}_t^{B,\$}, \bar{\theta}_t^{B,*})$. We suppose the mix is optimal for the bank given asset/liability management concerns and the currency mix of the rest of its balance sheet. If bank deposits offer convenience yields, than banks will create these deposits, and the limit on such deposit creation will be governed by bank costs in creating the deposits. See the model of Krishnamurthy and Vissing-Jorgensen (2015) for one specification of intermediaries doing asset/liability management and creating money where the cost is in terms of collateral backing. We have suppressed the specification of these costs to not stray from our primary analysis which is exchange rate determination. Think of the optimal mix $(\bar{\theta}_t^{B,\$}, \bar{\theta}_t^{B,*})$ as being driven by these costs.

Next, suppose that the bank also trades in the forward market. Clearly if the convenience yield on dollar deposits rises relative to foreign deposits, the bank will want to supply more of these dollar deposits and hedge these using the forward market to maintain its optimal currency mix. Then the bank chooses θ_t^B, the quantity of this swap, to achieve deposit mix $(\bar{\theta}_t^{B,\$} + \theta_t^B, \bar{\theta}_t^{B,*} - \theta_t^B)$. If there is greater demand for dollar deposits the bank will on the margin increase θ_t^B. Suppose the bank solves:

$$\max_{\theta_t^B} \theta_t^B \left(y_t^{*,Libor} - (f_t - s_t) - y_t^{\$,Libor} \right) - \frac{\kappa}{2} \left(\theta_t^B \right)^2 .$$

Here κ is a capital/leverage cost associated with doing the forward and hedging the dollar deposits. The term $y_t^{*,Libor} - (f_t - s_t) - y_t^{\$,Libor}$ is the funding cost reduction that the bank gets when taking advantage of the dollar convenience yield. The F.O.C. for the bank is,

$$\begin{aligned} -\kappa\theta_t^B &= y_t^{\$,Libor} - y_t^{*,Libor} + (f_t - s_t) \\ &= x_t^{LIBOR} \end{aligned}$$

where x_t^{LIBOR} denotes the LIBOR basis. If $y_t^{\$,Libor}$ is particularly low, e.g., driven by an increase in demand for dollar deposits, then x_t^{LIBOR} will rise and banks will increase the supply of dollar deposits, θ_t^B, while swapping these dollars deposits back into foreign currency to keep their exchange rate exposure unaffected. Suppose there are many banks and denote the aggregate quantity of dollar deposits supplied in equilibrium as Θ_t^B. Then, the equilibrium LIBOR basis is given by:

$$x_t^{LIBOR} = -\kappa\Theta_t^B. \tag{53}$$

Lemma 5. *The LIBOR basis depends on foreign demand for dollar deposits as follows: When banks face no capital/leverage costs in doing swaps and $\kappa = 0$, the LIBOR basis is zero and independent of Θ_t^B. When $\kappa > 0$, the LIBOR basis becomes more negative as the demand for dollar safe assets rises.*

In the frictionless case, as κ goes to zero, banks actively trade in the forward to earn the convenience yield on dollar deposits while not altering their exchange rate exposure. In equilibrium, the price of the forward will adjust to equalize these margins and the LIBOR C.I.P. deviation goes to zero. Perhaps surprisingly. the forward price, f_t^1, can embed a convenience yield. We come to this later when interpreting carry trade relations.

In an influential recent paper, Du, Tepper and Verdelhan (2017) document that the LIBOR basis was near zero pre-crisis and has often been significantly different than zero post-crisis. They show that the movements in the LIBOR basis are closely connected to frictions in financial intermediation that prevent arbitrage activities. Other papers have come to similar conclusions regarding the importance of financial frictions and capital controls (see Ivashina, Scharfstein and Stein, 2015; Gabaix and Maggiori, 2015;

Amador et al., 2017; Itskhoki and Mukhin, 2017). Our lemma shows, consistent with the findings of Du, Tepper and Verdelhan (2017), that when $\kappa > 0$, LIBOR C.I.P. will fail. More novel, our theory implies that when $\kappa > 0$, x_t^{LIBOR} will, like λ_t^*, reflect foreign investors's demand for safe dollar assets. We will verify this prediction in the data post-crisis.

We next reconsider the Treasury basis in light of the LIBOR basis:

$$x_t^{Treas} = y_t^{\$} - y_t^* + f_t - s_t = (y_t^{\$} - y_t^{\$,Libor}) - (y_t^* - y_t^{*,Libor}) + x_t^{LIBOR}. \tag{54}$$

The Treasury basis is the sum of the LIBOR basis and the difference between the two currency's Treasury-LIBOR spreads. We can further rewrite this expression as,

$$x_t^{Treas} = -(1-\beta^*)\lambda_t^* + x_t^{LIBOR}.$$

where we have used the relation that the Treasury-LIBOR spread is zero in the foreign country and equal to $(1-\beta^*)\lambda_t^*$ in dollars. We note that the Treasury basis measures the foreign demand for U.S. safe assets through both the Treasury convenience yield λ_t^* and through movements in the quantity of dollar deposits (Θ_t^B). When banks face no capital/leverage costs in doing swaps, the LIBOR basis is zero. Banks actively trade in the forward to earn the convenience yield on dollar deposits while not altering their exchange rate exposure. In equilibrium, the price of the forward will adjust to equalize these margins and the LIBOR C.I.P. deviation goes to zero.

In this case, we note another way of understanding why the Treasury basis measures the foreign convenience yield:

$$\lambda_t^* = -\frac{x_t^{Treas}}{1-\beta^*}. \tag{55}$$

The key behind this result is that both Treasury bond yields and LIBOR rates reflect the foreign convenience yield, but differentially. Thus the difference, as reflected in the basis, directly measures the foreign convenience yield. In this version of the model, where only U.S. bonds have convenience yields, the term $\left(y_t^{\$} - y_t^{\$,Libor}\right)$ is what captures the foreign convenience yield. When LIBOR C.I.P. holds, the convenience yield from a currency-hedged position in Treasurys equals the U.S. LIBOR-Treasury spread. From the standpoint of the U.S. investor, $x_t^{Treas} = -(1-\beta^{\$})\lambda_t^{\$}$. The Treasury basis is also related to U.S. investor's convenience valuation of Treasury bonds and LIBOR deposits. As noted earlier, for us to find a relation between convenience yields and exchange rates, we must have that $\lambda_t^{\$} \neq \lambda_t^*$, which further implies that $\beta^{\$} \neq \beta^*$. When both foreign and U.S. bonds carry convenience yields, the difference in the LIBOR-Treasury spreads across both countries is the right measure of the convenience yield as we show in the appendix.

C Robustness

Our robustness tests include (1) replacing the innovation to Treasury basis by the change in Treasury basis, (2) re-running regressions in subsamples with high and low Treasury basis, (3) replacing the 12-month Treasury basis by the 3-month Treasury basis, (4) changing the dependent variable from the dollar's excess return to its exchange rate movement, (5) using KfW bonds in dollar and foreign currencies instead of using U.S. and foreign government bonds, and (6) re-running the regression using quarterly averages.

C.1 Explaining Variation in the Dollar Using Change in Treasury Basis

In Table 3, we use the innovation in Treasury basis as the explanatory variable. Here we use the change $\Delta \bar{x}_t^{Treas} = \bar{x}_t^{Treas} - \bar{x}_{t-1}^{Treas}$ instead.

Table A.4: Average Treasury Basis and the USD Spot Nominal Exchange Rate

The dependent variable is the quarterly change in the log of the spot USD exchange rate against a basket. The independent variables are the change in the average Treasury basis, $\Delta \bar{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the change, the change in the LIBOR basis, and the change in the U.S.-to-foreign Treasury yield differential. Data is quarterly. OLS standard errors in parentheses.

	1988Q1−2017Q2					1988Q1−2007Q4		2008Q1−2017Q2	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta \bar{x}^{Treas}$	-3.93		-5.93		-5.79	-2.19		-9.27	
	(1.87)		(1.87)		(1.63)	(2.31)		(3.03)	
$\Delta \bar{x}^{LIBOR}$		3.32					8.09		-5.97
		(2.51)					(3.17)		(3.67)
Lag $\Delta \bar{x}^{Treas}$			-6.01		-4.65				
			(1.87)		(1.65)				
$\Delta(y^{\$} - \bar{y}^*)$				3.84	3.71				
				(0.66)	(0.62)				
R^2	3.7	1.5	12.0	22.6	33.4	1.1	7.7	21.1	7.0
N	117	117	116	117	116	80	80	37	37

C.2 Explaining Variation in the Dollar when Treasury Basis is Small

This section repeats the regressions in Table 3 in the subsample in which the average U.S. Treasury basis is above or equal to the 25th percentile (-68 basis points), which represent the 75% of the data in which the Treasury basis is not very negative. We find that the U.S. dollar's exchange rate also comoves with the U.S. Treasury basis in these calm periods.

Table A.5: Average Treasury Basis and the USD Spot Nominal Exchange Rate, Calm Periods

The dependent variable is the quarterly change in the log of the spot USD exchange rate against a basket. The independent variables are the innovation in the average Treasury basis, $\Delta\bar{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation, the innovation in the LIBOR basis, and the innovation in the U.S.-to-foreign Treasury yield differential. Data is quarterly. OLS standard errors in parentheses.

	1988Q1−2017Q2					1988Q1−2007Q4		2008Q1−2017Q2	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta\bar{x}^{Treas}$	-16.77		-17.43		-10.89	-14.38		-24.82	
	(3.06)		(3.22)		(3.36)	(3.56)		(6.21)	
$\Delta\bar{x}^{LIBOR}$		-11.11					-14.48		-7.73
		(4.06)					(6.37)		(5.82)
Lag $\Delta\bar{x}^{Treas}$			-2.06		-0.52				
			(2.93)		(2.72)				
$\Delta(y^{\$} - \bar{y}^{*})$				4.90	3.56				
				(0.80)	(0.87)				
Constant	0.002	-0.01	0.004	-0.01	-0.001	-0.002	-0.01	0.01	-0.004
	(0.004)	(0.004)	(0.005)	(0.004)	(0.004)	(0.01)	(0.01)	(0.01)	(0.01)
R^2	25.8	8.0	26.3	30.3	38.6	23.5	8.9	34.0	5.4
N	88	88	88	88	88	55	55	33	33

C.3 Explaining Variation in the Dollar Using 3M Treasury Basis

This section reports the regression results obtained when we use the 3-month Treasury basis instead of the 12-month Treasury basis, and use the 3-month interest rates instead of the 12-month interest rates. Because the time series of the 3-month Treasury basis is volatile, we also use the quarterly average instead of the last observation in each quarter. The results are much weaker.

Table A.6: Average 3M-Treasury Basis and the USD Spot Nominal Exchange Rate

The dependent variable is the quarterly change in the log of the spot USD exchange rate against a basket. The independent variables are the innovation in the average Treasury basis, $\Delta\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation, the innovation in the LIBOR basis, and the innovation in the U.S.-to-foreign Treasury yield differential. Data is quarterly. OLS standard errors in parentheses.

	1988Q1−2017Q2					1988Q1−2007Q4		2008Q1−2017Q2	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Last observation in 3 months									
$\Delta\overline{x}^{Treas}$	0.67		0.47		0.02	4.85		-1.79	
	(1.09)		(1.10)		(1.09)	(1.72)		(1.37)	
$\Delta\overline{x}^{LIBOR}$		-1.73					1.31		-3.21
		(1.72)					(2.76)		(2.36)
Lag $\Delta\overline{x}^{Treas}$			-1.24		-1.22				
			(1.10)		(1.08)				
$\Delta(y^{\$} - \overline{y}^{*})$				1.53	1.51				
				(0.59)	(0.60)				
Constant	0.0002	0.0002	0.0002	0.0002	0.0003	-0.002	-0.003	0.01	0.003
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.01)	(0.01)
R^2	0.3	0.8	1.3	5.2	6.3	8.5	0.3	4.6	5.0
N	125	125	124	125	124	88	88	37	37
Average across 3 months									
$\Delta\overline{x}^{Treas}$	-1.61		-1.76		-1.98	1.91		-8.92	
	(1.46)		(1.48)		(1.46)	(1.67)		(2.45)	
$\Delta\overline{x}^{LIBOR}$		-9.37					0.16		-16.44
		(3.57)					(5.26)		(5.27)
Lag $\Delta\overline{x}^{Treas}$			-2.58		-2.26				
			(1.46)		(1.44)				
$\Delta(y^{\$} - \overline{y}^{*})$				1.47	1.77				
				(0.70)	(0.72)				
Constant	-0.001	-0.001	-0.0004	-0.001	-0.0002	-0.003	-0.003	0.01	-0.0001
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.01)	(0.01)
R^2	1.0	5.3	3.2	3.5	7.8	1.5	0.0	27.4	21.7
N	125	125	124	125	124	88	88	37	37

C.4 Forecasting Returns and Exchange Rates Using 3-month Treasury Basis

This section reports predictability results obtained using the 3-month Treasury basis instead of the 12-month Treasury basis. Because the time series of the 3-month Treasury basis is volatile, we use the quarterly average instead of the last observation in each quarter.

Table A.7: Predicting Currency Excess Returns in the Panel using 3M Basis

The dependent variable is either $rx^{fx}_{t\to t+k}$ or $(4/k)\Delta s_{t\to t+k}$. The independent variables are the average Treasury basis, $\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation in the average Treasury basis, and the average nominal Libor rate difference ($y^{\$,Libor}_t - \overline{y}^{*,Libor}_t$) in units of log yield. Data is quarterly from 1988Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

Panel A: Dependent Variable is Excess Return				
Last observation in 3 months				
	3-month	1-year	2-year	3-year
$\overline{x}^{Treas}$	-4.61	3.47	4.31	5.95
	(4.14)	(5.44)	(3.59)	(2.87)
$y^{\$,Libor}_t - \overline{y}^{*,Libor}_t$	0.91	1.06	0.80	0.53
	(1.06)	(0.86)	(0.55)	(0.39)
$\Delta\overline{x}^{Treas}$		-3.80	-3.56	-4.69
		(3.88)	(2.21)	(1.84)
Constant	-0.02	0.01	0.01	0.02
	(0.02)	(0.03)	(0.03)	(0.02)
R^2	2.78	8.51	11.03	13.07
N	125	121	117	113
Average across 3 months				
	3-month	1-year	2-year	3-year
$\overline{x}^{Treas}$	-2.66	2.46	4.62	5.77
	(5.49)	(5.02)	(3.76)	(2.72)
$y^{\$,Libor}_t - \overline{y}^{*,Libor}_t$	1.35	1.27	1.00	0.71
	(0.81)	(0.78)	(0.53)	(0.39)
$\Delta\overline{x}^{Treas}$		-3.34	-5.44	-5.00
		(4.06)	(2.52)	(1.60)
Constant	-0.01	0.01	0.01	0.01
	(0.02)	(0.03)	(0.02)	(0.02)
R^2	4.49	11.17	16.54	18.69
N	125	121	117	113

C.5 Forecasting Exchange Rate Changes using 12M Treasury Basis

This section checks the predictability of nominal exchange rate movements instead of the predictability of currency excess returns.

Table A.8: Forecasting Exchange Rate Changes in Panel

The dependent variable is the annualized change in the exchange rate (in logs) $(4/k)\Delta s_{t\to t+k}$ on a long position in the dollar over k quarters. The independent variables are the average Treasury basis, $\overline{x}^{Treas}$, the nominal Treasury yield difference $(y^{\$} - \overline{y}^{*})$, the change in the average Treasury basis $\Delta\overline{x}_t^{Treas}$, and the lagged change in the average Treasury basis $\Delta\overline{x}_{t-1}^{Treas}$. Data is quarterly from 1988Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	$\overline{x}^{Treas}$	$y^{\$} - \overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2
			Panel A: 1988-2017		
3 months	-16.76	-1.17	-7.96	2.80	0.08
	(13.15)	(1.36)	(10.29)	(10.25)	
1 year	-0.65	-0.57	-7.02	-2.60	0.05
	(8.46)	(0.92)	(5.51)	(3.89)	
2 years	4.13	-0.28	-6.64	-4.48	0.06
	(4.81)	(0.56)	(3.25)	(2.41)	
3 years	6.94	0.05	-7.09	-4.21	0.10
	(3.64)	(0.34)	(2.54)	(1.91)	
			Panel B: 1988-2007		
3 months	-30.66	-1.37	5.79	14.55	0.13
	(15.04)	(1.47)	(9.51)	(10.52)	
1 year	-10.39	-0.71	-5.02	-3.32	0.14
	(9.22)	(0.88)	(6.83)	(5.13)	
2 years	-3.02	-0.47	-3.22	-3.44	0.07
	(5.09)	(0.56)	(4.00)	(2.92)	
3 years	3.77	-0.03	-6.39	-4.23	0.07
	(4.44)	(0.32)	(3.30)	(2.35)	
			Panel C: 2007-2017		
3 months	17.94	-4.58	-32.05	-29.41	0.22
	(10.92)	(3.04)	(11.78)	(10.40)	
1 year	18.94	-3.06	-6.31	-3.31	0.25
	(6.90)	(2.50)	(5.97)	(3.08)	
2 years	20.55	-0.07	-12.95	-8.59	0.50
	(3.13)	(0.74)	(1.91)	(1.96)	
3 years	13.35	-0.19	-7.26	-5.26	0.35
	(2.80)	(0.70)	(1.79)	(1.39)	

Table A.9: Forecasting 3-year Exchange Rate Changes in Panel

In Panel A (B), the dependent variable is the annualized nominal excess return (in logs) $(4/12)\Delta s_{t\rightarrow t+12}$ ($(4/8)\Delta s^{fx}_{t+4\rightarrow t+12}$) on a long position in U.S. Treasuries and a short position (equal-weighted) in all foreign bonds over k quarters. The independent variables are the Treasury basis, x^{Treas}, the nominal Treasury yield difference ($y^{\$} - \overline{y}^{*}$), the change in the Treasury basis $\Delta\overline{x}^{Treas}_{t}$, and the lagged change in the Treasury basis $\Delta\overline{x}^{Treas}_{t-1}$. Data is quarterly from 1988Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	$\overline{x}^{Treas}$	$y^{\$}-\overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2	$\overline{x}^{Treas}$	$y^{\$}-\overline{y}^{*}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	R^2
	Panel A: Change from year t to $t+3$					*Panel B: Change from year $t+1$ to $t+3$*				
1988-2017	2.17				0.01	6.31				0.07
	(2.33)					(2.74)				
	1.98	-0.09			0.01	6.52	0.10			0.07
	(2.24)	(0.29)				(2.59)	(0.34)			
	4.64	-0.01	-4.58		0.06	8.72	0.25	-4.59		0.09
	(2.96)	(0.31)	(1.70)			(3.48)	(0.34)	(2.36)		
	6.94	0.05	-7.09	-4.21	0.10	11.24	0.38	-7.58	-5.37	0.12
	(3.64)	(0.34)	(2.54)	(1.91)		(4.40)	(0.34)	(3.66)	(2.85)	
1988-2007	-0.36				0.00	6.44				0.07
	(2.84)					(3.13)				
	-0.75	-0.14			0.00	6.68	0.09			0.07
	(2.64)	(0.28)				(3.06)	(0.34)			
	1.51	-0.09	-3.55		0.03	8.73	0.22	-4.12		0.08
	(3.65)	(0.29)	(2.07)			(4.23)	(0.36)	(2.95)		
	3.77	-0.03	-6.39	-4.23	0.07	10.86	0.32	-7.07	-4.68	0.10
	(4.44)	(0.32)	(3.30)	(2.35)		(5.42)	(0.37)	(4.89)	(3.67)	
2007-2017	8.02				0.21	5.28				0.05
	(2.71)					(2.98)				
	8.03	-0.56			0.23	5.27	0.82			0.07
	(2.63)	(0.59)				(3.14)	(1.34)			
	10.41	-0.38	-5.36		0.29	7.87	1.02	-5.89		0.12
	(2.60)	(0.66)	(1.95)			(3.15)	(1.37)	(2.50)		
	13.35	-0.19	-7.26	-5.26	0.35	11.88	1.29	-8.47	-7.17	0.18
	(2.80)	(0.70)	(1.79)	(1.39)		(3.47)	(1.37)	(2.31)	(2.13)	

C.6 KFW Bonds

Figure A.1 plots the basis for KfW bonds. KfW is a German issuer whose bonds are backed by the German government, so that they are near default free. KfW issues bonds in different currencies allowing us to compute the basis for the bonds of the same issuer, i.e., holding safety fixed, in different currencies. We compute the basis for KfW bonds using one-year yields on these bonds for Australia, Euro, UK, and Switzerland against the U.S.. The yield data is from Bloomberg and corresponds to a fitted yield at the one-year maturity (one-year maturity bonds do not always exist). Clearly this measure is not as reliable as our Treasury basis measure which only uses information from traded instruments. Figure A.1 plots the cross-country mean KfW basis and the Treasury basis (cross-country mean for the same countries) over a sample with daily data from 2011Q2 to 2017Q2.

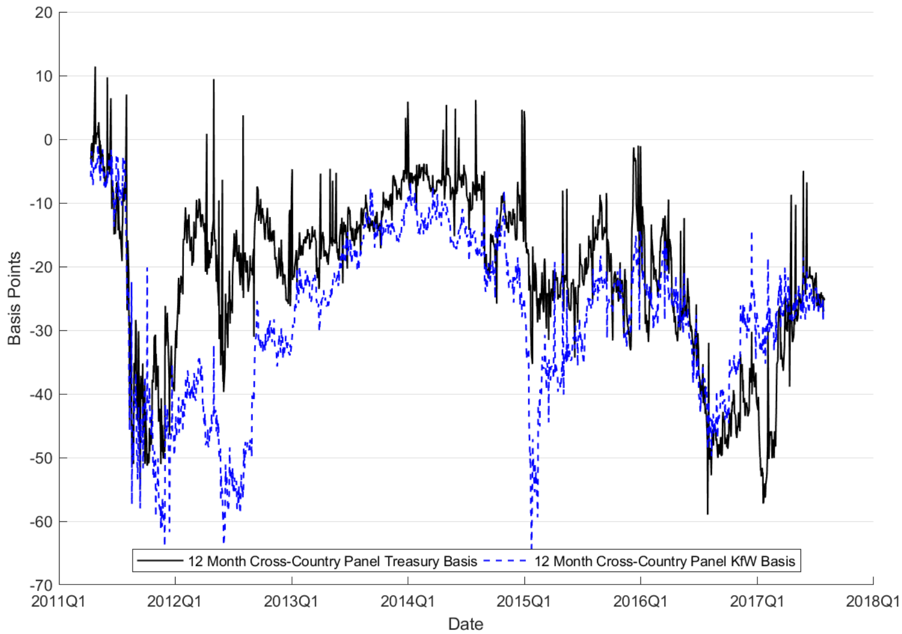

Figure A.1: KfW and Treasury Basis, 2011Q2 to 2017Q2

C.7 Market Microstructure

The FX markets in both spot and forward are large and liquid. Nevertheless, one may want to know the extent to which the relation we uncover stems from micro-structure order flow effects as in Evans and Lyons (2002) or Froot and Ramadorai (2005). Our theory does not involve these types of effects, and to test our theory ideally our data would reflect the mid of the bid and ask. By computing a quarterly average, we average out bid-ask bounce and thus likely measure true mid-market prices. The relation we uncover is quite strong in this averaged data (in fact it is stronger than the end-of-quarter data of Table 3).[5] The variation reflected in the exchange rate is an order of magnitude larger than typical bid-ask spreads. The standard-deviation of exchange rate changes in log points is 0.04, or 4%, which is well above typical bid-ask spreads. The standard-deviation of Treasury basis changes is 0.00134 (13.4 basis points). The slope coefficient on the fitted regression line of -14.5 implies that a one standard deviation change in the basis drives a 1.94% move in the exchange rate, which is also an order of magnitude larger than bid-ask spreads. Finally, the evidence in column (3) of Table 3 for momentum relates the lagged innovation in the basis to next quarter's change in the exchange rate.[6]

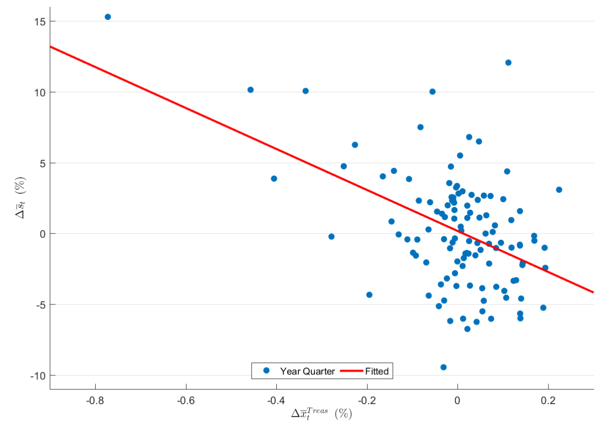

Figure A.2: Scatter plot of changes in the log exchange rate, averaged over a quarter, against shocks to the quarterly average basis. Data is from 1988Q1 to 2017Q2. In red we plot the fitted regression line. The R^2 is 22.8% and the slope coefficient is -14.6 with a t-statistic of 5.8.

[5]Figure A.2 presents a scatter plot of the change in the quarterly average log exchange rate against the change in the quarterly average basis.

[6]In section C.5 of the separate Appendix, we also show predictability evidence in Table A.8 and A.17 relating the current basis to future changes in the exchange rate. Our results are evidently not driven by micro-structure effects.

D The U.S. vs U.K. Treasury Basis and the USD/GBP Exchange Rate

Our second dataset covers the U.S./UK cross. This data begins much earlier, in 1970Q1 and ends in 2016Q2. The daily data quality is poor, with many missing values and implausible spikes in the constructed basis from one day to the next. To overcome these measurement issues, we take the average of the available data for a given quarter as the observation for that quarter. We construct the Treasury basis in the same manner as described earlier. We rely on Global Financial Data as the main data source.

Table A.10: Sources for U.S.-UK Time Series

	Source	Mnemomic	Range
Spot FX	GFD	*GBPUSD*	1960 – 2017
3M Forward	GFD	*GBPUSD*3*D*	1960 – 2017
12M Forward	GFD	*GBPUSD*12*D*	1960 – 2017
3M T-bill UK	GFD	*ITGBR*3*D*	1960 – 2017
1Y Note UK	GFD	*IGGBR*1*D*	1979 – 2017
1Y Note U.S.	FRED	*DTB*1*YR*	1960 – 2017
1Y Zero-Coupon	BoE		1970 – 1979

(GFD is Global Financial Data. FRED is the Federal Reserve Economic Database at the Federal Reserve Bank of St Louis. BoE is the Bank of England.)

Figure A.3 plots the resulting series. LIBOR rates do not exist back to 1971. The average U.S./UK Treasury basis is 0.84 bps per annum. On average, U.K. investors are close to indifferent between holding U.S. Treasurys on a currency-hedged basis and holding gilts. However, the standard deviation is 48 bps. per quarter. For comparison the figure also plots the mean basis from the cross-country panel. The two series track each other closely for the period where they overlap, but the U.S./UK basis is consistently higher than the panel basis. This result suggests that UK bonds also have a convenience yield, which is sometimes larger than that of U.S. bonds particularly in the 1970s, during which the basis is volatile and frequently positive. Suffering a balance-of-payments deficit in the early 1970s, the Nixon administration decided to suspend convertibility of the dollar into gold in 1973 and effectively ended the Bretton-Woods system. This action led to considerable uncertainty in the international monetary system, with some observers noting that foreigners became unwilling to continue to hold the dollar assets necessary to finance the balance-of-payments deficit (see Bach et al. (1972) and Farhi and Maggiori (2017)). Additionally, the U.K. suffered a balance-of-payments crisis in 1976, turning to the IMF for a large loan. These reductions in asset demand, first for U.S. and then for U.K. bonds, are apparent in the figure: the basis turns positive in 1973 before subsequently turning negative in 1976.

Figure A.4 plots the real exchange rate in units of GBP-per-USD (dashed line) against the U.S./UK Treasury basis (full line). Both series are based on quarterly averaged data. We use the real exchange rate here because there are clear trends in the price levels of both countries in the 1970s and early 1980s that we would expect to enter exchange rate determination. It is evident that the two series are negatively correlated. Table A.11 presents regressions analogous to that of Table 3. We again see a strong relationship between shocks to the basis and real exchange rate changes. The relation becomes stronger later in the sample. We think this is in part because of measurement issues with the basis during the 1970s. Note the spikey behavior of the basis in the 1970s in Figure A.4. In the sample from 1990 onwards, the regression R^2 is 28.4% which is a remarkably strong fit. The coefficient

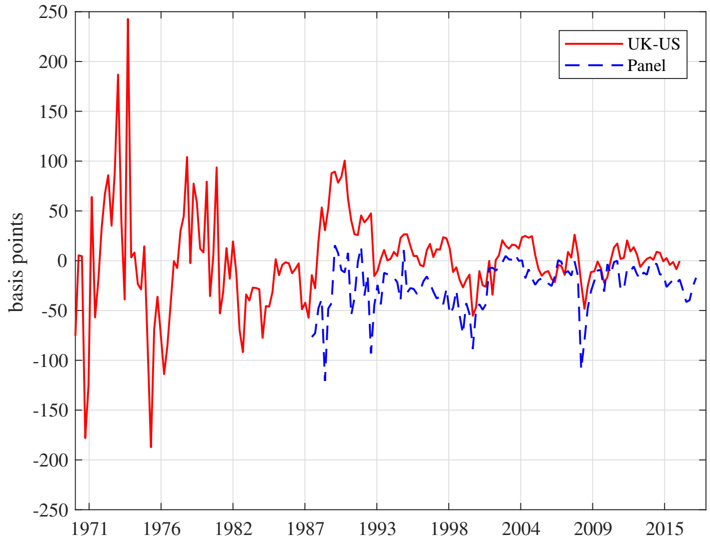

Figure A.3: U.S./UK Treasury Basis

U.S./UK Treasury basis from 1970Q1 to 2017Q2 and the mean Treasury basis across the panel of countries, in basis points. The maturity is one year.

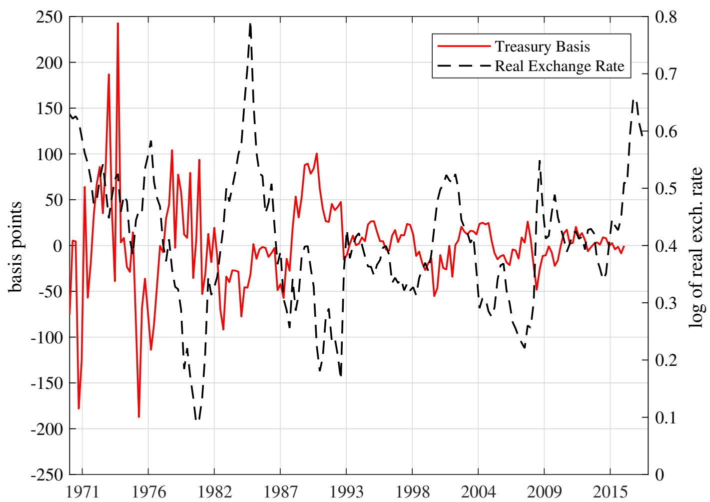

Figure A.4: U.S./U.K. Treasury Bases and Real Exchange Rate

One-year maturity Treasury basis from 1970Q1 to 2017Q2 for U.S./UK, in basis points, and the log real U.S./UK exchange rate.

Table A.11: U.S./UK Treasury Basis and the Spot Real Exchange Rate

The dependent variable is the quarterly change in the quarterly-mean of the log of the spot USD/UK real exchange rate (quoted in GBP-per-USD). The independent variables are the innovation in the quarterly average Treasury basis, $\Delta\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation, and the innovation in the real U.S.-UK interest rate differential. Data is quarterly. OLS standard errors in parentheses.

	1970Q1 - 2016Q2			1980Q1 - 2016Q2	1990Q1 - 2016Q2
	(1)	(2)	(3)	(4)	(5)
$\Delta\overline{x}^{Treas}$	-1.77		-1.74	-3.40	-11.67
	(0.78)		(0.77)	(1.57)	(2.40)
Lag $\Delta\overline{x}^{Treas}$	-1.70		-1.69	-4.59	-3.89
	(0.78)		(0.77)	(1.52)	(2.36)
$\Delta(y^{\$} - \overline{y}^{UK})$		0.13	0.13		
		(0.08)	(0.08)		
R^2	5.0	1.6	6.5	10.3	28.4
N	183	185	183	144	104

in column (4) of the Table indicates that a 10 basis point increase in the basis is correlated with an 0.34% depreciation in the U.S. dollar against the pound. The coefficients using the full sample are smaller than that of Table 3. For column (5), where the sample starts in 1990, the coefficient of -11.67 is similar in magnitude to our earlier estimates.

Column (2) considers the innovation in the interest rate differential as a regressor. In this sample in contrast to the cross-country sample, the interest rate differential has almost no explanatory power for the exchange rate. As noted in the introduction, the prior evidence linking interest rate changes and exchange rates is mixed and this is a clear example of this pattern. The source of the difference is the time period: If we focused on the sample from 1990 onwards, the interest rate differential has explanatory power similar to the result in Table 3. Column (3) includes basis innovations and interest rate differential innovations. The coefficients on the basis in column (3) are almost identical to those of column (1).

The results for the U.S./UK Treasury basis are quite similar to those obtained on the shorter sample for the Panel. Table A.14 reports the results obtained when forecasting the annualized excess returns on a long position in the dollar and a short position in the pound. Panel A considers the results obtained on the entire sample. At short horizons of 3 months, the slope coefficient on the Treasury basis is negative (-2.01). On the other hand, at horizons of 3 years, the slope coefficient is positive and statistically significant: 7.15. This is a quantitatively significant response as well: a one-standard-deviation shock to the U.S./U.K. Treasury basis increases the 3-year return by 2.78%. These regressors jointly explain 28% of the variation in the 3-year excess returns. Panel B and C report results for the pre-and post-crisis sample. The slope coefficients at the 3-year horizon vary from 7.11 in the pre-crisis sample to 8.80 in the post-crisis sample. As was the case in the Panel data, there is no evidence of a basis-induced momentum effect post-crisis: all of the coefficient estimates for the U.S./UK Treasury basis are positive at all forecasting horizons. At the 3-month horizon, these variables jointly explain 54% in the post-crisis sample.

Table A.15 provides more detail by reporting results for each subsample. Panel A examines the 3-year excess returns. Panel B excludes the first year. As was the case for the Panel, excluding the first year increases the size of the slope coefficient. The

Table A.12: U.S./UK Treasury Basis and the Spot Real Exchange Rate

The dependent variable is the quarterly change in the quarterly-mean of the log of the spot USD/UK real exchange rate (quoted in GBP-per-USD). The independent variables are the change in the quarterly average Treasury basis, $\Delta\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the change, and the change in the real U.S.-UK interest rate differential. Data is quarterly. OLS standard errors in parentheses.

	1970Q1 - 2016Q2			1980Q1 - 2016Q2	1990Q1 - 2016Q2
	(1)	(2)	(3)	(4)	(5)
$\Delta\overline{x}^{Treas}$	-1.23		-1.26	-3.05	-11.70
	(0.71)		(0.71)	(1.49)	(2.22)
Lag $\Delta\overline{x}^{Treas}$	-1.85		-1.67	-5.74	-11.69
	(0.71)		(0.71)	(1.45)	(2.22)
$\Delta(y^{\$} - \bar{y}^{UK})$		0.13	0.11		
		(0.06)	(0.06)		
R^2	4.0	2.4	5.8	10.3	33.2
N	183	190	183	144	104

Table A.13: U.S./UK 3M-Treasury Basis and the Spot Real Exchange Rate

The dependent variable is the quarterly change in the quarterly-mean of the log of the spot USD/UK real exchange rate (quoted in GBP-per-USD). The independent variables are the innovation in the quarterly average Treasury basis, $\Delta\overline{x}^{Treas}$, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation, and the innovation in the real U.S.-UK interest rate differential. Data is quarterly. OLS standard errors in parentheses.

	1970Q1 - 2016Q2			1980Q1 - 2016Q2	1990Q1 - 2016Q2
	Average across 3 months				
$\Delta\overline{x}^{Treas}$	-0.92		-0.82	-0.08	-1.99
	(0.48)		(0.48)	(1.02)	(1.23)
Lag $\Delta\overline{x}^{Treas}$	-0.53		-0.52	-3.75	-5.61
	(0.48)		(0.48)	(0.99)	(1.23)
$\Delta(y^{\$} - \bar{y}^{UK})$		0.16	0.14		
		(0.08)	(0.08)		
R^2	2.8	2.1	4.4	9.7	17.3
N	183	185	183	144	104

Table A.14: Forecasting Currency Excess Returns: U.S./UK

The dependent variable is the annualized nominal excess return (in logs) $rx^{fx}_{t\to t+k}$ on a long position in U.S. Treasuries and a short position (equal-weighted) in U.K. bonds over k quarters. The independent variables are the Treasury basis, x^{Treas}, the nominal Treasury yield difference $(y^{\$} - \overline{y}^{*})$, the change in the Treasury basis Δx^{Treas}_{t} from $t-1$ to t, and a lag of the change in the Treasury basis Δx^{Treas}_{t-1}. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses. Data is quarterly from 1970Q1 to 2017Q2.

	x^{Treas}	$y^{\$} - \overline{y}^{*}$	Δx^{Treas}	Lag Δx^{Treas}	R^2
	Panel A: 1970-2017				
3 months	-2.01	2.73	-6.08	-3.03	0.09
	(4.38)	(1.26)	(4.68)	(3.01)	
1 year	0.21	1.79	-3.28	-2.31	0.08
	(3.07)	(0.86)	(2.02)	(1.34)	
2 year	3.26	1.49	-5.40	-3.56	0.13
	(2.17)	(0.59)	(1.72)	(1.06)	
3 year	7.15	1.49	-6.64	-4.31	0.28
	(1.70)	(0.44)	(1.43)	(0.96)	
	Panel B: 1970-2007				
3 months	-2.54	3.09	-4.52	-2.34	0.11
	(4.54)	(1.33)	(4.55)	(2.94)	
1 year	-0.13	1.86	-3.26	-2.59	0.10
	(3.12)	(0.92)	(2.04)	(1.35)	
2 year	3.08	1.48	-5.23	-3.57	0.14
	(2.22)	(0.65)	(1.75)	(1.08)	
3 year	7.11	1.47	-6.61	-4.35	0.30
	(1.76)	(0.51)	(1.47)	(0.98)	
	Panel C: 2007-2017				
3 months	25.73	-8.84	-88.53	-28.34	0.54
	(11.39)	(2.88)	(20.56)	(16.67)	
1 year	22.48	-7.74	-1.24	9.75	0.42
	(7.66)	(2.56)	(7.17)	(7.78)	
2 year	10.89	-3.73	-9.13	3.99	0.30
	(12.09)	(1.48)	(9.12)	(7.33)	
3 year	8.80	-1.82	-4.48	0.57	0.20
	(6.90)	(0.86)	(5.29)	(3.78)	

univariate slope coefficient on the U.S./U.K. increases from 2.86 to 9.18. When we include the other regressors, the slope coefficient changes from 7.15 to 13.84. This coefficient estimate implies that a one-standard-deviation shock to the Treasury basis increases the expected excess return by 5.39% per annum over the next 3 years.

Table A.15: Forecasting 3-year Currency Excess Returns: U.S./UK

In Panel A (B), the dependent variable is the annualized nominal excess return (in logs) $rx^{fx}_{t\to t+12}$ ($rx^{fx}_{t+4\to t+12}$) on a long position in U.S. Treasuries and a short position (equal-weighted) in all foreign bonds over k quarters. The independent variables are the Treasury basis, x^{Treas}, the nominal Treasury yield difference ($y^{\$} - \overline{y}^{*}$), the change in the Treasury basis Δx^{Treas}_t from $t-1$ to t, and a lag of the change in the Treasury basis Δx^{Treas}_{t-1}. Data is quarterly from 1970Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	x^{Treas}	$y^{\$} - \overline{y}^{*}$	Δx^{Treas}	Lag Δx^{Treas}	R^2	x^{Treas}	$y^{\$} - \overline{y}^{*}$	Δx^{Treas}	Lag Δx^{Treas}	R^2
	Panel A: $rx^{fx}_{t\to t+12}$					*Panel B: $rx^{fx}_{t+4\to t+12}$*				
1970-2017	2.86				0.05	9.18				0.06
	(1.16)					(4.12)				
	2.99	1.19			0.14	9.28	0.89			0.06
	(1.07)	(0.48)				(4.16)	(0.89)			
	5.14	1.31	-4.04		0.21	12.08	1.07	-5.67		0.08
	(1.49)	(0.46)	(1.00)			(3.76)	(0.90)	(2.38)		
	7.15	1.49	-6.64	-4.31	0.28	13.84	1.26	-7.84	-4.08	0.09
	(1.70)	(0.44)	(1.43)	0.96		(3.70)	(0.91)	(2.84)	(1.91)	
1970-2007	2.85				0.05	9.68				0.07
	(1.17)					(4.23)				
	2.96	1.16			0.14	9.75	0.76			0.07
	(1.11)	(0.54)				(4.28)	(1.00)			
	5.09	1.28	-3.97		0.21	12.45	0.94	-5.44		0.09
	(1.54)	(0.53)	(1.01)			(3.85)	(1.01)	(2.44)		
	7.11	1.47	-6.61	-4.35	0.30	13.99	1.13	-7.36	-3.64	0.09
	(1.76)	(0.51)	(1.47)	0.98		(3.76)	(1.02)	(2.94)	(1.93)	
2007-2017	6.90				0.07	-19.12				0.05
	(4.52)					(10.64)				
	7.22	-2.06			0.18	-19.60	3.05			0.07
	(4.49)	(0.63)				(9.52)	(1.74)			
	9.11	-1.82	-4.64		0.20	-7.81	4.54	-28.89		0.14
	(5.74)	(0.83)	(4.54)			(8.00)	(1.76)	(13.37)		
	8.80	-1.82	-4.48	0.57	0.20	-3.29	4.62	-31.15	-8.57	0.15
	(6.90)	(0.86)	(5.29)	(3.78)		(13.58)	(1.77)	(12.36)	(15.73)	

Table A.16: Predicting Currency Excess Returns in the U.S./UK Data using 3M Basis

The dependent variable is either $rx^{fx}_{t\to t+k}$ or $(4/k)\Delta s_{t\to t+k}$. The independent variables are the Treasury basis, x^{Treas}, as log yield (i.e. 50 basis points is 0.005), the lagged value of the innovation in the Treasury basis, and the nominal Treasury yield difference $(y^{\$} - \overline{y}^{*})$ in units of log yield. Data is quarterly from 1970Q1 to 2016Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

Panel A: Dependent Variable is Excess Return				
Average across 3 months: Full Sample				
	3-month	1-year	2-year	3-year
$\overline{x}^{Treas}$	-3.07	-1.04	-0.43	0.32
	(1.32)	(1.19)	(1.06)	(0.82)
$y^{\$} - \overline{y}^{*}$	2.88	1.67	1.21	0.99
	(1.17)	(0.86)	(0.58)	(0.46)
Lag $\Delta\overline{x}^{Treas}$		-1.23	-1.49	-1.63
		(0.80)	(0.67)	(0.45)
Constant	0.03	0.02	0.02	0.01
	(0.02)	(0.02)	(0.02)	(0.02)
R^2	10.1	8.9	8.8	9.7
N	185	184	184	180
Average across 3 months: From 1980				
	3-month	1-year	2-year	3-year
$\overline{x}^{Treas}$	-10.28	1.19	-0.35	4.26
	(5.32)	(5.59)	(5.39)	(3.85)
$y^{\$} - \overline{y}^{*}$	2.30	1.95	1.30	1.33
	(1.57)	(0.95)	(0.74)	(0.57)
Lag $\Delta\overline{x}^{Treas}$		-4.62	-2.72	-5.67
		(4.16)	(4.20)	(2.58)
Constant	0.03	0.03	0.02	0.02
	(0.03)	(0.02)	(0.02)	(0.02)
R^2	15.5	11.6	11.9	11.9
N	145	144	144	140

Table A.17: Forecasting Exchange Rate Changes: U.S./UK

The dependent variable is the annualized change in the exchange rate (in logs) $(4/k)\Delta s_{t\rightarrow t+k}$ on a long position in the dollar over k quarters. The independent variables are the Treasury basis, x^{Treas}, the nominal Treasury yield difference $(y^{\$} - \overline{y}^{*})$, the change in the Treasury basis Δx_t^{Treas}, and the lagged change in the Treasury basis Δx_{t-1}^{Treas}. Data is quarterly from 1970Q1 to 2017Q2. Heteroskedasticity and autocorrelation adjusted standard errors in parentheses.

	$\overline{x}^{Treas}$	$\Delta\overline{x}^{Treas}$	Lag $\Delta\overline{x}^{Treas}$	$y^{\$} - \overline{y}^{*}$	R^2
			Panel A: 1970-2017		
3 months	-2.01	1.73	-6.08	-3.03	0.06
	(4.38)	(1.26)	(4.68)	(3.01)	
1 year	0.51	1.02	-3.29	-2.20	0.04
	(3.18)	(0.88)	(2.10)	(1.36)	
2 years	3.72	0.92	-5.49	-3.51	0.09
	(2.16)	(0.59)	(1.70)	(1.04)	
3 years	7.51	1.04	-6.66	-4.24	0.28
	(1.63)	(0.42)	(1.37)	(0.87)	
			Panel B: 1970-2007		
3 months	-2.54	2.09	-4.52	-2.34	0.07
	(4.54)	(1.33)	(4.55)	(2.94)	
1 year	0.18	1.10	-3.29	-2.51	0.05
	(3.25)	(0.93)	(2.13)	(1.37)	
2 years	3.55	0.94	-5.35	-3.55	0.10
	(2.21)	(0.64)	(1.74)	(1.06)	
3 years	7.48	1.07	-6.67	-4.31	0.30
	1.69	(0.47)	(1.41)	(0.90)	
			Panel C: 2007-2017		
3 months	25.73	-9.84	-88.53	-28.34	0.55
	(11.39)	(2.88)	(20.56)	(16.67)	
1 year	22.28	-8.53	-0.10	10.54	0.45
	(7.57)	(2.55)	(7.25)	(7.78)	
2 years	10.12	-4.25	-8.06	4.99	0.33
	(12.25)	(1.47)	(9.26)	(7.62)	
3 years	8.03	-2.19	-3.64	1.33	0.22
	(6.92)	(0.88)	(5.41)	(3.99)	

E Reduced-Form VAR and Impulse Response Functions

We use a Vector Autoregression (VAR) to model the joint dynamics of the interest rate difference, the exchange rate and the Treasury basis. We estimate the VAR separately in both the panel and the U.S./UK data. For this exercise, we define the 12-month U.S. real interest rate $r_t^{\$}$ as $y_t^{\$} - \pi_{t\to t+4}^{\$}$. The foreign real interest rate is similarly defined as $y_t^* - \pi_{t\to t+4}^*$. For the panel, we run a VAR with three variables: the basis, the real interest rate difference, and the log of the real exchange rate $\overline{x}_t^{Treas}$, $r_t^{\$} - \overline{r}_t^*$, and $\overline{q}_t$. The VAR includes one lag of all variables:

$$\boldsymbol{z}_t = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 \boldsymbol{z}_{t-1} + \boldsymbol{a}_t,$$

We identified the VAR(1) as the optimal specification using the BIC. This specification assumes that the log of the real U.S. dollar index is stationary, which seems to be case in this sample period. We order the VAR so that shocks to the basis affect all variables contemporaneously, shocks to the interest rate affect the exchange rate and the interest rate differential but not the basis, and shocks to the exchange rate only affect itself. This ordering implies that nominal and real exchange rates can respond instantaneously to all of the structural shocks. As we discuss, the evidence from the VAR provides support for interpreting our regression evidence causally: shocks to convenience yields drive movements in the exchange rate.

Figure A.5 plots the impulse response from orthogonalized shocks to the basis. The top left panel plots the dynamic behavior of the basis (in units of percentage points), the top right panel plots the dynamic behavior of the interest rate difference (in percentage points), and the bottom left panel plots the behavior of the exchange rate (in percentage points). The pattern in the

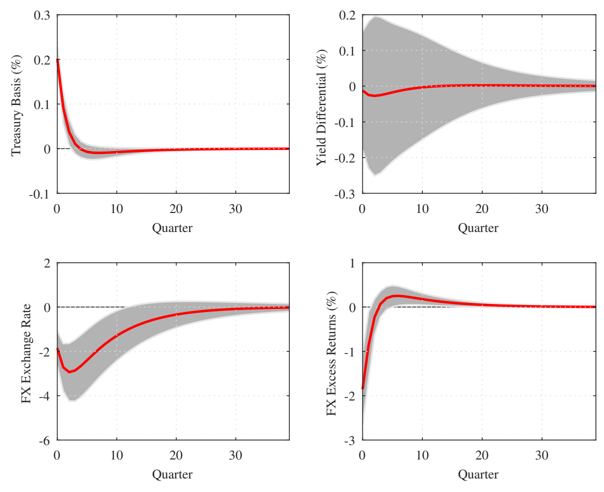

Figure A.5: Dynamic Response to Treasury Basis Shocks: Panel.

The red line plots the impulse response of a one-standard-deviation orthogonalized shock to the average Treasury basis on the basis (top left panel), the real interest rate differential (top right panel), the log real spot exchange rate (bottom left panel), and the quarterly log excess return on a long position in dollars (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. Standard errors were generated using 10,000 Monte Carlo simulations. The VAR is estimated using a sample from 1988Q1 to 2017Q2. The ordering is $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^*, q_t\right]$.

figure is consistent with the regression evidence from the Tables. An increase in the basis of 0.2% (decrease in the convenience yield) depreciates the real exchange rate contemporaneously by about 4% over two quarters. The finding that the depreciation persists over 2 quarters is consistent with the time-series momentum effect discussed earlier. Thus, the exchange rate exhibits classic Dornbusch (1976) overshooting behavior. Then there is a gradual reversal over the next 5 years over which the effect on the level of the dollar gradually dissipates. There is no statistically discernible effect of the basis on the interest rate differential. Finally, the bottom right panel plots the quarterly log excess return on a long position in dollars, $rx_t = q_t - q_{t-1} + i_{t-1}$. Initially, the quarterly excess return drops, but after the first 2 quarters, it is higher than average for the next 15 to 18 quarters, consistent with higher expected returns on long positions in Treasurys.

Interestingly, once we add the basis shock, U.I.P. roughly holds for the dollar against this panel of currencies. Figure A.6 plots the response to the interest rate shocks. The dollar appreciates in real terms in the same quarter by more than 100 basis points in response to a 100 bps increase in the U.S. yields above the foreign yields. Recently, Engel (2016) and Dahlquist and Penasse (2016) have documented that an increase in the short-term U.S. interest rate initially causes the dollar to appreciate, but they subsequently depreciate on average. Once we allow for shocks to the basis, the initial appreciation effect disappears. The bottom right panel of the figure plots the excess return on the currency, and we see that this return is zero after the first quarter indicating that U.I.P. holds once we account for shocks to the basis.

Basis shocks account for a large fraction of the exchange rate forecast error variance, especially at longer horizons. At the one-quarter horizon, basis shocks account for more than 20% of the variance; this fraction increases to 60% at longer horizons. In

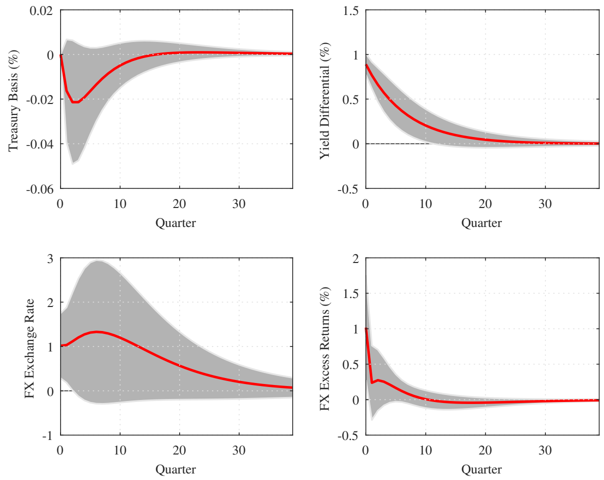

Figure A.6: Dynamic Response to Interest Rate Shocks: Panel.

The red line plots the impulse response of a one-standard-deviation orthogonalized shock to the yield difference on the basis (top left panel), the real interest rate differential (top right panel), the log real spot exchange rate (bottom left panel), and the quarterly log excess return on a long position in dollars (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. Standard errors were generated using 10,000 Monte Carlo simulations. The VAR is estimated using a sample from 1988Q1 to 2017Q2. The ordering is $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^*, q_t\right]$.

contrast, the interest rate shocks account for less than 15% at all horizons. While the initial impact of a one-standard deviation interest rate shock on the dollar is similar to that of a one-standard deviation basis shock (roughly 2%), its effect does not initially build up and is much less persistent. Figure A.7 reports all of the impulse responses.

Importantly, the results are not sensitive to switching the order of the basis and interest rate differential, indicating that we can plausibly interpret the relation between the basis and exchange rate causally: A shock to convenience yields moves both the basis and the exchange rate. We say this because we have allowed for other known determinants of the exchange rate, relative price levels and relative interest rates, and yet recover the same relation between the basis and the exchange rate. Figure A.8 in the Separate Appendix switches the ordering of the interest rate difference and the basis in the VAR. The impulse responses to a basis shock are nearly identical to those of Figure A.6. The exchange rate falls a little under 4% over two quarters and then gradually reverts over the subsequent 2 years. Note that our finding that ordering does not matter need not have been the result. It occurs simply because the reduced form VAR innovations to the basis and the interest rate difference are only weakly correlated. Finally, the variance decomposition also looks independent of ordering.

We report the estimated impulse responses for the U.S./UK, too. The variables included in the VAR are the basis, the interest rate differential and the log of the real exchange rate (GBP-per-USD). The impulse response patterns in this figure are similar to those documented in Figure A.5, but have smaller magnitudes and are less persistent. An increase in the basis of 40 basis points leads to a real depreciation in the dollar against the pound of about 1.8% over two quarters. Then, the effect gradually reverses out over 3 years. We also report the impulse responses that obtain when we switch the ordering of the interest rate differences and the basis. The responses to the basis shock again look identical.[7]

[7] Finally, we also adopted a local projection approach by projecting returns $rx_{t+k-1\to t+k}$ on $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^*, q_t\right]$ and $[\Delta\overline{x}_t, \Delta\overline{x}_{t-1}]$. These yield impulse responses that are quite similar to the ones produced by the Cholesky decomposition. The results are not reported.

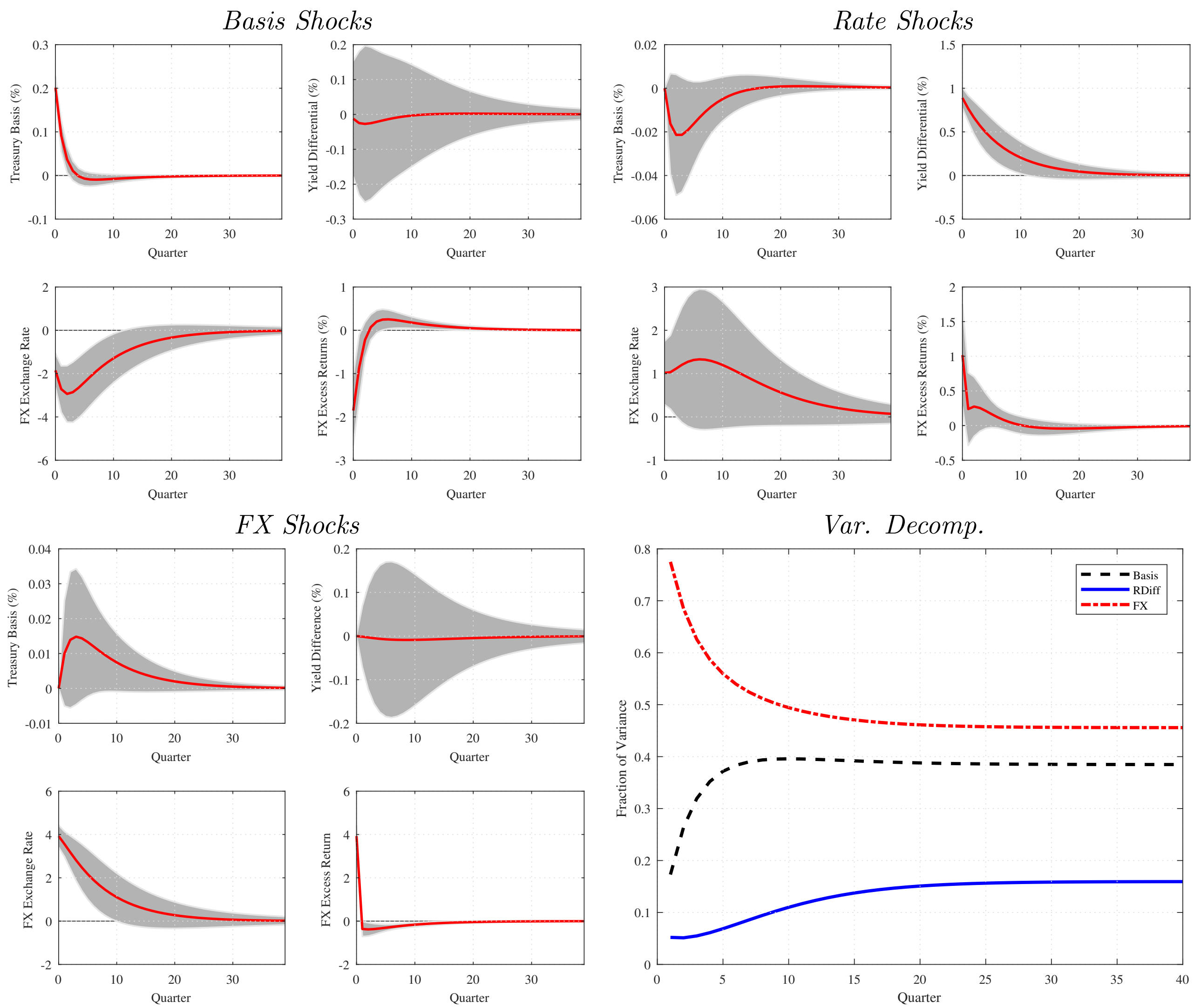

Figure A.7: Panel Impulse Responses.

The red line plots the impulse response of an orthogonalized one-standard-deviation shock on the basis (top left panel), the real interest rate differential (top right panel), the log real spot exchange rate (bottom left panel), and the quarterly log excess return on a long position in dollars (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. The VAR is estimated using a sample from 1988Q1 to 2017Q2. The ordering is $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^*, q_t\right]$.

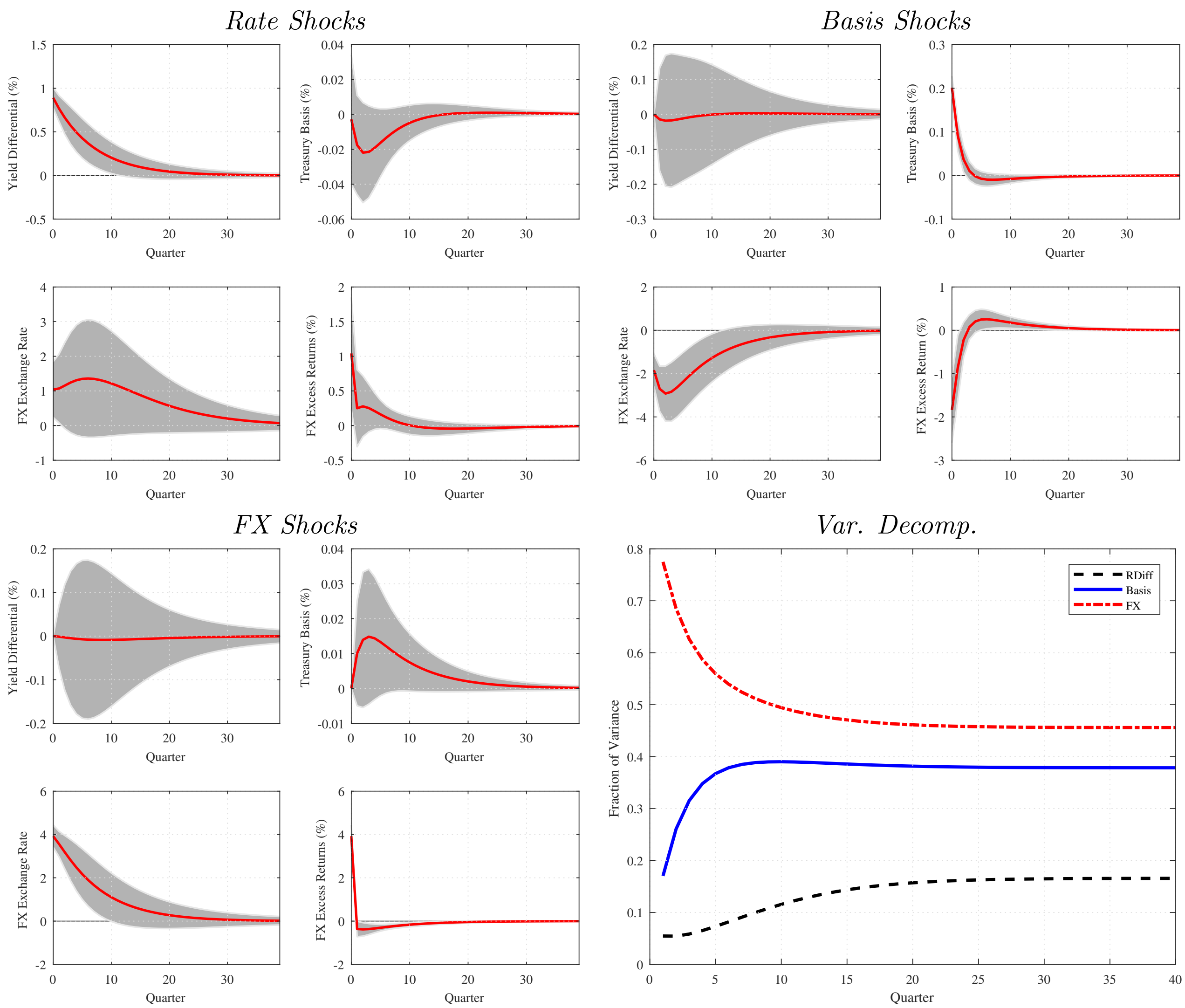

Figure A.8: Panel Impulse Responses: Alternate Ordering.

The red line plots the impulse response of an orthogonalized one-standard-deviation shock on the real interest rate differential (top left panel), the the basis (top right panel), the log real spot exchange rate (bottom left panel), and the quarterly log excess return on a long position in dollars (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. The VAR is estimated using a sample from 1988Q1 to 2017Q2. The ordering is $\left[r_t^{\$} - \overline{r}_t^{*}, \overline{x}_t, q_t\right]$.

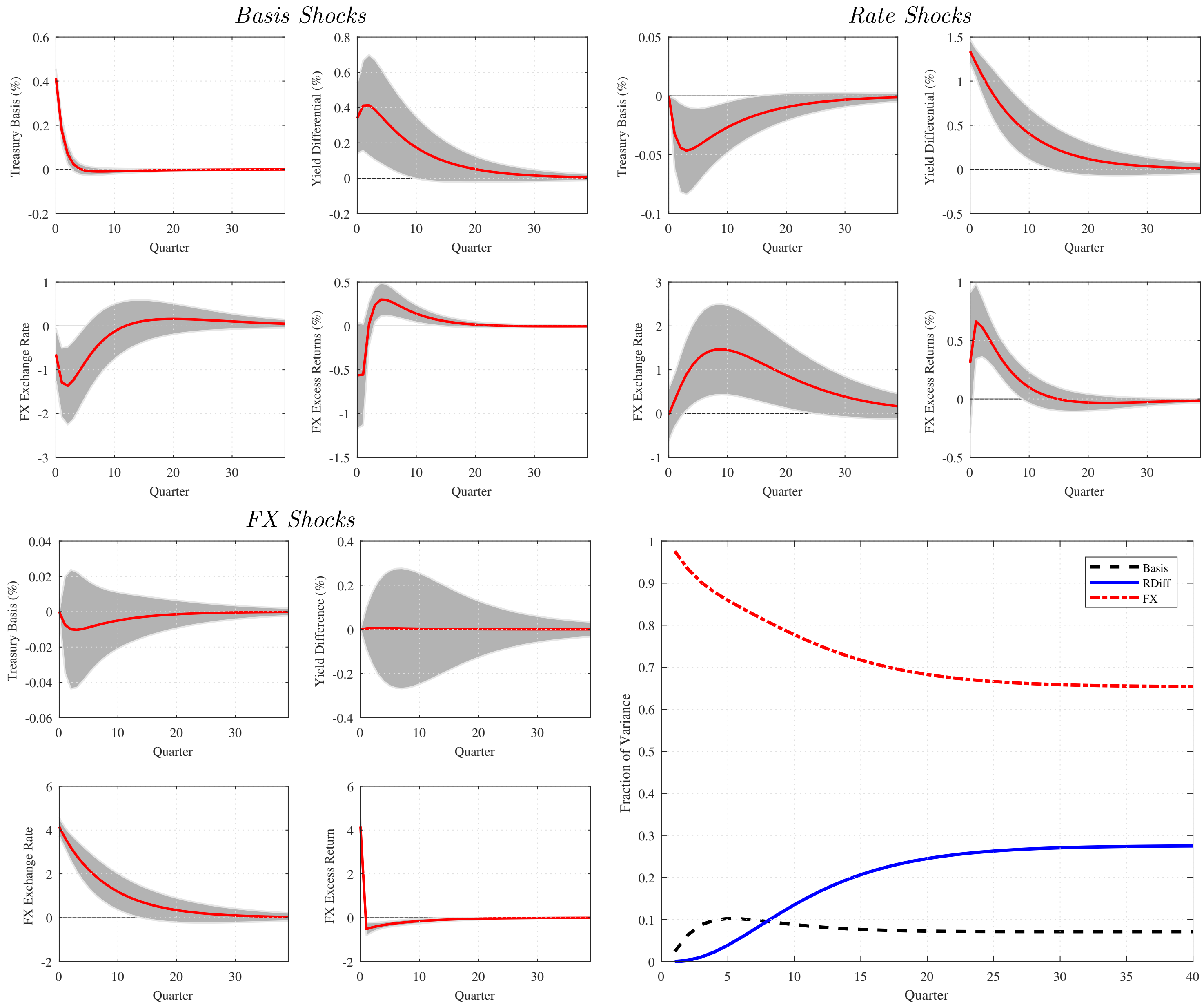

Figure A.9: UK/U.S. Impulse Responses.

The red line plots the impulse response of of a one-standard-deviation orthogonalized shock to the U.S./UK Treasury basis on the basis (top left panel), the real U.S./UK interest rate differential (top right panel), and the log real GBP-per-USD spot exchange rate (bottom left panel), as well as the quarterly excess return (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. The VAR is estimated using a sample from 1970Q1 to 2016Q2. The ordering is $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^{*}, q_t\right]$.

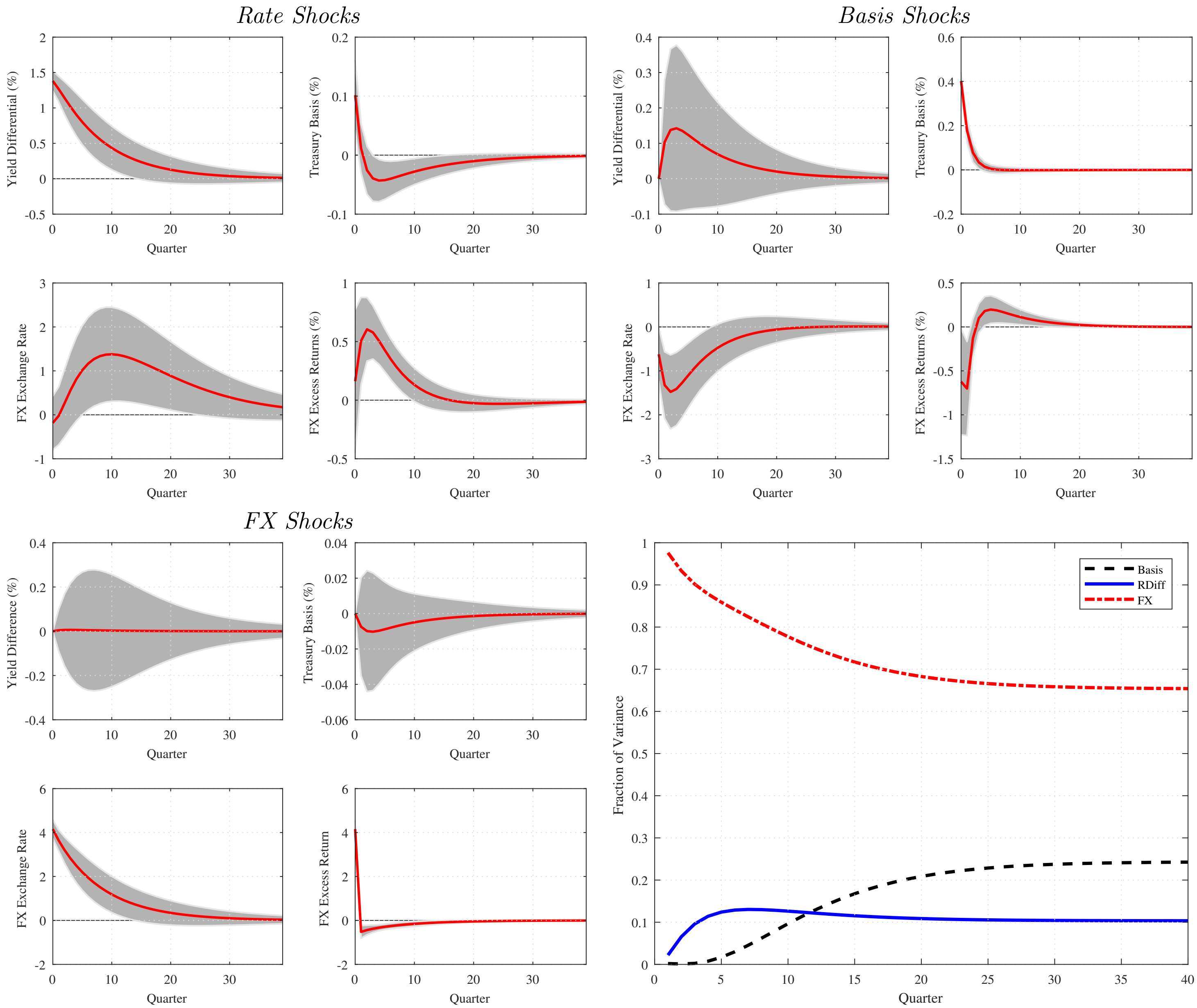

Figure A.10: UK/U.S. Impulse Responses: Alternate Ordering.

The red line plots the impulse response of of a one-standard-deviation orthogonalized shock to the U.S./UK Treasury basis on the basis (top left panel), the real U.S./UK interest rate differential (top right panel), and the log real GBP-per-USD spot exchange rate (bottom left panel), as well as the quarterly excess return (bottom right panel). The units for the y-axis are in percentage points. The grey areas indicates 95% confidence intervals. The VAR is estimated using a sample from 1970Q1 to 2016Q2. The ordering is $\left[\overline{x}_t, r_t^{\$} - \overline{r}_t^{*}, q_t\right]$.