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Liquidity and Exchange Rates: An Empirical Investigation

CHARLES ENGEL
University of Wisconsin, NBER and CEPR

and

STEVE PAK YEUNG WU
University of California, San Diego

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We find strong empirical evidence that the liquidity yield on government bonds in combination with standard economic fundamentals can well account for nominal exchange rate movements. We find impressive evidence that changes in the liquidity yield are significant in explaining exchange rate changes for all the G10 countries, and we stress that the US dollar is not special in this relationship. We show how these relationships arise out of a canonical two-country New Keynesian model with liquidity returns. Additionally, we find a role for sovereign default risk and currency swap market frictions.

Key words: Convenience yield, Liquidity, Meese–Rogoff puzzle

JEL Codes: F31, F41

1. INTRODUCTION

In October 2008 during the 2008 Global Financial Crisis and again in March 2020 at the onset of coronavirus disease 2019 (COVID-19) crisis, the US Federal Reserve cut interest rates more quickly and sharply than central banks of most other major economies, but the US dollar appreciated. According to the standard interest parity condition, low interest rates should depreciate the currency. During both of these periods, analysts refer to a "dash for cash", highlighting the role of demand for liquid assets in influencing exchange rates.[1]

We examine empirically the role of the liquidity return on government bonds in driving exchange rates. Theoretically, Engel (2016) suggests that this return—the non-monetary return that government short-term bonds provide because of their safety, the ease with which they can be sold, and their value as collateral, which is sometimes referred to as the "convenience yield"—may be important in understanding exchange rate puzzles.[2] The intuition for why the government

1. See Bianchi, Bigio and Engel (2021) for a day-by-day analysis of the March 2020 event.

2. Krishnamurthy and Vissing-Jorgensen (2012) and Nagel (2016) study the convenience yield on US Treasury assets. del Negro, Giannone, Giannoni and Tambalotti (2019) find that convenience yields account for the long-run drop in global real interest rates.

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bond convenience yield influences the exchange rate is straightforward. The liquidity that these bonds provide is attractive to investors and influences their investment decisions as if the bonds were paying an unobserved convenience dividend. An increase in the liquidity yield, as measured by the difference between the riskless private bond return and government bond return, will ceteris paribus lead to a currency appreciation much in a similar way that an increase in the interest rate would affect the currency value.

We find for each of the so-called G10 currencies that the relative liquidity yield (the home country yields relative to foreign country yields) has significant explanatory power for exchange rate movements.[3] That is, the role of the liquidity yield in driving exchange rates is not limited to the US but is evident across all the major currencies. Moreover, using guidance from a standard New Keynesian model augmented with a role for liquidity returns on government bonds, we find that the customary determinants of exchange rate movements are statistically and quantitatively important after controlling for the liquidity yields: interest rate differentials and a lagged adjustment term for the real exchange rate (as in Eichenbaum, Johannsen and Rebelo 2021) also drive exchange rate movements.

Our study uses measures of the liquidity yield on government bonds, as constructed by Du, Im and Schreger (2018a). These measures take the difference between a riskless market rate and the government bond rate to quantify the implicit liquidity yield on the government bond. While Du *et al.* (2018a) examine the convenience yield of US dollar government securities, an important point of emphasis for us is that the relationship between this liquidity yield and the exchange rate is not exclusively, or even especially, a dollar phenomenon or a crisis phenomenon. Much of the exchange rate literature in recent years has focused on the special role of the dollar, especially at times of heightened global risk, but the role of the liquidity yield goes beyond that.[4] In fact, when we eliminate the U.S. dollar from our empirical work entirely, the importance of the convenience yield is not diminished. Relatedly, we find, using extensive sub-sample analysis, that the role of the liquidity premium is not a phenomenon driven by financial crises. We find that the relationship is particularly strong for the post-2008 sample.

We subject our results to many robustness tests. We refine our liquidity measures with adjustment for covered interest parity (CIP) deviation and default risk suggested by Du *et al.* (2018a) and construct country-specific liquidity yields rather than relative liquidity yields and find the models perform consistently well. We also find the same pattern for emerging market

3. Namely, Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and US Dollar (USD).

4. A partial list of these studies include the contributions of Bruno and Shin (2015, 2017, 2019), Avdjiev, Du, Koch and Shin (2019a,b) who emphasize the two-way feedback between the dollar and the financial system. Maggiori (2017) links the reserve currency role of the dollar and exchange rates. Lustig, Roussanov and Verdelhan (2011, 2014), Verdelhan (2018), Hassan and Mano (2019), Avramov and Xu (2019), and Sarno, Schneider and Wagner (2012) point to a "dollar factor" in asset returns and returns from a dollar carry trade strategy. Related studies such as Kekre and Lenel (2020), Gourinchas *et al.* (2017), and Greenwood, Hanson, Stein and Sunderam (2020) build economic models of risk premiums and the US dollar. Gabaix and Maggiori (2015), Adrian *et al.* (2018), Itskhoki and Mukhin (2021), Adrian and Xie (2020), and Bianchi *et al.* (2021) relate exchange rates to the global financial system's demand for dollar assets. Caballero, Farhi and Gourinchas (2017, 2020), Gourinchas *et al.* (2017), Lilley, Maggiori, Neiman and Schreter (2019), and Jiang *et al.* (2018, 2020, 2021) relate the exchange rate to demand for safe dollar assets. Rey (2016) and Miranda-Agrippino and Rey (2020) relate US monetary policy, exchange rates and the global financial cycle. Mukhin (2018), Gopinath, Boz, Casas, Díez, Gourinchas and Plagborg-Møller (2019), and Gopinath and Stein (2021) link "dollar currency pricing" to exchange rate behaviour. Lustig and Richmond (2020), Richmond (2019), and Jiang and Richmond (2020) examine gravity models of exchange rate risk premiums in which the dollar plays an outsized role.

currencies but weaker quantitatively. In an exercise in the spirit of Meese and Rogoff (1983), we find our empirical model has a significantly better out-of-sample fit than a random walk model.

The liquidity return or convenience yield may arise from the usefulness of some government securities either as collateral for very short-term loans, or from the ease with which they can be sold for cash. Nickolas (2018) defines liquid assets as: "cash on hand or an asset that can be readily converted to cash. An asset that can readily be converted into cash is similar to cash itself because the asset can be sold with little impact on its value." But there is only a fine distinction between liquidity so defined and "safety" as defined by Gorton (2017): "A safe asset is an asset that is (almost always) valued at face value without expensive and prolonged analysis. By design, there is no benefit to producing (private) information about its value, and this is common knowledge." From these definitions, it is clear that safe assets will be liquid, and liquid assets are safe. The role of safe assets in the global economy has been studies extensively in recent literature. In Caballero, Farhi and Gourinchas (2008), Mendoza, Quadrini and Rios-Rull (2009), Gourinchas and Rey (2011), Maggiori (2017), and Farhi and Maggiori (2018), safe assets play a key role in accounting for global imbalances. Caballero, Farhi and Gourinchas (2015) and Caballero *et al.* (2017) explore the role of a shortage of safe assets and their role in the global financial crisis. Gourinchas and Jeanne (2012) explore the consequences of a shortage of safe assets for the stability of the global financial system.

Engel (2016) proposes a model of liquidity return that can reconcile puzzling empirical evidence on failures of uncovered interest parity. Engel, Lee, Liu, Liu and Wu (2019) and Valchev (2020) document empirical findings that can be explained by a US Treasury liquidity yield.[5] Bianchi, Bigio and Engel (2021) propose an endogenous liquidity yield model from a general equilibrium banking setup with deposit shocks. We contribute to the literature by explicitly measuring the liquidity yield and draw direct linkage to exchange rate determination.[6] Engel (2020) uses the same liquidity yield measures to account for the size of "exorbitant privilege" for the US using a VAR estimation.

Our study is closely related to Jiang, Krishnamurthy and Lustig (2018, 2021), but with the following differences: First, our empirical specification is derived from a simple theoretical general equilibrium model, which is important in understanding the endogenous relationship of liquidity yield, interest rate, inflation, and exchange rate.[7] It also pins down the permanent component of non-stationary nominal exchange rates. The model augments the "canonical" three-equation open economy New Keynesian model with a model of the liquidity yield, which emphasizes that the liquidity yield is a "missing link" that helps solve exchange rate puzzles. Second, consistent with the model, our empirical work finds strong evidence for the role of government liquidity yields, interest rates and adjustment toward purchasing power parity in monthly data for all G10 currencies, while Jiang *et al.* look only at quarterly changes in the US dollar. That is, we emphasize that the liquidity yield is not solely a US dollar story. We also extend the results to emerging market currencies. Finally, using the decomposition of Du *et al.* (2018a), we find additional explanatory power arising from default risk and forward

5. Linnemann and Schabert (2015) also posit a relationship between liquidity returns and exchange rate behaviour. Their paper does not provide an empirical test of the relationship between the liquidity return and exchange rates. Their model postulates a negative relationship between the liquidity yield and interest rates, contrary to the model of Nagel (2016), Engel (2016), and this article, and contrary to the evidence in Nagel (2016) and this article.

6. A bit of our preliminary findings were first reported at a conference at the Bank for International Settlements on "International macro, price determination and policy cooperation" in September 2017. The publicly available slides for that lecture can be found at https://www.bis.org/events/confresearchnetwork1709/programme.htm

7. For example, we note that in our equilibrium model, the liquidity return and the interest rate play somewhat different roles arising from the role of government bond returns and liquidity yield in the monetary policy rule. Thus, interest rates respond endogenously to inflation in a way that the convenience yield does not.

market frictions in a way that is compatible with our model. This latter is important because the premium on government bonds is influenced not only by the liquidity yield, or "convenience yield", of government bonds, but also by default risk and frictions in forward markets for foreign exchange.[8] Jiang *et al.* (2021) further explore the relationship between the LIBOR basis—the deviation from covered interest parity for LIBOR rates—and the US dollar exchange rate. That study also examines the relationship between monetary policy shocks and the convenience yield, investigates the term structure of the convenience yield, and estimates a VAR in order to estimate a variance decomposition that measures the contribution of the convenience yield to exchange rate movements. It also tests the predictability of excess returns using the relative convenience yield.

The exchange rate determination literature has pointed to market imperfections of various stripes—arising for example from balance-sheet constraints, incomplete risk sharing, default risk, noise traders—as potentially playing an important role in exchange rate determination. These studies shed light on what has come to be known at the "exchange-rate disconnect" puzzle, as coined by Obstfeld and Rogoff (2000).[9] Many of these explanations can be collapsed into deviation from uncovered interest parity, which is now introduced as a standard feature in open-economy New Keynesian models to reproduce to some extent the observed volatility of real exchange rates.[10] Indeed, Itskhoki and Mukhin (2021) show that this deviation is key to being able to account for the disconnect puzzle. These models inevitably treat the deviation as an unobserved variable. One interpretation of our model and findings is that the uncovered interest parity deviation is partly observable and can be well-measured by the relative liquidity yield on government bonds.

Liquidity and its role in exchange-rate determination has been explored from a variety of angles. Grilli and Roubini (1992) and Engel (1992) are earlier, related works. Brunnermeier, Nagel and Pedersen (2009), Adrian, Etula and Shin (2018), and Bruno and Shin (2015) consider a liquidity effect on exchange rates arising from banks' balance sheets. One can identify the notions of liquidity in these studies with "funding" liquidity, as defined in Brunnermeier and Pedersen (2009), but other work has looked at the role of "market" liquidity. A prominent recent study is Gabaix and Maggiori (2015) that considers financial constraints that prevent full liquidity to arbitrage international money markets. A related study is Pavlova and Rigobon (2008) which investigates the role of portfolio constraints. Melvin and Taylor (2009), Banti, Phylaktis and Sarno (2012), and Mancini, Ranaldo and Wrampelmeyer (2013) empirically study of the role of liquidity in foreign exchange markets.

There is a long history of attributing a role to the "safe haven" effect on currency values. Fatum and Yamamoto (2016), look at this phenomenon during the global financial crisis, defining a safe currency as "a currency that increases its relative value against other currencies as market uncertainty increases". The idea of a safe haven effect is an old one—see, for example, Dooley and Isard (1985), Isard and Stekler (1985), or Dornbusch (1986). Here, we might argue that during times of global uncertainty, certain assets such as short-term government securities become more valued for their liquidity. There can be other channels through which the safe

8. Avdjiev *et al.* (2019b) document the role of deviations from covered interest parity for the value of the US dollar.

9. Engel (2014) provides a recent survey. Itskhoki and Mukhin (2021) is a recent attempt to build a model to account for the disconnect. One notable determinant of nominal exchange rate movements is the lagged real exchange rate, which arises from adjustment to real exchange rate disequilibrium. This point was made clearly by Mark (1995), and has found strong recent support by Eichenbaum *et al.* (2021).

10. See Kollmann (2002) for an early example.

haven phenomenon works. Farhi and Gabaix (2015) model safe haven currencies as ones that appreciate during times of global downturns, a concept that has been tested empirically by Ranaldo and Söderlind (2010). Obstfeld and Rogoff (2003) speak of risk more generally, which could encompass both the liquidity channel and the hedging channel.

Section 2, which guides our empirical work, presents an equilibrium New Keynesian model in which government bonds pay a liquidity return. Section 3 presents the results of our empirical investigation. Section 4 concludes.

2. A MODEL OF LIQUIDITY AND EXCHANGE RATES

Our model is an analytical version of an off-the-shelf New Keynesian open economy model, but with the addition of the liquidity yield. When augmented with these convenience yields, the model provides guidance for the empirical specification under a general equilibrium framework.

Following Krishnamurthy and Vissing-Jorgensen (2012), Engel (2016), Nagel (2016), and Jiang *et al.* (2018, 2021), we posit that the *ex ante* excess return on short-term government bonds in one country relative to another is attributable to an unobserved liquidity payoff. Let i_t be the one-period interest rate in the "home" country government bonds (we present the model in the context of two countries, "home" and "foreign") i_t^m is the return on a short-term, one-period market instrument. The liquidity premium is the difference in these two rates: $\gamma_t = i_t^m - i_t$. Assume that the market rate is collateralized and the government rate adjusts for the cost of credit default swaps, so these rates represent the default-risk-free returns. The empirical section will adjust the returns for default risk using credit default swap (CDS) data.

Under this formulation, we should observe $\gamma_t > 0$ if the government bond is more liquid. Investors are willing to hold the government bond instead of the market instrument because the government bond is more easily sold on markets or is more readily accepted as collateral. It may be that some agents in the economy have no need for liquidity, in which case their holdings of the government bonds are zero. In particular, foreign agents may hold no home government bonds because they do not value the liquidity of those assets. But private agents cannot short government bonds—that is, private agents (in either economy) cannot borrow at the rate i_t, because the assets they issue do not have the same liquidity as government bonds.

Analogously, in the foreign country, there is a liquidity yield given by $\gamma_t^* = i_t^{*m} - i_t^*$, where the * superscript denotes the foreign-country equivalents of the home-country variables.

We assume there is a deviation from uncovered interest parity for the market instruments, r_t, that is stochastic, exogenous, and uncorrelated with the other shocks (monetary and liquidity) in the model. We remain agnostic about the source of this deviation. r_t could be a deviation from rational expectations, some sort of market friction, or perhaps a foreign exchange risk premium. In Jeanne and Rose (2002), Devereux and Engel (2002), and Itskhoki and Mukhin (2021), this term arises because of the presence of noise traders. We assume that r_t is uncorrelated with other shocks introduced into the model, to the monetary policy rule and to the liquidity return:

$$i_t^{*m} + E_t s_{t+1} - s_t - i_t^m = r_t, \tag{1}$$

where s_t is the log of the exchange rate (expressed as the home currency price of the foreign currency.)[11]

11. A simplification implicit here is that the standard "Fama" regression would not reject the null hypothesis. Under our specification for monetary policy, introduced below, with inflation predetermined and monetary shocks that are independent of the shocks to r_t, we should not be able to reject uncovered interest parity using our measures of returns on market instruments. Supplementary Appendix A1. shows, however that during our time sample, the null hypothesis is not rejected in the Fama regression for any of the G10 currencies.

Let η_t be the liquidity return on home relative to foreign government bonds:

$$\eta_t = \gamma_t - \gamma_t^* = \left(i_t^m - i_t\right) - \left(i_t^{m*} - i_t^*\right) = \left(i_t^m - i_t^{m*}\right) - \left(i_t - i_t^*\right). \tag{2}$$

Then, we can rewrite (1) as:

$$i_t^* + E_t s_{t+1} - s_t - i_t = \eta_t + r_t. \tag{3}$$

The expected excess return on foreign one-period government bonds (relative to home bonds) is determined in part by the liquidity yield of home government bonds relative to foreign bonds. When home bonds are more liquid, the foreign bonds must pay a higher expected monetary return.

Now, iterate equation (3) forward, as in Campbell and Clarida (1987) and others:[12]

$$\begin{aligned} s_t = & -E_t \sum_{j=0}^{\infty} \left(i_{t+j} - i_{t+j}^* - \left(\overline{i - i^*}\right) \right) \\ & - E_t \sum_{j=0}^{\infty} \left(\eta_{t+j} - \bar{\eta}\right) - E_t \sum_{j=0}^{\infty} \left(r_{t+j} - \bar{r}\right) + \lim_{k \to \infty} \left(E_t s_{t+k} - k\left(\overline{s_{+1} - s}\right)\right). \end{aligned} \tag{4}$$

We assume that the interest differential, $i_t - i_t^*$; the liquidity return, η_t, and the u.i.p. deviation, r_t, are all stationary random variables, but s_t follows a unit root process.[13] An overbar represents unconditional means: $\overline{i - i^*}, \bar{\eta}, \bar{r}$. Here, $\lim_{k \to \infty} \left(E_t s_{t+k} - k\left(\overline{s_{+1} - s}\right)\right)$, which is a random variable when the exchange rate has a unit root, is the permanent component of the nominal exchange rate—in the sense that Beveridge and Nelson (1981) use that term in their permanent-transitory decomposition. The term $\overline{s_{+1} - s}$ represents the trend in the log of the nominal exchange rate.

There is consensus that nominal exchange rates among high-income countries contain unit roots. For example, if monetary policy is set by a rule for money supplies, any permanent change in the money supply would lead to a permanent change in the nominal exchange rate. If monetary policy is set by an interest-rate rule such as a Taylor rule, the exchange rate will contain a unit root unless the interest rate rule targets the nominal exchange rate.[14]

Equation (4) already points to the intuition of our empirical specification. It says that when the infinite sum of the expected current and future home interest rates rises relative to the expected infinite sum of current and future foreign interest rate, the home currency appreciates (s_t falls.) That is a well-known channel of influence, which is at work in, for example, the famous Dornbusch (1976) model. A higher relative liquidity return on home government bonds also leads to an appreciation of the domestic currency. In this equation, the liquidity return and the interest rate are just two components of the return on government bonds, and so their impact on the exchange rate is identical. In the model below, the interest rates and liquidity return play different roles in the monetary policy rule and are endogenously determined.

Equation (4) is not a full model of nominal exchange rate determination. Comparative statics exercises that change the interest rate differential or the components of expected excess returns holding the permanent component of the exchange rate constant can be misleading. A full macro

12. Engel (2016) and Jiang *et al.* (2021) relate the expected excess returns to the liquidity yield.

13. Technically, we assume $i_t - i_t^*, \eta_t$, and r_t are square summable, which insures that the infinite sums converge. Any finite order ARMA process, for example, is square summable.

14. See Benigno and Benigno (2008). See Engel and Wu (2021) for recent evidence on the unit root in nominal exchange rates.

model is necessary to understand how the components on the right-hand side of equation (4) relate to the exchange rate in equilibrium. For example, not all nominal interest rate changes are the same. In a traditional monetarist model of exchange rates, a permanent one-time increase in the monetary growth rate in the home country would immediately raise inflation and therefore raise the inflation premium incorporated in the nominal interest rate. $i_{t+j}-i^*_{t+j}$ would increase for all time periods but that also implies an increase in the unconditional mean of the relative interest rates, $\overline{i-i^*}$. In that case, there would be no change in the first term on the right-hand side of equation (4): $E_t\sum_{j=0}^{\infty}\left(i_{t+j}-i^*_{t+j}-\left(\overline{i-i^*}\right)\right)$ would be unaffected. However, this change would lead to an increase in the permanent component of the exchange rate. The size of the increase is model-dependent, but a classic result is that an increase in the growth rate of x percent leads to an immediate permanent depreciation of greater than x percent, which the literature referred to as the "magnification effect".[15] The conclusion is that equation (4) by itself, which represents the international financial market equilibrium condition, is not sufficient to determine the exchange rate. In order to determine the exchange rate, we need a model of the determination of interest rates, and of the permanent component of the nominal exchange rate. One cannot infer the effects of the liquidity yield or interest rates on exchange rates by changing one of the components of (4) holding the others constant, because in a dynamic general equilibrium, the components interact.[16]

We adapt the model from Engel (2016), based in turn on Nagel (2016), in which the liquidity return on the home bond is positively related to the interest rate:

$$\eta_t=\alpha\left(i_t-i^*_t\right)+v_t,\alpha>0. \tag{5}$$

Appendix A1 derives this equation, extending the analysis of Engel (2016). The positive relationship between the relative liquidity return and the interest differential arises as in Nagel (2016). When the monetary authority tightens monetary policy by reducing the supply of money and raising interest rates, liquid assets that can substitute for money become more valued for their liquidity services and so pay a higher liquidity return.[17]

The remainder of the model adopts a New Keynesian framework. We assume that prices in each country are sticky in nominal terms and set one period in advance. We posit that there is local-currency pricing, so that each firm, in both countries, sets two prices—one in home currency for sale in the home country, and one in foreign currency for sale in the foreign currency. Supplementary Appendix B1 derives the home relative to foreign Phillips curve,

$$\pi_t-\pi^*_t=\theta q_{t-1}+E_{t-1}s_t-s_{t-1}. \tag{6}$$

In this equation, q_{t-1} is the log of the real exchange rate (the price of the consumer basket in the foreign country relative to the home country), π_t is home consumer price inflation between $t-1$ and t, and π^*_t is foreign consumer price inflation. Note that because prices are set one period in advance, the inflation rates, π_t and π^*_t, are observable at time $t-1$. Under this specification of price adjustment, the real exchange rate is a stationary random variable and long-run purchasing power parity holds. The pricing to market disequilibria are expected to dissipate over time.

15. See, for example, Frenkel (1976).

16. Here, we differ from Jiang *et al.* (2021), who take nominal interest rates as exogenous and assume the nominal exchange rate is stationary.

17. This equation does not necessarily imply a positive relationship between the relative liquidity yields and the relative interest rates in equilibrium. We introduce below a monetary policy rule in which these two variables are negatively related. An increase in v_t here leads to an increase in η_t, which in turn will lead monetary policy makers to reduce $i_t-i^*_t$, leading to a negative correlation between $i_t-i^*_t$ and v_t.

Small modifications to the standard open-economy Phillips curve are introduced here to motivate our empirical model of the exchange rate in an intuitive way. As is well-known from Benigno (2004), price stickiness would not matter at all for the adjustment of the real exchange rate with a standard Calvo-pricing equation, unless interest-rate smoothing is introduced into the monetary policy rule. Engel (2019) shows how the Phillips curve here, along with serially correlated errors in the monetary policy rule produces very similar real exchange rate behaviour as the Calvo pricing model with interest rate smoothing, but this model is more analytically convenient.

The final component of the model is the characterization of monetary policy behaviour. We assume that the monetary authority has control of a policy rate that is a weighted average of the government bond rate and the market rate, which it uses to target inflation. In the home country:

$$(1-\psi)\, i_t + \psi i_t^m = \sigma \pi_t + u_t,\ 0 \leq \psi \leq 1. \tag{7}$$

In practice, the policy rate is closely aligned with the government bond rate in the countries in this study.[18] We use this specification to avoid introducing yet another interest rate, the monetary policy rate, and remain agnostic on the correct value of ψ. None of the qualitative conclusions from the model depend on the value of ψ. Another interpretation of this equation is that the government interest rate is the policy rate, and the policy maker targets the liquidity premium by lowering the policy rate when the market places a high value on liquidity:

$$i_t = \sigma \pi_t - \psi \left(i_t^m - i_t\right) + u_t.$$

The stability condition in this model is given by $\sigma > \frac{1+\alpha}{1+\alpha\psi}$, and we assume $\sigma > 1$. u_t is a deviation from the monetary policy rule. There is an analogous equation in the foreign country. Subtracting the foreign Taylor rule from the home Taylor rule gives us:

$$i_t - i_t^* = \sigma\left(\pi_t - \pi_t^*\right) - \psi \eta_t + u_t - u_t^*. \tag{8}$$

The relative error terms in the monetary rules follow a first-order autoregressive process:

$$u_t - u_t^* = \delta\left(u_{t-1} - u_{t-1}^*\right) + \xi_t,\ 0 \leq \delta < 1 \tag{9}$$

where ε_t is a mean-zero, i.i.d. random variable.

The exogenous variables are monetary shocks (in equation (8)), the uncovered interest parity shocks in equation (1), and liquidity shocks in equation (5). We have already noted that monetary shocks are assumed to be follow an AR(1) process. We assume that there is persistence in liquidity, and that v_t also follows a first-order autoregressive process:

$$v_t = \rho v_{t-1} + \varepsilon_t, \tag{10}$$

where ε_t is mean-zero, i.i.d., and $0 \leq \rho < 1$. Furthermore, we assume that the deviation from uncovered interest parity for market interest rates also follows an autoregressive process given by:

$$r_t = \varsigma r_{t-1} + \omega_t,\ 0 \leq \varsigma < 1. \tag{11}$$

Equations (3), (5), (6), and (8)—the international financial market equilibrium condition, the model of the liquidity premium, the (relative home to foreign) open economy Phillips curve,

18. Supplementary Appendix A2 provides evidence to support this statement.

and the (home relative to foreign) monetary policy rule—give us a complete dynamic system for the real exchange rate, inflation and interest rates. The model incorporates slow adjustment of the real exchange rate because of nominal price stickiness, governed by the parameter θ, the fraction of the firms that reset their price optimally each period. As Eichenbaum *et al.* (2021) have recently emphasized, empirically almost all the adjustment of real exchange rate comes through adjustment by the nominal exchange rate. Eichenbaum *et al.* (2021) demonstrate that this empirical regularity can be captured in a New Keynesian model with strong inflation targeting (large value of σ) which leads to a low inflation variance even if the variance of the real exchange rate is large. That regularity indeed does not depend on sluggish price adjustment but would be true even under flexible goods prices when monetary policy stabilizes the nominal price level.

The model can be solved by hand.[19] For the real exchange rate, we find:

$$
\begin{aligned}
q_t = & -\left(\frac{\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)}{\theta(1+\alpha\psi)}\right)(\pi_t-\pi_t^*) \\
& -\left(\frac{(1+\alpha)[\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)]}{\theta(1+\alpha\psi)(\sigma(1+\alpha)-\delta(1+\alpha\psi))}\right)(u_t-u_t^*) \\
& -\left(\frac{(1-\psi)[\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)]}{\theta(1+\alpha\psi)[\sigma(1+\alpha)-\rho(1+\alpha\psi)]}\right)v_t-\left(\frac{\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)}{\theta[\sigma(1+\alpha)-\zeta(1+\alpha\psi)]}\right)r_t
\end{aligned} \quad . \tag{12}
$$

The inflation variables at time t are predetermined, so (12) expresses the real exchange rate in terms of predetermined and exogenous variables. A relative monetary tightening in the home country (an increase in $u_t-u_t^*$) causes a real appreciation of the home currency. Similarly, an increase in the liquidity yield on home government bonds leads to a real appreciation. Note that as inflation targeting becomes more stringent, so σ is larger, the real exchange rate reacts more to monetary policy shocks if $\delta<1-\theta$. If $\rho<1-\theta$, a larger σ increases the response of the real exchange rate to changes in the relative liquidity return. We assume in all following discussion that both preceding inequalities are satisfied. Also, the greater price stickiness (smaller θ), the larger the response of the real exchange rate to monetary policy shocks and the relative liquidity returns.

Referring to equation (4), the expected values of future interest rates depend on all of the shocks to the model, as do the expected values of future convenience yields. In this model, the expected future interest rates, liquidity yield and risk premiums are all endogenous functions of the state variables on the right hand side of equation (12). The real exchange rate can be solved in terms of these time t variables because shocks have only first-order serial correlation. The final term in (4), the permanent component of the nominal exchange rate, is endogenously determined by the convergence of prices and the exchange rate toward the unconditional mean of the real exchange rate.

Our empirical analysis aims at explaining movements in the log of the nominal exchange rate, s_t-s_{t-1} in terms of observable variables. We derive:

$$
s_t-s_{t-1}=\beta_1 q_{t-1}+\beta_2(\eta_t-\eta_{t-1})+\beta_3\left(i_t-i_t^*-(i_{t-1}-i_{t-1}^*)\right)+\beta_4\eta_{t-1}+\beta_5\left(i_{t-1}-i_{t-1}^*\right)+z_{j,t} \tag{13}
$$

19. The full derivation of the model is in Supplementary Appendix B2.

where

$$\beta_1 = -\left(\frac{\sigma(1+\alpha)(1-\theta-\delta)}{\sigma(1+\alpha)-\delta(1+\alpha\psi)}\right) < 0$$

$$\beta_2 = -\left(\frac{[\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)][\sigma(1+\alpha)-(\delta(1-\psi)+\rho\psi(1+\alpha))]}{\theta(\sigma(1+\alpha)-\delta(1+\alpha\psi))(\sigma(1+\alpha)-\rho(1+\alpha\psi))}\right) < 0,$$

$$\beta_3 = -\left(\frac{[\sigma(1+\alpha)-(1-\theta)(1+\alpha\psi)][(1+\alpha)(\sigma-\rho)+\alpha\delta(1-\psi)]}{\theta(\sigma(1+\alpha)-\delta(1+\alpha\psi))(\sigma(1+\alpha)-\rho(1+\alpha\psi))}\right) < 0.$$

and $z_t = -\left(\frac{\sigma(1+\alpha)+\theta-1}{\theta(\sigma(1+\alpha)-\varsigma)}\right)\omega_t + z_1 r_{t-1}$.

The full expressions for β_4, β_5, and z_1 are presented in Supplementary Appendix B3.

This specification for the depreciation of the exchange rate includes, first, an error correction term as the nominal exchange rate adjusts to disequilibrium in the real exchange rate. Second, the change in the interest differential affects the exchange rate as in standard New Keynesian models. Third, the change in the relative liquidity return on government bonds plays a role in influencing the exchange rate. Lagged levels of the relative interest differentials and liquidity returns capture the dynamic adjustment. Under the parameter restrictions of the model—$\sigma > 1$, $\alpha > 0$, $0 \leq \theta < 1, 0 \leq \rho < 1$, and $\rho < 1-\theta$—ceteris paribus, an increase in q_{t-1}, and increase in $i_t - i_t^* - \left(i_{t-1} - i_{t-1}^*\right)$, and an increase in $\eta_t - \eta_{t-1}$ all lead to a decline in $s_t - s_{t-1}$. That is, the home currency appreciates to correct for a real undervaluation, and it appreciates in response to a relative increase in either the home interest rate or the home liquidity return. The error term, z_t, is a function of the dynamics of the deviation from uncovered interest parity, which is assumed not to be observable by the econometrician. It is by construction uncorrelated with the explanatory variables in the regression. The derivation implies that there may be serial correlation in the regression error. However, as shown in Supplementary Appendix B5, for parameters calibrated to the data, the serial correlation of the residual is very low (around 0.02, for example), as it is in the data.[20]

The estimating equation (13) isolates the effects of an increase in liquidity demand. Controlling for the government interest rate differential, $i_t - i_t^*$, as in (13), means that an increase in the liquidity premium, η_t, raises the relative market interest differential, $i_t^m - i_t^{m*}$. Intuitively, if government bonds in the home country become more valuable for their liquidity services, but $i_t - i_t^*$ does not change, then the market interest rate on home securities must rise, leading to incipient capital inflows and an immediate appreciation of the home currency.

Before turning to the data, we note a few features of our empirical specification based on (13). As in our model, we follow convention and treat nominal exchange rates as non-stationary random variables. Considering much evidence, from Mark (1995) to more recent empirical evidence in Engel (2016) and Eichenbaum *et al.* (2021), we choose to treat the real exchange rate as stationary, and the nominal exchange rate adjusts in the direction of restoring purchasing power parity. Relative interest rates and relative liquidity returns are also modelled as stationary. We allow dynamics by including contemporaneous and lagged values of these variables. Because these variables are serially correlated, we enter them in the specification as in (13) with the first difference in the returns and the lagged level of the returns. This reduces the multi-collinearity that would be present if these variables were included in contemporaneous and lagged levels and gives us the natural interpretation that changes in relative interest rates and changes in relative liquidity yields influence changes in the log of the nominal exchange rate.

20. See Supplementary Appendix B3 for the derivation of equation (13), and of the serial correlation of the residual.

β_3 measures the impact of monetary policy shocks, $u_t - u_t^*$, on s_t, while β_2 quantifies the effect of shocks to the relative liquidity yield, v_t, on the log of the exchange rate. To see this, first observe from the relative Taylor rules, (8), that the home relative to foreign interest rate differential depends on relative inflation, η_t and the monetary policy shocks. However, from equations (6) and (3), we see that relative inflation, $\pi_t - \pi_t^*$ is predetermined and a function of the lagged interest rate differential and liquidity yield, $i_{t-1} - i_{t-1}^*$ and η_{t-1}.[21] Because these latter two variables and η_t are controlled for in (13), the independent effect of $i_t - i_t^*$ arises only from the monetary policy shocks. Equation (5) finds η_t is a function of the interest rate differential as well as the independent shocks to liquidity. Since the regression equation controls for $i_t - i_t^*$, the independent effects of the shocks to liquidity are measured by the coefficient on η_t.

3. EMPIRICAL INVESTIGATION OF GOVERNMENT BOND LIQUIDITY AND EXCHANGE RATES

In this section, we present our empirical results. We first describe how we construct the measure of government bond liquidity in 3.1. Section 3.2 presents our baseline result that the change in the relative government bond liquidity returns is strongly correlated with exchange rate movements. We show our results are robust to controlling for certain market frictions in Section 3.3. In Section 3.4, we further confirm that country-specific government bond liquidity matters. Finally, in Section 3.5, we conduct an out-of-sample fit exercise a la Meese and Rogoff (1983) and find that our model's prediction significantly outperforms a random walk model.

Throughout the section, we denote the foreign variable as X_t^* if the context is not country j specific. For example, we use i_t^* for the foreign interest rate on a government bond. Whenever needed, we denote the variables of a foreign country j as $X_{j,t}^*$, for example, $i_{j,t}^*$ for the interest rate of a government bond for the foreign country j.

3.1. *Construction of liquidity measure*

The word "liquidity" appears in different economic contexts with different meanings. Here, it refers to a non-observable non-pecuniary return that investors enjoy when holding the asset.

We consider two measures of the term $i_t^m - i_t^{m*}$ in equation (1). The first uses LIBOR swap rates:

$$\hat{\eta}_t = IRS_t - IRS_t^* + i_t^* - i_t, \tag{14}$$

where IRS_t (IRS_t^*) refers to the home (foreign) return on LIBOR swaps. The second uses the forward premium:

$$\tilde{\eta}_t \equiv f_{t,t+1} - s_t + i_t^* - i_t, \tag{15}$$

where $f_{t,t+1}$ is the log of forward rate and s_t is the log of the spot exchange rate, both expressed in home currency price of a foreign currency.

There are two ways to interpret η_t. First, as the term $(i_t^m - i_t) - (i_t^{m*} - i_t^*)$ suggests, it is a relative measure of difference between marketable securities and government bond yield in the home and foreign country. This interpretation accords well with the $\hat{\eta}_t$ measure in equation (14), where we are using LIBOR swap rates as the empirical counterpart of our model's market interest rates.

21. Relative inflation would also be a function of lagged r, but we have already argued that because serial correlation is essentially zero in our regressions, the impact of lagged r as a "left out" variable in our regressions is minimal.

Second, as described by $f_{t,t+1}-s_t+i_t^*-i_t$, the first three terms can be understood as the payoff of a synthetic home government bond that is constructed by buying the foreign government bond and eliminating exchange rate risk by entering a forward contract. Since the home government bond and the synthetic home government bond pay equivalent pecuniary returns, the difference between the two gives a measure of the relative difference in liquidity services the home and foreign government bonds provide. Under this interpretation, we are measuring relative liquidity yields by looking at relative government interest rates, correcting for foreign exchange risk. This motivates our baseline measure $\tilde{\eta}_t$ in equation (15).

We employ the procedure developed by Du *et al.* (2018a) to obtain $\hat{\eta}_t$ and $\tilde{\eta}_t$ for any pair of home currency *i* and foreign currency *j* (90 pairs in total) for the G-10 currencies. To give a sense of how this liquidity measure behaves, we plot the liquidity measure ($\tilde{\eta}_t$) against the nominal exchange rate of each home currency *i*and foreign currency *j* in Figure 1. For each time period, we take a simple average across foreign currency *j* to improve visual representation. There is already a negative relationship between the mean exchange rate and mean liquidity measure, meaning a higher government bond liquidity relative to the rest of the G10 currency country is associated with a strong currency contemporaneously.

Unless otherwise specified, our study uses end-of-month monthly data from January 1999 to January 2018.[22] We use exchange rates and forward rates from Thomson Reuters Datastream. The consumer price indexes and unemployment rates are from the IMF IFS. The government yield data is obtained from Bloomberg, Datastream and central banks. The LIBOR swap rates are from Bloomberg. The credit default swap data is from Bloomberg and IHS Markit. We provide the data source details in Appendix A2 and summary statistics for the variables we used in Appendix A3. Supplementary Appendix A reports a large number of robustness checks. We employ panel fixed effect regressions in all the reported estimates to make use of cross-country time series information but at the same time allow for time-invariant heterogeneity. To account for the possibility of cross-sectional correlated estimation errors, we report standard errors that allow for non-diagonal covariance of the error terms. We estimate the regression using ordinary least squares (OLS). The error terms estimated from the OLS are then used to construct estimates of the variance–covariance matrix of the error term. Consistent statistical inference (e.g. significance) can be conducted using this estimated variance–covariance matrix.[23]

3.2. *Baseline results*

To investigate the empirical relationship between government bond liquidity and exchange rates for the G10 countries, we estimate the following panel monthly fixed effect regression from equation (13):[24]

$$\Delta s_{j,t}=\alpha_j+\beta_1 q_{j,t-1}+\beta_2\left(\Delta\eta_{j,t}\right)+\beta_3\left(\Delta i_{j,t}^R\right)+\beta_4\eta_{j,t-1}+\beta_5 i_{j,t-1}^R+u_{j,t}, \tag{16}$$

22. Whenever needed, we linearly interpolate the quarterly variable to monthly variable. For example, we interpolate the Australia and New Zealand CPI to obtain monthly real exchange rates.

23. The standard errors reported in the tables do not correct for time dependence of the residuals. The G10 exchange rates are famously nearly random walks at the monthly frequency. The standard errors are more precisely estimated if this assumption is true. In Supplementary Appendix A15, we report three of our baseline regressions results with Driscoll and Kraay (1998) standard errors, which account for both cross-sectional correlation and autocorrelation. The conclusions on statistical significance are unchanged.

24. We have followed the practice of the empirical exchange rate literature, which might be described as regularization, to avoid overfitting the model with unrestricted coefficients across foreign countries. See Supplementary Appendix B6 for a detailed discussion.

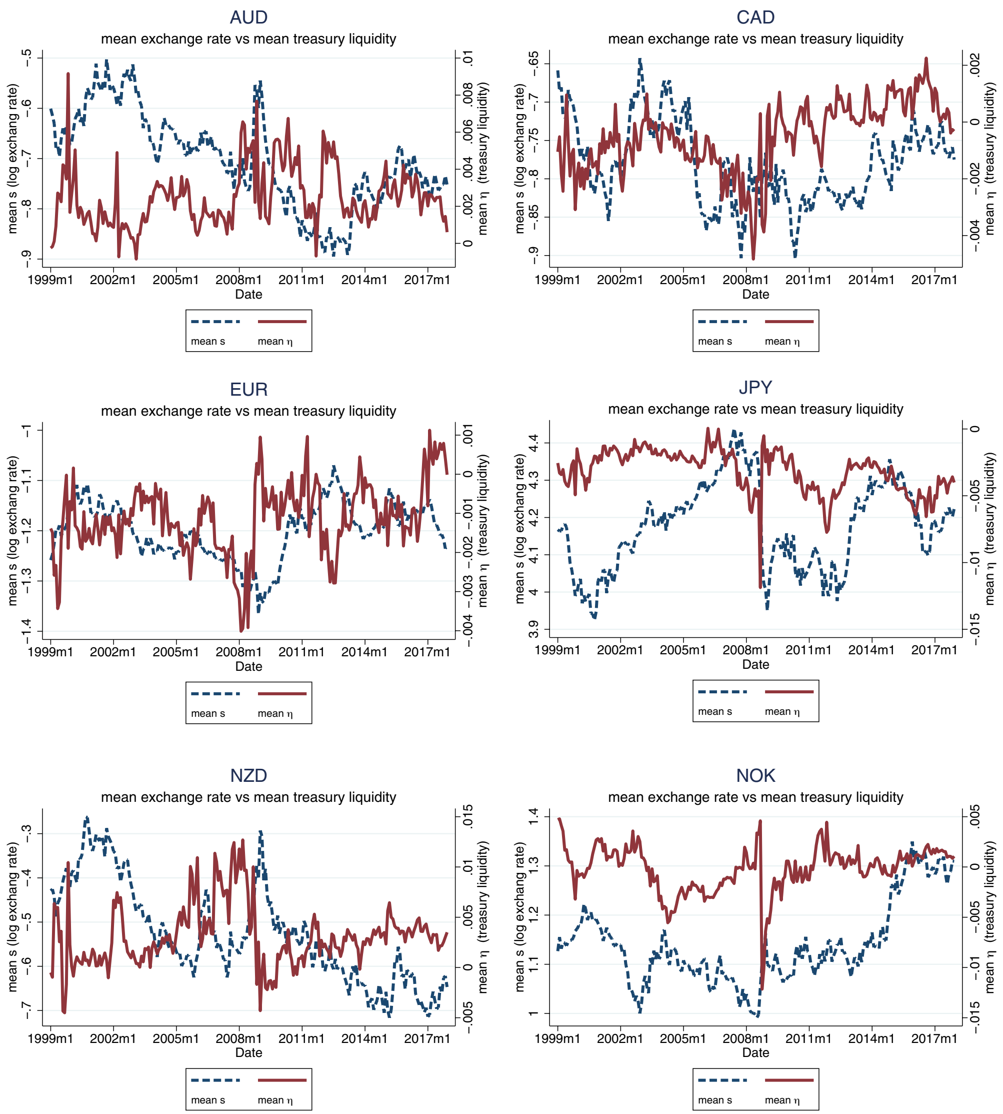


FIGURE 1

Time-series plot of country average exchange rate and average liquidity premium

Notes: An increase in the value of exchange rate is a depreciation of the header currency relative to the average of all other G10 currencies. An increase in treasury liquidity is an increase of the treasury liquidity of the header currency relative to the average of all other G10 currencies.

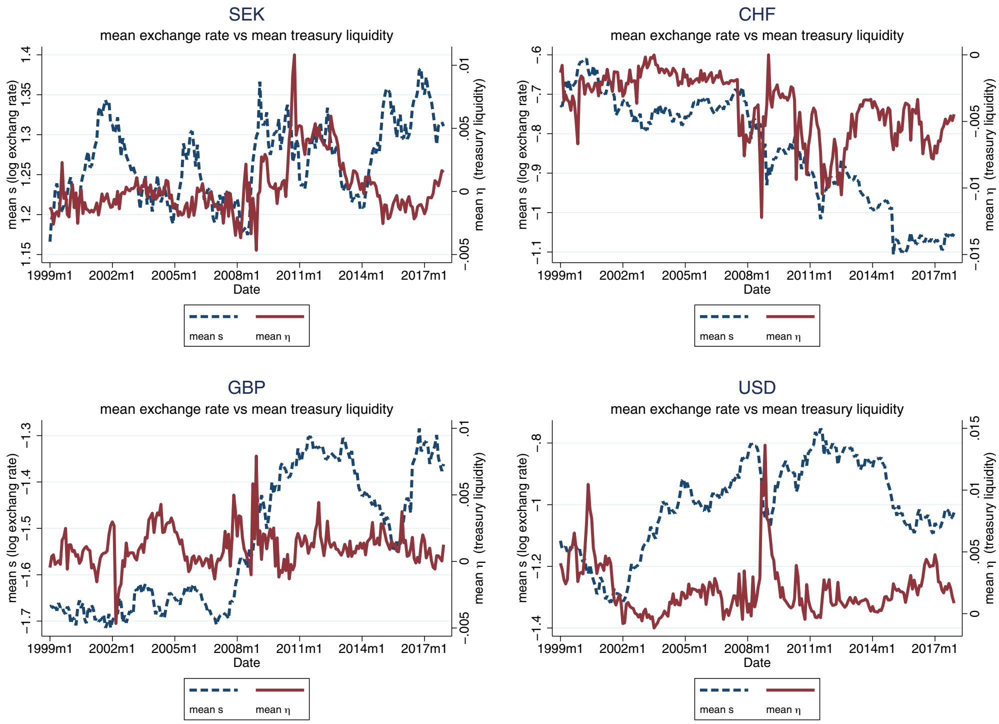
where s_t is the log nominal exchange rate, q_t is the log real exchange rate, η_t is a measure of liquidity, $i_t^R = i_t - i_t^*$ is the home minus foreign government bond interest rate differential, $\Delta X_t = X_t - X_{t-1}$ for any variable X.

As will become clear, our results—qualitatively, quantitatively, and by statistical significance—are essentially the same for both measures of η_t.[25] We first present the results

25. The difference between the two measures, $\tau_t = f_{t,t+1} - s_t + IRS_t^* - IRS_t$, is a measure of deviations from covered interest rate parity for interest rate swaps. In Section 3.3, we explore the relationship between the covered interest parity deviation *per se* and exchange-rate changes.

FIGURE 1

(Continued)



using the measure of η_t given in equation (14) that uses LIBOR swap rates, but we present many of our detailed results and robustness tests using the measure given in (15) for three reasons. First, Baba and Packer (2009) note that financial institutions that prefer a short position in one currency and long in another find it cheaper to use a foreign currency swap (earning $f_{t,t+1}-s_t$) rather than taking a long deposit in one currency and borrowing in the other (earning $IRS_t - IRS_t^*$.) Second, we use definition (15) in order for our results to be most directly related to those in other recent studies. In the case where the US is assumed to be the home country, Du *et al.* (2018a) denotes the $\tilde{\eta}_t$ term here as the US Treasury Premium, $\Phi_{j,n,t}$, which is the n-year deviation from covered interest parity between government bond yields in the United States and country j. Jiang *et al.* (2021) take the US as the home country and define $-\tilde{\eta}_t$ as a cross-country average over nine large markets relative to the dollar. Third, for some currencies, there are longer samples using the measure from (15) rather than from (14). However, the first two sets of results we report demonstrate that the conclusions do not depend on the measure, either qualitatively, or to a large extent, quantitatively.

Table 1 reports the regression coefficient estimates of (16), using $\hat{\eta}_t$ from equation (14) as the measure of η_t.[26] Each row of the table represents the estimation results that take the country of the currency in the first column as the home country and rest of the nine countries as the foreign

26. To keep the table visibly clear, we only report the main coefficient estimates of interest and refer readers to Supplementary Appendix A5 for the full regression tables.

TABLE 1

Estimation result of baseline panel regression using interest rate swaps to construct liquidity measure

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\hat{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \hat{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	−0.028***	−5.84***	−5.64***	2028	0.19
	(0.007)	(0.78)	(0.53)		
CAD	−0.029***	−3.68***	−6.22***	1836	0.18
	(0.007)	(0.74)	(0.54)		
EUR	−0.021***	−3.52***	−5.03***	2028	0.12
	(0.006)	(0.58)	(0.42)		
JPY	−0.038***	−4.20***	−6.19***	2028	0.14
	(0.010)	(1.02)	(0.76)		
NZD	−0.029***	−6.06***	−5.88***	2028	0.17
	(0.009)	(0.81)	(0.63)		
NOK	−0.018***	−2.89***	−4.4***	2028	0.12
	(0.007)	(0.65)	(0.50)		
SEK	−0.024***	−4.20***	−4.46***	2028	0.11
	(0.006)	(0.64)	(0.48)		
CHF	−0.010*	−3.13***	−3.08***	2028	0.05
	(0.006)	(0.78)	(0.56)		
GBP	−0.020***	−4.09***	−5.35***	2028	0.13
	(0.007)	(0.74)	(0.52)		
USD	−0.014*	−6.04***	−4.74***	2028	0.13
	(0.007)	(0.84)	(0.60)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the first column currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\hat{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

countries. When constructing the variables, we use 1-year swap rates and 1-year government yields.[27] The real exchange rates are constructed using consumer price levels.

First, consistent with our theoretical prediction and the empirical results of Eichenbaum *et al.* (2021), the coefficient estimates for $q_{j,t-1}$ are all negatively significant, implying that real exchange rates adjust through nominal exchange rates. The average coefficient estimate is approximately −0.023, implying a 2.3% adjustment of the nominal exchange rate in the direction of the long-run real exchange rate, per month. It is interesting to note that the estimated adjustment of the dollar exchange rate is around half the size of the average (across currencies) adjustment coefficient, suggesting a more persistent real exchange rate.

Second, we find that a positive change in the relative interest rate (home minus foreign) drives a contemporaneous home currency appreciation, which matches the traditional interest rate and exchange rate relationship. While almost all monetary, sticky-price models of exchange rates predict such a relationship, empirical support for even a contemporaneous relationship between interest rates and exchange rates has not been universally strong in previous studies.[28] It may be

27. See the discussion and robustness tests below for the choice of 1-year tenor. Supplementary Appendix A2 shows that for our countries, the average correlation rate between the policy rate and the 1-year bond rate that we use is 0.967.

28. See Engel (2014) for a recent survey.

TABLE 2

Estimation result of baseline panel regression with baseline liquidity measure

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\tilde{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	−0.028***	−5.27***	−5.74***	2052	0.19
	(0.007)	(0.72)	(0.54)		
CAD	−0.027***	−4.61***	−5.46***	2052	0.17
	(0.006)	(0.62)	(0.49)		
EUR	−0.02***	−4.64***	−5.02***	2052	0.14
	(0.006)	(0.52)	(0.41)		
JPY	−0.04***	−4.39***	−6.32***	2052	0.17
	(0.010)	(0.95)	(0.74)		
NZD	−0.028***	−6.29***	−6.02***	2052	0.2
	(0.008)	(0.73)	(0.61)		
NOK	−0.019***	−4.01***	−4.87***	2052	0.15
	(0.007)	(0.61)	(0.49)		
SEK	−0.023***	−4.52***	−4.60***	2052	0.13
	(0.006)	(0.58)	(0.46)		
CHF	−0.013**	−2.32***	−2.76***	2052	0.05
	(0.007)	(0.71)	(0.56)		
GBP	−0.023***	−3.35***	−5.24***	2052	0.13
	(0.007)	(0.67)	(0.52)		
USD	−0.011*	−6.44***	−4.77***	2052	0.17
	(0.007)	(0.72)	(0.57)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

that the importance of the interest rate channel requires controlling for the error-correction term and liquidity yields, as in our specification. We find the interest rate effect is strongly statistically significant for all 10 currencies. The average coefficients, across the currencies, is −5.10, which means that a 100 basis point increase in the annualized interest rate in the home currency relative to the foreign country leads on average to a 5.10% appreciation from the previous month.

Our main novel results concern the effects of the liquidity yield on exchange rates. The coefficient estimates for $\Delta\tilde{\eta}_{j,t}$ are all negative and statistically significant at the 1% level, with a range from −2.89 to −6.06. This indicates a 2.89% to 6.06% home currency appreciation in a month when there is a positive change of 100 basis points (annualized rate) in relative liquidity. We find that the relative government bond liquidity exhibits a very strong relationship with exchange rate movements for all the G-10 countries.

Table 2 displays the findings from estimation of the model using the $\tilde{\eta}_t$ measure of the relative convenience yield as defined in equation (15). The results are almost identical to those of the baseline estimation in Table 1, qualitatively, quantitatively, in precision of parameter estimates and in overall fit of the equation.

Tables 1 and 2 point to two important aspects of the impact of the liquidity yield. First, it is not just a US dollar phenomenon. While a great deal of attention has been paid to the convenience yield on US government bonds, our regression results show that the relative liquidity yield is an important factor in explaining exchange rate changes for all of the G10 currencies. Further results reported in Table 6 (discussed below), and in Section 3.4 emphasize this point.

However, secondly, the US is still a special case in the sense that the impact of the relative convenience yields on the exchange rate is the largest in the US. The estimated coefficient on the liquidity yield is largest in absolute value for the US (along with New Zealand), and the size of that impact is substantially larger than for most other currencies.

As a check on the reasonableness of the coefficient estimates in Table 1, we can use the model for the estimating equation, (13), and the price adjustment equation, (6), to derive estimates of the structural parameters, $\theta, \sigma, \psi, \alpha, \rho$, and δ. We use the average first-order serial correlation of the real exchange rates for the G10 currencies to estimate θ, and then the average estimates of $\beta_i, i=1,\ldots,5$ to estimate the other five parameters. Supplementary Appendix B4 describes the procedure in detail.

We find an unbiased estimate of θ is 0.017, which gives the average half-life of the G10 real exchange rates as 40 months, consistent with Rogoff's (1996) estimate that the half-life of real exchange rates of major currencies is 3–5 years. We find the responsiveness of interest rates to inflation in the monetary policy rule, σ, of 4.69. While this is large compared to standard estimates, our monetary policy rule does not include an interest-rate smoothing term but instead incorporates persistence in the interest rate through serial correlation in the shocks to the policy rule. We estimate that serial correlation, δ, as 0.964. The estimated value of $(1-\delta)\sigma$ is 0.169, which is in line with estimates of the inflation responsiveness of interest rates to inflation in monetary rules that include interest-rate smoothing.

We estimate α from equation (5), which captures the relation between the interest rate differential and the relative liquidity yield—the Nagel (2016) effect—to be 0.211. The measure of the reaction of interest rates to the liquidity yield in the monetary policy rule, ψ, from equation (7), is 0.241. Our estimates of α and ψ are less than one, as in the model, and modest in size but not negligible. Finally, we find the serial correlation of the liquidity yield shocks, ρ, is 0.267.

All the structural parameter estimates are plausible, giving us some more confidence in our interpretation of the exchange-rate regression.

These parameters imply from the model equation (13) that the effects of $\Delta\eta_{j,t}$ and $\Delta i^R_{j,t}$ on the change in the log of the exchange rate will be similar. That is, the regression coefficients from the estimating equation (16) will be alike, and indeed in Table 1, the average value (across the 10 currencies) of β_2 is -4.35 and of β_3 is -5.10. If those coefficient estimates were identical, then we could consolidate the effects of $\Delta\eta_{j,t}$ and $\Delta i^R_{j,t}$ on the exchange rate into one variable, $\Delta\eta_{j,t}+\Delta i^R_{j,t}=\Delta\left(IRS_t-IRS^*_t\right)$. The currency value in this case appears to be driven only by the market interest rate differential, and it might be tempting to simply add together the liquidity yield and the interest payments as two components of the return on investments in a country. But our results show that would be a mistaken approach because the strong statistical significance of both $\Delta\eta_{j,t}$ and $\Delta i^R_{j,t}$ tells us that there are independent forces driving the relative liquidity yield and government interest rate differentials that impact the exchange rate. The relationship between the market interest differential and exchange rates can be decomposed into separate effects, one driven primarily by liquidity shocks, and the other primarily by monetary policy.[29]

29. In Supplementary Appendix A16, we report Table 1 regression but control for market interest rate rather than government bond interest rate.

TABLE 3

Estimation result of baseline panel regression without liquidity measure $\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta i^R_{j,t} + \beta_3 i^R_{j,t-1} + u_{j,t}$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	−0.032***	−4.07***	−0.28***	2052	0.11
	(0.007)	(0.51)	(0.11)		
CAD	−0.028***	−4.54***	−0.25**	2052	0.12
	(0.006)	(0.47)	(0.10)		
EUR	−0.023***	−3.83***	−0.20**	2052	0.09
	(0.006)	(0.39)	(0.09)		
JPY	−0.034***	−5.32***	−0.15	2052	0.10
	(0.011)	(0.73)	(0.13)		
NZD	−0.031***	−2.67***	−0.09	2052	0.06
	(0.009)	(0.60)	(0.13)		
NOK	−0.018**	−3.50***	−0.13	2052	0.08
	(0.007)	(0.47)	(0.10)		
SEK	−0.027***	−3.21***	−0.14	2052	0.07
	(0.006)	(0.44)	(0.10)		
CHF	−0.010	−2.00***	−0.22**	2052	0.02
	(0.006)	(0.52)	(0.10)		
GBP	−0.023***	−3.72***	−0.33***	2052	0.09
	(0.007)	(0.49)	(0.10)		
USD	−0.014*	−3.68***	−0.14	2052	0.07
	(0.008)	(0.58)	(0.11)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

3.2.1. Omitting the liquidity return. For comparison, we also conduct the regression (16) but excluding the liquidity yield variables. That is:

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2(\Delta i^R_{j,t}) + \beta_3(i^R_{j,t-1}) + u_{j,t}. \tag{17}$$

The regression estimates are reported in Table 3. The coefficient estimates on lagged real exchange rates and change in interest rate differential remain negatively significant for all country pairs. However, the within R^2 for this specification are universally much lower compared to Tables 1 or 2.[30] This indicates including relative government bond liquidity returns brings strong explanatory power to exchange rate determination, in addition to and independent of the traditional factors.

3.2.2. Estimation on sub-samples. Next, we investigate whether the relationship between government bond liquidity and exchange rates are driven by (1) the Global Financial

30. The average R^2 in the baseline regression in Table 2, which uses the same measure of the liquidity yield as Table 3, is 0.150, but only 0.081 in the regressions that omit the liquidity yield.

TABLE 4

Estimation result of baseline panel regression, pre- and post-2008

$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\tilde{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$

Home Currency (1)	$\Delta\tilde{\eta}_{j,t}$ (2)	Within R^2 (3)	$\Delta\tilde{\eta}_{j,t}$ (4)	Within R^2 (5)
	1999M1–2007M12		2008M1–2018M1	
AUD	−3.78***	0.086	−6.03***	0.296
	(1.21)		(0.88)	
CAD	−2.71**	0.090	−5.73***	0.292
	(1.09)		(0.73)	
EUR	−2.89***	0.046	−5.30***	0.259
	(0.85)		(0.65)	
JPY	−1.17	0.041	−5.73***	0.33
	(1.32)		(1.24)	
NZD	−4.47***	0.099	−6.92***	0.321
	(1.12)		(0.95)	
NOK	−3.58***	0.089	−4.88***	0.258
	(1.00)		(0.77)	
SEK	−2.98***	0.074	−5.60***	0.228
	(0.91)		(0.73)	
CHF	−1.13	0.026	−2.86***	0.092
	(1.01)		(1.01)	
GBP	−4.10***	0.099	−3.42***	0.209
	(0.88)		(0.92)	
USD	−3.98***	0.079	−7.11***	0.326
	(1.11)		(0.86)	

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP) and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country$j$, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2007M12 and 2008M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

Crisis, (2) the post-crisis period, or (3) only by the US dollar. In Table 4, we re-estimate the model but split the sample period into two periods, pre-2008 and post (and including)-2008. We see that the contemporaneous relationship between the change of the liquidity measure and the change of exchange rates holds in both time periods. As in the full sample, all the estimated coefficients on the impact of the estimated government liquidity return are negative. They are all individually statistically significant at the 1% level in the post-crisis period. In the pre-crisis data, the p-values for these coefficients are all less than 0.01 except for Japan and Switzerland but both of them still have a negative coefficient. The coefficient estimates in all cases have larger values in absolute terms after 2008, ranging from −2.86 to −7.11. In addition to the significant and larger coefficients, the post-2008 R^2 are markedly improved, with a maximum of 33%, reflecting the importance of the relationship between the government bond liquidity and exchange rate determination.[31]

31. In a country by country estimation of (16) reported in Online Appendix A6, the maximum R^2 is 49%, which is the AUD – JPY pair.

TABLE 5

Estimation result of baseline panel regression, excluding 2007–2009

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\tilde{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	−0.023***	−5.02***	−4.90***	1728	0.128
	(0.007)	(0.92)	(0.62)		
CAD	−0.023***	−4.72***	−4.63***	1728	0.127
	(0.006)	(0.77)	(0.55)		
EUR	−0.017***	−5.15***	−3.75***	1728	0.090
	(0.006)	(0.70)	(0.49)		
JPY	−0.035***	−2.21*	−3.46***	1728	0.058
	(0.010)	(1.21)	(0.88)		
NZD	−0.023***	−5.85***	−4.49***	1728	0.109
	(0.008)	(0.95)	(0.72)		
NOK	−0.015**	−4.19***	−4.14***	1728	0.107
	(0.007)	(0.85)	(0.59)		
SEK	−0.018***	−4.18***	−3.38***	1728	0.078
	(0.006)	(0.77)	(0.55)		
CHF	−0.008	−2.19**	−0.90	1728	0.021
	(0.006)	(0.88)	(0.63)		
GBP	−0.016**	−4.75***	−4.00***	1728	0.083
	(0.006)	(0.86)	(0.60)		
USD	−0.008	−7.11***	−4.32***	1728	0.125
	(0.007)	(0.96)	(0.67)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other 9 currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1, but excludes data from 2007M1–2009M12. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

It is not the case that the results are driven by the global financial crisis years, 2007–2009. Table 5 reestimates the model using the full sample but excluding those crisis years. Compared to our baseline results, the findings excluding the crisis years are nearly identical.[32] The estimated coefficients on the change in $\hat{\eta}_t$ and the change in $i_t - i^*_t$ are very similar in magnitude to those we report in Table 1, and all are statistically significant as in our baseline estimates. The error-correction terms for the adjustment of the nominal exchange rate to lagged real exchange rates is marginally statistically insignificant for Switzerland and the US in this sample, and the overall fit of the model is not as good as indicated by generally lower values of the R^2 statistics. But the message of this table is that the principal findings are clearly not determined solely by the crisis years.

Table 6 displays estimates of the model that exclude the US dollar from the sample. Each of the panel regressions is left with eight foreign currencies. We find that the relationship between government bond liquidity and exchange rates is largely unchanged. The coefficients on the

32. Online Appendix A10 reports regressions that excludes 2000-2001 (dot com bubble and 9/11), 2007-2013 (GFC and European debt crisis). The main findings are robust to this.

TABLE 6

Estimation result of baseline panel regression, excluding the US dollar

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\tilde{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	-0.032^{***}	-4.80^{***}	-5.64^{***}	1824	0.184
	(0.008)	(0.72)	(0.53)		
CAD	-0.033^{***}	-4.31^{***}	-5.42^{***}	1824	0.178
	(0.007)	(0.65)	(0.50)		
EUR	-0.026^{***}	-4.32^{***}	-4.95^{***}	1824	0.141
	(0.006)	(0.53)	(0.41)		
JPY	-0.044^{***}	-4.78^{***}	-6.64^{***}	1824	0.18
	(0.011)	(1.01)	(0.77)		
NZD	-0.033^{***}	-5.95^{***}	-5.97^{***}	1824	0.194
	(0.009)	(0.72)	(0.60)		
NOK	-0.023^{***}	-3.49^{***}	-4.93^{***}	1824	0.153
	(0.007)	(0.60)	(0.48)		
SEK	-0.029^{***}	-4.02^{***}	-4.52^{***}	1824	0.125
	(0.006)	(0.57)	(0.45)		
CHF	-0.015^{**}	-1.88^{***}	-2.71^{***}	1824	0.049
	(0.007)	(0.71)	(0.55)		
GBP	-0.024^{***}	-3.43^{***}	-5.40^{***}	1824	0.136
	(0.007)	(0.71)	(0.55)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The nine currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), and British Pound (GBP). We exclude United States Dollar (USD) from all the regressions. Each row represents a regression estimation using the column (1) currency as the home currency and the other 8 currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

liquidity measure are negative and significant, though generally slightly smaller in absolute term than those estimated in Table 1 (except for JPY and GBP). The R^2 statistics are also largely unchanged. This exercise shows that our results are not simply a US dollar phenomenon, instead observed in all currency pairs.[33] These findings do not preclude the view that the US dollar is special, or that its convenience yield is central in the global economy, as in Bianchi *et al.* (2021), and Jiang, Krishnamurthy and Lustig (2020), and Jiang *et al.* (2021). Apropos Verdelhan (2018), the US convenience yield may be a significant factor driving exchange rates, and in our setting, the relationship between relative convenience yield of two non-US countries, i and j, on the ij exchange rate might reflect the relative liquidity value of each country's bonds relative to the US Treasury bonds. In other words, in a counterfactual world in which the US is absent, country i's bonds might be more liquid than country j's, but in the actual world, country j's might have a

33. Table 12 and Supplementary Appendix A6 summarize and report the regression results country by country. Jiang *et al.* (2021) find a weaker relationship between the relative liquidity yield and non-US exchange rates when the US dollar is omitted from the study. We find that the main reasons for the differences in our findings are that we use monthly data rather than the quarterly data in Jiang *et al.*, and that we estimate by panel methods, while Jiang *et al.* use univariate regressions of one currency relative to an average of the others. Supplementary Appendix A details the effects of these differences.

higher convenience yield because i's bonds are a close substitute for US bonds, while j's offer some independent liquidity return.

3.2.3. One-month forward rates. As we have noted, in our baseline regressions we use 1-year forward rates and 1-year government yields as regressors, while the regressions are conducted in monthly frequency. The choice of 1-year tenor is a tradeoff between model consistency and data availability. Ideally, for model consistency, we would use 1-month forward rates and government yields to construct the variables. However, the data availability of 1-month government yields is rather limited for some of the sample countries. In addition, in Section 3.3, we use credit default swap (CDS) data to make an adjustment for the probability of non-repayment of government debt. The CDS data are more extensively available only for tenors of one year or above. Therefore, we use 1-year forward rates and 1-year government yields to construct the variables in our analysis. To be fully consistent with the model, investors would need to have no uncertainty about the 1-month own-currency return on 1-year bonds, but the variation in that return (annualized) relative to the 1-year interest rate is very small relative to changes in exchange rate. The monthly correlation of 1-year and 1-month interest rates is over 0.90 in our sample for all countries.

Nevertheless, to make sure our result is robust to the choice of tenor, we report in Table 7 the regression coefficient estimates of equation (16), using 1-month forward rates and one-month government yield data.[34] The empirical relationship between the change of nominal exchange rate and the independent variables is largely consistent with the result we discussed in Table 2, which uses 1-year forward rates and 1-year government yields data. Considering this, to make our empirical results comparable across different specifications, we use 1-year forward rates and 1-year government yields throughout the analysis.

3.2.4. Using survey measures of expectations. While uncertainty is an important driver of the liquidity yield, the question arises whether our findings arise simply because our measure of this relative convenience yield is a proxy for other deviations from uncovered interest parity, such as default risk or foreign exchange risk. Below, we control for default risk using CDS premiums. Also, our regressions include the interest rate differential as a separate variable, so if it is a risk premium at work, it must be a part of the risk premium that is uncorrelated with the interest rate differential. Here, we undertake an exercise to see if there is an independent role for the relative liquidity premium.

Supplementary Appendix A3 includes some estimates that incorporate a direct measure of the risk premium using the difference between forward rates and a survey measure of expected future exchange rates. Several recent studies have made use of data on exchange rate expectations of foreign exchange traders to measure the risk premium.[35] Our data on expectations come from the Bloomberg exchange rate forecasts. Letting $s^e_{t,t+1}$ represent the time t expectation of the exchange rate at time $t+1$, the risk premium can be measured as:

$$rp_t = i^*_t + s^e_{t,t+1} - s_t - i_t.$$

34. Norway is excluded in this exercise as a home country and foreign country due to lack of Norway 1-month government yield data.

35. Bussière, Chinn, Ferrara and Heipertz (2019), Chinn and Frankel (2020), Kalemli-Özcan (2019), Kalemli-Özcan and Varela (2019), and Stavrakeva and Tang (2018, 2019).

TABLE 7
Estimation result of baseline panel regression, using 1-month rates
$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\tilde{\eta}_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$

Home Currency (1)	$q_{j,t-1}$ (2)	$\Delta\hat{\eta}_{j,t}$ (3)	$\Delta i^R_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	−0.054**	−7.23	−22.9	360	0.054
	(0.026)	(10.60)	(14.90)		
CAD	−0.03***	−12.73***	−29.13***	1228	0.056
	(0.010)	(4.64)	(7.28)		
EUR	−0.083***	−18.61***	−20.04**	609	0.087
	(0.017)	(7.47)	(9.03)		
JPY	−0.102***	−23.34***	−11.66	462	0.140
	(0.024)	(10.13)	(11.69)		
NZD	−0.033***	−18.71***	−28.59***	1228	0.063
	(0.010)	(4.85)	(6.98)		
SEK	−0.033***	−14.97***	−17.22***	1228	0.046
	(0.009)	(3.88)	(5.61)		
CHF	−0.057***	−10.89	9.75	731	0.063
	(0.014)	(8.91)	(8.52)		
GBP	−0.022*	−8.17*	−19.29***	1228	0.031
	(0.012)	(4.91)	(7.03)		
USD	−0.025***	−17.17***	−15.16**	1228	0.062
	(0.009)	(3.94)	(6.15)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The nine currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Norwegian Krone (NOK) is missing due to data availability. Each row represents a regression estimation using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. *, **, and *** for $q_{j,t}$ is based on critical values from distribution for Augmented Dickey Fuller test with a constant. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in in Supplementary Appendix A5.

The risk premium augmented regression is

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \left(\Delta\tilde{\eta}_{j,t}\right) + \beta_3 \left(\Delta i^R_{j,t}\right) + \beta_4 \tilde{\eta}_{j,t-1} + \beta_5 i^R_{j,t-1} + \beta_6 \left(\Delta rp_{j,t}\right) + \beta_7 rp_{j,t-1} + u_{j,t}$$

We find that the change in the foreign risk premium does imply a home appreciation as theory would predict and is significant for all ten currencies. Importantly, the relative liquidity yield remains highly statistically significant in 9 of the 10 these regressions.

It is often argued in the literature that certain “safe haven” currencies (such as the US Dollar, the Yen, or the Swiss Franc) are less risky because their currencies appreciate during times of global turmoil. That relationship, however, does not explain why these currencies get stronger during global downturns. In Farhi and Gabaix (2015), marginal utilities of consumption in the safe-haven countries rise less than in other countries during such times, in turn because productivity declines are smaller in these countries. Gabaix and Maggiori (2015) build a model in which countries that are net debtors suffer a depreciation during times of global financial disruption, as financial intermediaries wish to unload debt securities as financial constraints tighten. As the paper notes, this conclusion runs counter to the evidence for the US. In Gourinchas, Ray and Vayanos (2020), during periods of weakness, markets fear even greater downturns in which government debt of some countries will default, which leads markets to find shelter in safer currencies. Perhaps the liquidity return story can complement the risk premium story by providing one other reason why some currencies are safe havens—because their government assets are more liquid, and therefore during global recessions demand for the safe haven currencies increases as the demand for liquidity rises.

3.2.5. Emerging market currencies. Our model and our empirical analysis is designed for analysis of low-inflation (e.g. inflation-targeting) countries with low default risk, as in the countries of the G10 currencies. We modify the model to extend our empirical analysis to emerging market currencies, focusing on eighteen emerging market currencies that Du *et al*. (2018) dataset provides.[36] We report two specifications in Supplementary Appendix A4. The first one takes an emerging market currency as the home country and the rest of the 17 emerging market countries as the foreign countries. The second one takes the rest of the 17 emerging market countries and all G10 currencies as the foreign countries. To account for larger heterogeneity in emerging countries, we further control for inflation and default risk (using credit default swaps), and allow for different coefficients for home and foreign variables.[37] Our key finding of the importance of liquidity yield carries forward to the emerging market regressions. Coefficients of change of liquidity measure are estimated negative in 16 out of the 18 panels. Eleven and 13 of them are significantly negative at 5% level in the two specifications. We also note that the coefficient estimates are generally smaller (in absolute value) for emerging markets. For those estimated negatively, they range from −0.003 (TRY) to −0.88 (RUB). This indicates a weaker role of Treasury bonds as a liquid and safe instrument in emerging markets.

We have seen so far that the inclusion of the liquidity variable greatly increases the explanatory power of the model for all G10 currencies, and the model is not very sensitive to estimation over sub-samples of the time span of our data, or to alternative measures of the relative convenience yield. We next dig deeper into the data to get a better understanding of what drives these results.

3.3. *Decomposing the liquidity measure*

In this section, we decompose the relative liquidity measure to assess the impact of three components on exchange rate movements: deviations from covered interest parity for interest rate swaps, $\tau_t \equiv f_{t,t+1} - s_t + IRS_t^* - IRS_t$; the difference between home and foreign credit default swap premiums, $l_t^R = CDS_t - CDS_t^*$; and the relative liquidity yield, corrected for default swap premiums, $\lambda_t = IRS_t - i_t - \left(IRS_t^* - i_t^*\right) + l_t^R$.[38]

In relation to our previous measures, $\hat{\eta}_t$ from equation (14) is given by:

$$\hat{\eta}_t = \lambda_t - l_t^R, \tag{18}$$

and $\tilde{\eta}_t$ from equation (15) is given by:

$$\tilde{\eta}_t = \hat{\eta}_t + \tau_t = \lambda_t - l_t^R + \tau_t. \tag{19}$$

First, the government bond yields might incorporate expected default risk. A credit default swap (CDS) contract insures the buyer against credit events. In the case of sovereign default,

36. They are the more commonly traded currencies in the emerging market sample, including Brazilian Real (BRL), Chilean Peso (CLP), Chinese Yuan (CNY), Colombian Peso (COP), Hungarian Forint (HUF), Indonesian Rupiah (IDR), Israeli New Shekel (ILS), Indian Rupee (INR), Korean Won (KRW), Mexican Peso (MXN), Malaysian Ringgit (MYR), Peruvian Sol (PEN), Philippine Peso (PHP), Polish Zloty (PLN), Russian Ruble (RUB), South African Rand (ZAR), Thai Baht (THB), and Turkish Lira (TRY).

37. See Section 3.3 for a more detailed discussion on accounting for default risk in the liquidity measure.

38. Details of the full derivation of these expressions are available at Du *et al.* (2018a).

the CDS sellers make payments to the buyers to compensate for the loss in the credit event. Buyers of the CDS pay a premium to CDS sellers for the insurance. The return to a riskless home government bond—that is, a home government bond protected by insurance—is therefore $i_t - CDS_t$, where CDS_t is the CDS premium. To measure the relative convenience yield, we therefore adjust the interest rate on government bonds to get the true riskless return. Then, λ_t can be understood as the relative government bond convenience yield after adjusting for credit default risk.

The home minus foreign difference of CDS premium is a measure of the relative premium investors are willing to pay to avoid default, i.e. $l^R_{j,t} = CDS_t - CDS^*_{j,t}$. Della Corte, Sarno, Schmeling and Wagner (2022) examine the effects of sovereign default on exchange rates. When $l^R_{j,t} > 0$, investors are willing to pay more for protection against home default compared to foreign default. When $\Delta l^R_{j,t+1} > 0$, there is an increase in home default risk relative to foreign default risk which we posit is associated with an immediate home currency depreciation. From one perspective, an increase in $l^R_{j,t}$, holding the nominal government bond interest rate fixed, simply implies that the return on home bonds falls because of the increase in the cost of the CDS.[39] Alternatively, as Della Corte S. (2022) demonstrate, an increase in the CDS rate reflects an increase in default probability and in default risk, which leads to a depreciation of the currency as its sovereign bonds are both riskier and offer lower expected returns.

The third component is the deviation from covered interest parity. If covered interest parity held for market returns, we should find $f_{t,t+1} - s_t + IRS^*_t = IRS_t$, where IRS_t (IRS^*_t) refers to the home (foreign) return on LIBOR swaps. Baba, Packer and Nagano (2008), Baba and Packer (2009), and Griffoli and Ranaldo (2011) attribute the failure of covered interest arbitrage in the years immediately following the global financial crisis to both a liquidity and a default factor. In particular, there appeared to be profitable arbitrage opportunities that involved borrowing in dollars and making covered investments in foreign interest-earning assets. These papers provide evidence that investors were reluctant to take advantage of such opportunities both because of counterparty risk, and because there was a global demand for liquid dollar assets. Du *et al.* (2018b) find that in recent years, for some currencies (particularly, when the US dollar is the home currency), $IRS_t < f_{t,t+1} - s_t + IRS^*_t$, but financial institutions do not undertake the arbitrage that would result in riskless profits. In order to earn those profits, banks would need to go short in dollars, and purchase the foreign currency on the spot market and go long in foreign currency (which they sell forward). Such an arbitrage investment, while risk free, expands the size of the financial institutions' balance sheets, and may cause them to run afoul of regulatory constraints. Financial institutions that held home assets could sell those and acquire synthetic home assets, but they might be unwilling to do so if they value the home assets for non-pecuniary reasons. Hence, when home assets are especially valued, then τ_t will be high, and the home currency will be strong. The same relationship could arise if there were default risk on LIBOR rates, as might have been the case in 2008 during the global financial crisis. When foreign LIBOR is considered risky, τ_t is high, and the home currency is strong. We note that Cerutti, Obstfeld and Zhu (2021) associate the failure of covered interest parity for the US dollar with periods of a strong dollar.[40]

Jiang *et al.* (2020, 2021) interpret τ_t as a convenience yield on home (US) LIBOR. That is, $f_{t,t+1} - s_t + IRS^*_t$ denotes the return on a foreign LIBOR, swapped into dollars. If this return is greater than IRS_t, so that $\tau_t > 0$, the home LIBOR pays a lower return because of its liquidity relative to the liquidity of the foreign LIBOR that is swapped into dollars in the forward market.

39. See Supplementary Appendix B7 for a more detailed discussion of the model solution with default risk.

40. See Du *et al.* (2018b) investigate deviations from covered interest parity and Avdjiev *et al.* (2018) consider the relationship between the currency swap friction and the exchange rate.

So, conceptually, we have two different sorts of convenience yields. τ_t is the convenience yield of home LIBOR relative to foreign LIBOR, while $\hat{\eta}_t$ is the convenience yield on the home government bond compared to the market rate, relative to the analogous convenience yield in the foreign country.

In all cases, we use IRS and CDS data with 1-year tenor as the CDS data are extensively available only for tenors of 1-year or above.

With these decomposed components, we modify the baseline regression by putting each into the equation. Specifically,

$$\begin{aligned}\Delta s_{j,t} = & \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta\lambda_{j,t} + \beta_3 \Delta\tau_{j,t} + \beta_4 \Delta l^R_{j,t} + \beta_5 \Delta i^R_{j,t} \\ & + \beta_6 \lambda_{j,t-1} + \beta_7 \tau_{j,t-1} + \beta_8 l^R_{j,t-1} + \beta_9 i^R_{j,t-1} + u_{j,t}\end{aligned} \tag{20}$$

As discussed above, we expect to find a negative estimate of β_3, because a larger $\Delta\tau_{j,t}$ indicates an unwillingness to sell home assets to buy the foreign currency, which appreciates the home currency. The estimated β_4 should be positive, since a larger $\Delta l^R_{j,t}$ means there is a greater default risk for home government bonds. $\Delta\lambda_{j,t}$ is the residual measure of the change in the home relative to foreign liquidity yields, and for that we posit a negative value of β_2. As in our model, we should also find negative values for the estimates of β_1 and β_5.

We estimate the regression in two ways. First, since CDS data for many of the sample countries are only available after 2008, we start the sample from 2008M1 and estimate (20). Second, to make use of the full sample information and test whether the adjusted liquidity measure is important in explaining the change of exchange rates throughout the sample, we estimate the regression from 1999M1, but excluding the CDS variables (dropping $\Delta l^R_{j,t}$ and $l^R_{j,t-1}$).[41]

In Table 8, the coefficient estimates on $\Delta\lambda_{i,t}$, which represents the effect of changes in government bond liquidity after adjusting for credit risk and derivative market friction, are still significantly negative in all cases. The range of coefficient is from -3.04 to -8.91 for the left panel, indicating a monthly 3.04% to 8.91% immediate home currency appreciation when there is a monthly positive change of 100 basis points (annualized rate) in relative liquidity. These coefficients are also larger than the coefficients of $\Delta\eta_{i,t}$ estimated in Tables 1 or 2. These results reaffirm our baseline result that there is a strong linkage between government bond liquidity and exchange rates.

In many cases, we also see that credit risk variation and derivative market frictions are important variables in explaining the change of exchange rates.[42] The positive coefficient on $\Delta l^R_{j,t}$ indicates that an increase in home default risk relative to foreign default risk is associated with an immediate home currency depreciation. Holding the nominal government bond interest rate fixed, an increase in default risk implies the default risk adjusted nominal interest rate goes down, resulting in a home currency depreciation. The negative coefficient on $\Delta\tau_{j,t}$ can be interpreted as the influence of an increase in the convenience yield on home bonds relative to foreign bonds. The channel could go through default risk on interest rate swaps, or there may be a liquidity yield of the home currency asset.

To confirm our results are robust to different specifications, we conduct the estimation in (20) by including one or two sub-components at a time. The results are reported at Table 9. Once again, we find the regression coefficients for $\Delta\lambda_{j,t}$ are significantly negative in all cases.

41. In the second case, the $\lambda_{j,t}$ is effectively $\eta_{j,t} - \tau_{j,t}$

42. See Della Corte S. (2022) who find similar findings of the relationship between exchange rate and sovereign risk.

TABLE 8

Estimation result of panel regression with decomposed liquidity measure

$\Delta s_{j,t}=\alpha_j+\beta_1 q_{j,t-1}+\beta_2\Delta\lambda_{j,t}+\beta_3\Delta\tau_{j,t}+\beta_4\Delta l^R_{j,t}+\beta_5\Delta i^R_{j,t}+\beta_6\lambda_{j,t-1}+\beta_7\tau_{j,t-1}+\beta_8 l^R_{j,t-1}+\beta_9 i^R_{j,t-1}+u_{j,t}$

Home Currency (1)	$\Delta\lambda_{j,t}$ (2)	$\Delta\tau_{j,t}$ (3)	Within R^2 (4)	$\Delta\lambda_{j,t}$ (5)	$\Delta\tau_{j,t}$ (6)	$\Delta l^R_{j,t}$ (7)	Within R^2 (8)
	Full sample, no default risk			Post 2008, with default risk			
AUD	−6.15***	−2.98**	0.200	−7.00***	−3.11**	14.37***	0.288
	(0.79)	(1.23)		(1.12)	(1.54)	(2.36)	
CAD	−4.6***	−5.20***	0.207	−8.86***	−6.89***	8.32***	0.305
	(0.72)	(1.13)		(1.52)	(1.80)	(2.57)	
EUR	−4.66***	−4.91***	0.147	−6.10***	−3.92**	8.21***	0.259
	(0.57)	(0.87)		(0.80)	(0.95)	(1.68)	
JPY	−4.16***	−4.99***	0.173	−6.79***	−4.43***	10.2***	0.339
	(1.00)	(1.59)		(1.38)	(1.85)	(3.17)	
NZD	−6.62***	−5.77***	0.201	−7.79***	−5.84***	12.25***	0.314
	(0.80)	(1.30)		(1.14)	(1.48)	(2.51)	
NOK	−3.84***	−5.08***	0.157	−5.11***	−5.74***	4.08**	0.26
	(0.65)	(1.04)		(0.81)	(1.20)	(1.96)	
SEK	−4.46***	−5.02***	0.133	−5.56***	−4.18***	7.43***	0.205
	(0.64)	(0.98)		(0.89)	(1.15)	(1.90)	
CHF	−3.04***	−1.17	0.055	−3.24**	−1.2	5.60**	0.031
	(0.77)	(1.20)		(1.49)	(1.81)	(2.61)	
GBP	−4.19***	−1.4	0.137	−6.13***	−0.45	5.61**	0.229
	(0.75)	(1.12)		(1.12)	(1.41)	(2.35)	
USD	−6.32***	−6.74***	0.171	−8.91***	−3.20**	12.56***	0.376
	(0.82)	(1.19)		(1.11)	(1.25)	(2.16)	

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP) and United States Dollar (USD). Each row represents a regression estimation using the column (1) currency as the home currency and the other 9 currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tau_{j,t}$ is the measure of currency derivative friction, $l^R_{j,t}$ is the measure of home minus foreign default risk, $\lambda_{j,t}$ is the measure of the government bond liquidity after adjusting for derivative market friction and default risk, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1 (Columns (2) to (4)) and 2008M1–2018M1 (Columns (5) to (8)). Germany government interest rate and default risk are used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

How much of the variation of $\tilde{\eta}_t$ is driven by each of the sub-components? We can answer this with a variance decomposition. Table 10 reports the decomposition given by:

$$1=\frac{var(\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}+\frac{var(\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}+\frac{var(\Delta l^R_t)}{var(\Delta\tilde{\eta}_t)}+2\frac{cov(\Delta\lambda_t,\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}-2\frac{cov(\Delta\tau_t,\Delta l^R_t)}{var(\Delta\tilde{\eta}_t)}-2\frac{cov(\Delta l^R_t,\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}. \tag{21}$$

For most of the countries, the variation of $\Delta\lambda_t$ contributes a large share of variation of $\Delta\tilde{\eta}_t$. However, the sums of the variance shares of $\Delta\lambda_t$, $\Delta\tau_t$, and Δl^R_t are greater than one. This arises because of the negative correlation of $\Delta\lambda_t$ with $\Delta\tau_t$ and $-\Delta l^R_t$. We see that even controlling for default and swap-market frictions, the liquidity yield is still a significant determinant of exchange rates.

TABLE 9

Estimation result of panel regression with decomposed liquidity measure one by one

$$\Delta s_{j,t} = \alpha_j + \beta_1 q_{j,t-1} + \beta_2 \Delta X_{j,t} + \beta_3 \Delta i^R_{j,t} + \beta_4 X_{j,t-1} + \beta_5 i^R_{j,t-1} + u_{j,t}$$

where $X_{j,t}$ is the column head variable

Home currency (1)	$\Delta\lambda_{j,t}$ (2)	$\Delta\tau_{j,t}$ (3)	$\Delta l^R_{j,t}$ (4)	$\Delta(\eta + l^R)_{j,t}$ (5)	$\Delta(\eta - \tau)_{j,t}$ (6)	$\Delta(\tau - l^R)_{j,t}$ (7)
AUD	−4.33***	−1.42	9.95***	−4.25***	−5.84***	−4.65***
	(1.09)	(1.28)	(2.22)	(0.96)	(0.78)	(1.31)
CAD	−3.63***	−4.36***	0.75	−5.29***	−3.68***	−1.82
	(1.33)	(1.12)	(2.27)	(1.23)	(0.74)	(1.44)
EUR	−3.12***	−3.16***	2.08	−3.76***	−3.52***	−2.39***
	(0.74)	(0.87)	(1.55)	(0.62)	(0.58)	(0.81)
JPY	−4.90***	−5.81***	4.57	−4.52***	−4.20***	−6.16***
	(1.36)	(1.64)	(3.10)	(1.14)	(1.02)	(1.64)
NZD	−4.93***	−4.06***	4.54*	−5.19***	−6.06***	−5.21***
	(1.13)	(1.41)	(2.58)	(0.91)	(0.81)	(1.37)
NOK	−4.14***	−3.02***	−1.21	−4.73***	−2.89***	−3.38***
	(0.77)	(1.06)	(1.95)	(0.67)	(0.65)	(1.08)
SEK	−4.04***	−4.37***	3.08*	−4.17***	−4.20***	−3.71***
	(0.82)	(0.10)	(1.80)	(0.68)	(0.64)	(0.10)
CHF	−1.79	−1.48	2.45	−1.54	−3.13***	−2.06
	(1.32)	(1.20)	(2.29)	(1.07)	(0.78)	(1.45)
GBP	−5.04***	−0.67	0.78	−3.36***	−4.09***	−0.45
	(1.03)	(1.14)	(2.23)	(0.87)	(0.74)	(1.20)
USD	−5.13***	−6.13***	5.14**	−4.57***	−6.04***	−4.68***
	(1.11)	(1.23)	(2.11)	(0.85)	(0.84)	(1.16)
G10 average within R^2	0.204	0.099	0.169	0.222	0.134	0.181

Notes: The table reports the OLS estimates of the coefficient of change of liquidity measure of the regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tau_{j,t}$ is the measure of currency derivative friction, $l^R_{j,t}$ is the measure of home minus foreign default risk, $\lambda_{j,t}$is the measure of the government bond liquidity after adjusting for derivative market friction and default risk, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Regressions involving default risk $l^R_{j,t}$ are only estimated through 2008M1–2018M1 period (Columns (2), (4), (5), (7)). Germany default risk is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

3.4. *Country-specific government bond liquidity*

So far, we have conducted all our analysis with different measures of bilateral relative government bond liquidity. However, the impact of the own-country liquidity service and the aggregate foreign country liquidity service might have different effects on the home exchange rate.

We measure the home and foreign liquidity returns on government bonds as $\gamma_t = IRS_t - i_t$ and $\gamma^*_{j,t} = IRS^*_t - i^*_t$. Motivated by the decomposition above, we will include also the currency derivative market friction, $\tau_{j,t}$. We have then decomposed $\tilde{\eta}_{j,t}$ used in our baseline regressions:

$$\tilde{\eta}_{j,t} = \tau_{j,t} + \gamma_t - \gamma^*_{j,t}. \tag{22}$$

TABLE 10

Variance share of component of liquidity measure:

$$1=\frac{var(\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}+\frac{var(\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}+\frac{var(\Delta l_t^R)}{var(\Delta\tilde{\eta}_t)}+2\frac{cov(\Delta\lambda_t,\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}-2\frac{cov(\Delta\tau_t,\Delta l_t^R)}{var(\Delta\tilde{\eta}_t)}-2\frac{cov(\Delta l_t^R,\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}$$

Home currency	$\frac{var(\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}$ (%)	$\frac{var(\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}$ (%)	$\frac{var(\Delta l_t^R)}{var(\Delta\tilde{\eta}_t)}$ (%)	$\frac{2cov(\Delta\lambda_t,\Delta\tau_t)}{var(\Delta\tilde{\eta}_t)}$ (%)	$\frac{2cov(\Delta\tau_t,\Delta l_t^R)}{var(\Delta\tilde{\eta}_t)}$ (%)	$\frac{2cov(\Delta l_t^R,\Delta\lambda_t)}{var(\Delta\tilde{\eta}_t)}$ (%)
(1)	(2)	(3)	(4)	(5)	(6)	(7)
AUD	71	36	13	−7	−4	19
CAD	140	55	56	−46	−6	110
EUR	120	55	23	−50	−8	54
JPY	63	37	15	12	1	27
NZD	80	39	12	−5	1	26
NOK	98	25	10	−10	2	18
SEK	87	26	14	0	−3	31
CHF	67	41	18	12	1	38
GBP	81	36	16	−6	1	26
USD	73	39	25	4	0	41

The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP) and United States Dollar (USD). Each row represents the variance and covariance using the column (1) currency as the home currency and the other 9 currencies as foreign currency j. $\tilde{\eta}_t$ is the measure of government bond liquidity, $\tau_{j,t}$ is the measure of currency derivative friction, $l_{j,t}^R$ is the measure of home minus foreign default risk, $\lambda_{j,t}$ is the measure of the government bond liquidity after adjusting for derivative market friction and default risk, $i_{j,t}^R$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 2008M1–2018M1. Germany default risk and government interest rate are used for EUR case.

We estimate the following equation with the country-specific liquidity measures, controlling for the derivative market friction:

$$\Delta s_{j,t}=\alpha_j+\beta_1 q_{j,t-1}+\beta_2\Delta\tau_{j,t}+\beta_3\Delta\gamma_t+\beta_4\Delta\gamma_{j,t}^*+\beta_5\Delta i_{j,t}^R+\beta_6\tau_{t-1}$$
$$+\beta_7\gamma_{t-1}+\beta_8\gamma_{j,t-1}^*+\beta_9 i_{j,t-1}^R+u_{j,t}. \tag{23}$$

Estimates of β_3 and β_4 in (23) show how the change of country-specific government bond liquidity affects exchange rate movements. We expect a negative sign for β_3 and a positive sign for β_4.

Table 11 presents the estimation results for the country-specific government bond liquidity. The second column gives the coefficient estimates for the change in government bond liquidity for all the foreign currencies. The coefficient estimates are all significantly positive, indicating an increase in government bond liquidity of the foreign country is associated with a depreciation of the home currency, which is consistent with our theory and the empirical finding above. All the coefficient estimates of the home government bond liquidity, $\Delta\gamma_t$, term are significantly negative with the exception of the Japanese Yen and Swiss Franc. Both estimates are negative but with smaller absolute size compared to others.

These results then show that our findings regarding the effect of the relative liquidity returns on exchange rates are, for each country, driven at least in part by the liquidity return of that country. That is, the effects on exchange rates of the relative liquidity returns are not all determined by liquidity returns in one or a few larger countries.

We provide further evidence that our findings are not driven by one or a few countries by performing our baseline regression (16) country-by-country. Table 12 provides a summary of those regression results. (We report all 45 country-by-country regressions in Supplementary Appendix A6.) For the analysis that uses the entire 1999–2018 sample, among the 45 country pairs, 37 country pairs have coefficients on $\Delta\tilde{\eta}_t$ that are negatively statistically different

TABLE 11

Estimation result of panel regression with country-specific liquidity measure

$$\Delta s_{j,t}=\alpha_j+\beta_1 q_{j,t-1}+\beta_2\Delta\tau_{j,t}+\beta_3\Delta\gamma_t+\beta_4\Delta\gamma^*_{j,t}+\beta_5\Delta i^R_{j,t}+\beta_6\tau_{t-1}+\beta_7\gamma_{t-1}+\beta_8\gamma^*_{j,t-1}+\beta_9 i^R_{j,t-1}+u_{j,t}$$

Home Currency (1)	$\Delta\gamma^*_{j,t}$ (2)	$\Delta\gamma_{j,t}$ (3)	$\Delta\tau_{j,t}$ (4)	Observations (5)	Within R^2 (6)
AUD	5.48***	−6.76***	−2.74***	2028	0.202
	(0.69)	(1.21)	(1.21)		
CAD	4.42***	−5.91***	−5.23***	1836	0.208
	(0.69)	(1.98)	(1.13)		
EUR	4.58***	−4.96***	−5.08***	2028	0.147
	(0.56)	(1.09)	(0.91)		
JPY	4.27***	−2.09	−4.82***	2028	0.174
	(0.99)	(5.17)	(1.58)		
NZD	6.11***	−6.78***	−5.55***	2028	0.208
	(0.86)	(0.90)	(1.29)		
NOK	5.59***	−3.31***	−5.02***	2028	0.164
	(0.73)	(0.76)	(1.04)		
SEK	4.79***	−4.04***	−5.03***	2028	0.134
	(0.63)	(1.22)	(0.99)		
CHF	3.09***	−2.73	−1.22	2028	0.056
	(0.75)	(1.89)	(1.19)		
GBP	4.75***	−3.68***	−1.17	2028	0.139
	(0.70)	(1.08)	(1.11)		
USD	6.22***	−5.84***	−6.47***	2028	0.188
	(0.80)	(1.37)	(1.20)		

Notes: The table reports the OLS estimates of the coefficient of the panel fixed effect regression listed above. The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a regression estimation using the Column (1) currency as the home currency and the other 9 currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tau_{j,t}$ is the measure of currency derivative friction, γ'' is the measure of foreign government bond liquidity, γ'' is the measure of the home government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The sample period is 1999M1–2018M1. Germany government interest rate is used for EUR case. Standard errors in parentheses are standard errors adjusted for cross-sectional correlation. *, **, and *** indicate that the alternative model significantly different from zero at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. The table only reports the coefficient estimates of interest. A table that reports all coefficient estimates is available in Supplementary Appendix A5.

from zero at the 10% level. We find that 42 of the 45 pairs have a negatively significant coefficient on Δi^R_t at the 1% level. This evidence makes manifest that our results are not driven by a single country. Country-by-country regressions also allow the coefficients to be unconstrained and leave room for higher explanatory power. While the median adjusted R^2 of the full sample regressions (17%) is close to the average R-squared of the panel regressions, the maximum adjusted R^2 is 33% for the full sample and 49% post-2008 (in both cases for the AUD-JPY pair).

3.5. *Out-of-sample fit*

The influential work by Meese and Rogoff (1983) shows that standard macroeconomic exchange rate models, even with the aid of ex post data on the fundamentals, forecast exchange rates at short to medium horizons no better than a random walk. In this subsection, we conduct an

TABLE 12
Summary of country-by-country estimation with baseline liquidity measure
$$\Delta s_t = \alpha_+ \beta_1 q_{t-1} + \beta_2 \Delta\tilde{\eta}_t + \beta_3 \Delta i_t^R + \beta_4 \tilde{\eta}_{t-1} + \beta_5 i_{t-1}^R + u_t$$

	q_{t-1} (1)	$\Delta\tilde{\eta}_t$ (2)	Δi_t^R (3)	Adjusted R^2 (4)
		Whole sample: 1999M1–2018M1		
Max	−0.003	1.714	0.250	0.334
Min	−0.116	−9.985	−9.208	0.003
Median	−0.031	−4.160	−5.151	0.170
Mean	−0.038	−4.398	−4.956	0.160
		Pairs that are negatively significant at:		
10%	29	37	42	
5%	25	33	42	
1%	13	29	42	
		2008M1–2018M1		
Max	−0.011	2.905	0.767	0.487
Min	−0.198	−11.860	−14.336	0.012
Median	−0.069	−5.251	−6.993	0.281
Mean	−0.070	−5.393	−7.616	0.267
		Pairs that are negatively significant at:		
10%	31	38	41	
5%	25	36	40	
1%	11	26	39	

Notes: Total number of country pair is 45 (9*10/2). $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i_{j,t}^R$ is the home minus foreign interest rates. Δ is a difference operator. A table that reports all coefficient estimates is available in Supplementary Appendix A6.

out-of-sample forecasting exercise as in Meese and Rogoff (1983) and find that our empirical model significantly outperforms the random walk prediction.[43]

We estimate (17), the model with only interest rate differential and the lagged real exchange rate as explanatory variables, and (16), the empirical model that also includes the liquidity return, using a rolling regression approach. We first use the sample from 1999M1 to 2007M12 for the estimation of regression coefficients.[44] The rolling window is therefore 108 months and the forecast horizon is one month. The first prediction is 2008M1 and the last prediction is 2018M1. We then compare the root-mean-square-error (RMSE) of these models verse the RMSE of a random walk no change prediction ($\Delta\hat{s}_{j,t}^{RW} = 0$). As in the Meese–Rogoff exercise, we use actual realized values of the right-hand-side variables to generate the forecasts.

Table 13 reports the RMSEs of the predictions of models (16), (17) and the random walk prediction. The RMSEs of forecasts from (16) and (17) are lower than the RMSEs of the random walk prediction in 9 out of the 10 countries. We are also interested in testing whether these differences in RMSEs are statistically significant. We adopt the test statistics by Diebold and Mariano (1995) and West (1996) (DMW) which tests the following three null hypothesis: (A) mean-square-error (MSE) of the prediction model (17) and the random walk

43. Engel and Wu (2021) considers pitfalls in assessing the forecastability of the US dollar in the 21st century using various recently proposed predictive variables.

44. We provide robustness of different end date for estimation of regression coefficients (ends at 2005M12 and 2009M12. We also estimate using a recursive regression approach. The results are robust both and are reported in Supplementary Appendix A13.

TABLE 13

Out-of-sample fit comparison of different models. Model (16): Rolling window prediction error of panel regression with liquidity return: $\Delta\hat{s}_{j,t}=\hat{\alpha}_j+\hat{\beta}_1 q_{j,t-1}+\hat{\beta}_2\Delta\tilde{\eta}_{j,t}+\hat{\beta}_3\Delta i^R_{j,t}+\hat{\beta}_4\tilde{\eta}_{j,t-1}+\hat{\beta}_5 i^R_{j,t-1}$. *Model (17): Rolling window prediction error of panel regression without liquidity return:* $\Delta\hat{s}_{j,t}=\hat{\alpha}_j+\hat{\beta}_1 q_{j,t-1}+\hat{\beta}_2\Delta i^R_{j,t}+\hat{\beta}_3 i^R_{j,t-1}$ *and random walk (RW) model:* $\Delta\hat{s}^{RW}_{j,t}=0$

Home currency	RMSE of RW	RMSE of model (17)	RMSE of model (16)	DMW statistics of (17) vs. RW	DMW statistics of (16) vs. RW	DMW statistics of (16) vs. (17)
(1)	(2)	(3)	(4)	(5)	(6)	(7)
AUD	0.0333	0.0311	0.0295	3.212	3.229	1.978
CAD	0.0305	0.0281	0.0268	4.321	4.469	3.007
EUR	0.0285	0.0269	0.0259	3.008	3.685	3.319
JPY	0.0428	0.0404	0.0395	2.768	3.296	2.139
NZD	0.0311	0.0298	0.0282	1.989	2.844	2.716
NOK	0.0363	0.0349	0.0316	2.859	3.539	2.124
SEK	0.0296	0.0287	0.0276	1.722	2.967	2.357
CHF	0.0328	0.0333	0.0332	−0.549	−0.291	0.222
GBP	0.0332	0.0317	0.0313	2.290	3.329	0.770
USD	0.0349	0.0336	0.0312	2.762	4.172	3.570
Home currency	CW statistics of (17) vs. RW	p-value of CW test (17) vs. RW	CW statistics of (16) vs. RW	p-value of CW test (16) vs. RW	CW statistics of (16) vs. (17)	p-value of CW test (16) vs. (17)
	(8)	(9)	(10)	(11)	(12)	(13)
AUD	5.929	0.000***	5.206	0.000***	3.000	0.001***
CAD	6.665	0.000***	6.160	0.000***	3.854	0.000***
EUR	5.340	0.000***	5.915	0.000***	4.437	0.000***
JPY	4.858	0.000***	5.228	0.000***	4.053	0.000***
NZD	4.068	0.000***	4.099	0.000***	3.395	0.007***
NOK	5.279	0.000***	4.901	0.000***	2.452	0.000***
SEK	4.303	0.000***	5.367	0.000***	3.647	0.000***
CHF	0.361	0.359	0.870	0.192	1.202	0.115
GBP	3.910	0.000***	5.388	0.000***	2.010	0.022**
USD	5.335	0.000***	5.521	0.000***	4.299	0.000***

Notes: The 10 currencies used are Australian Dollar (AUD), Canadian Dollar (CAD), Euro (EUR), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF), British Pound (GBP), and United States Dollar (USD). Each row represents a rolling window predictive regression using the Column (1) currency as the home currency and the other nine currencies as foreign currency j. $s_{j,t}$ is the nominal exchange rate between home and foreign country j, defined as home currency price of foreign currency, $q_{j,t}$ is the real exchange rate. $\tilde{\eta}_t$ is the measure of government bond liquidity, $i^R_{j,t}$ is the home minus foreign interest rates. Δ is a difference operator. The rolling window is 108 months. The first estimated coefficient uses sample from 1999M1 to 2007M12. Germany government interest rate is used for EUR case. DMW stands for Diebold and Mariano (1995) and West (1996) and CW stands for Clark and West (2007). The null hypotheses are that the models MSE are equal. The alternative hypotheses are that the larger models MSE are smaller than the nested models. *, **, and *** indicate that the alternative model significantly outperforms the smaller nested model at 10%, 5%, and 1% significance level, respectively, based on standard normal critical values for the one-sided test. Standard errors adjusted for cross-sectional correlation.

model are equal, (B) MSE of the prediction model (16) and the random walk model are equal, and (C) MSE of the prediction model (17) and MSE of the prediction model (16) are equal. The DMW statistics are reported in Columns (5)–(7) in Table 13. Since model (16) nests model (17) and the random walk model, Clark and West (2006) shows that the DMW test statistic should be corrected to account for the fact the regression coefficients are estimated. The Clark–West adjusted test statistic (CW statistics) is asymptotically standard normal and suitable for usual statistics inference. We report the CW statistics and the corresponding p-value of the one-sided alternative test in Columns (8)–(14). We find that the prediction model (17), which includes only the lagged real exchange rate and the interest-rate differential, performs significantly better than random walk in 9 out of the 10 cases (p-values in Column (9).) We find the baseline model with liquidity returns, (16), outperforms the random walk model in nine cases (p-values in Column (11).) In nine cases, we find that the MSE of model (16) are significantly lower than model (17) (p-values in Column (13).) Thus, the random-walk model and the model that does not include liquidity returns are rejected in favour of our baseline model for nine currencies, except for CHF, using the Meese–Rogoff criterion.

3.5.1. Switzerland, January 2015. The model with the liquidity yield included significantly outperforms the random walk model and the traditional model for almost all currencies, but the RMSE for both the liquidity model and the traditional model are higher than the random walk for the single exception of Switzerland.[45] If we eliminate 1 month from the Swiss sample, January 2015, the models also have lower RMSEs than the random walk.

Until that month, the Swiss National Bank had been trying to keep the Swiss franc from appreciating, setting very low interest rates and engaging in massive foreign exchange intervention to keep a ceiling on the value of the franc of CHF1.20 per euro. The SNB lifted the cap in January 2015, which led to a very large franc appreciation that month, despite the low Swiss interest rates. The model performs poorly in that month because the low interest rates should have led to a depreciation of the franc, as the SNB desired.

In fact, all our results reported in previous table are improved, sometimes markedly, for the Swiss franc if that 1 month is eliminated from the sample. It is an extreme outlier.[46] The absolute value of the change in the log of the exchange rate during that month is much greater than for any currency during any month, and it is also a month in which Swiss interest rates were at extremely low values. In particular, in a few of the regressions reported above, the sign on the interest differential was positive rather than negative for Switzerland, but that anomaly disappears when January 2015 is dropped from the sample.

4. CONCLUSIONS

Our empirical findings are good news for macroeconomic models of exchange rates. The government liquidity yield is the “missing link” in exchange rate determination. Not only do we find that liquidity yields are a significant determinant of exchange rate movements for all the largest countries, but we also find that with these included, traditional determinants of exchange rate movements are also important. Our simple regressions have high R^2 values, so can account for a large fraction of exchange rate movements. Our empirical specification is based on a model that is a straightforward extension of the canonical open-economy New Keynesian model to allow for liquidity yields.

An important next step is to dig deeper into the origins of the liquidity yield. Does it arise because government bonds are useful as collateral, perhaps because the market can assess their value without any fear that the counterparty has private information? Or because markets for government bonds are deeper, and therefore more liquid in the traditional sense that they are assets that are less expensive to buy and sell? Or perhaps for some reason, our measure of the liquidity yield is correlated with some other fundamental driver, such as a foreign exchange risk premium that is, in fact, the true driving variable for foreign exchange rates. These possibilities suggest avenues for further theoretical and empirical research.

45. Note that because the CW statistic takes into account estimation error, both models are found to have a positive value of CW statistic, even including January 2015, even though their RMSEs are higher than the random walk model’s RMSE.

46. In Supplementary Appendix A11, we report the results with the extreme outlier in January 2015 dropped.

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Supplementary Data

Supplementary data are available at *Review of Economic Studies* online. And the replication packages are available at https://dx.doi.org/10.5281/zenodo.7038811.

Data Availability Statement

The data underlying this article are available in Zenodo at https://doi.org/10.5281/zenodo.7038811.

APPENDIX

Appendix A1 Derivation of Model of Liquidity Returns

Consider first the problem of the home-country investor. As in Krishnamurthy and Vissing-Jorgensen (2012), Nagel (2016), and Engel (2016), we take a very simple approach to modelling the liquidity service of some assets, by including them in the utility function. In particular, we assume home households maximize:

$$E_0\left\{\sum_{t=0}^{\infty}\beta^t\left[u(c_t)+v\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)\right]\right\}. \tag{B.1}$$

There are six assets in the world economy:
M_t - home country money
M_t^* - foreign country money
B_t - home country government bonds
B_t^* - foreign country government bonds
B_t^m - home country "market" bonds
B_t^{*m} - foreign country "market" bonds

The H subscript in the asset holdings refers to home country holdings of each asset, while F will denote foreign country holdings. $c_t(c_t^*)$ is home (foreign) country consumption.

The utility function for the home household shows that it may get liquidity services from home money, home government bonds and foreign government bonds. We require that holdings of each of these assets must be weakly positive. We will assume that the supplies of the assets and the parameterization of the utility function is such that the home household will always hold home money and government bonds and get liquidity services from those assets, but it may hold a zero amount of foreign government bonds in equilibrium. The utility function $v(\cdot)$ is assumed to be concave, but Inada conditions do not hold for the foreign government bond, so its holdings may be zero. Furthermore, we will follow Nagel (2016) and assume that home bond are a perfect substitute for κ units of money, $0<\kappa<1$. An example of such a utility function is:

$$v\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)=\frac{1}{1-\gamma}\left[\left(\frac{M_{H,t}+\kappa B_{H,t}}{P_t}\right)^{\frac{\varepsilon-1}{\varepsilon}}+\left(\frac{\eta S_tB^*_{H,t}}{P_t}+\mu\right)^{\frac{\varepsilon-1}{\varepsilon}}\right]^{\frac{(1-\gamma)\varepsilon}{\varepsilon-1}} \tag{B.2}$$

where we assume $\varepsilon>1,\gamma>0,0<\kappa<1,0<\eta<1,\mu\geq 0$.

This specification is a slight generalization of that of Nagel (2016) because we assume that there are two non-money assets that might deliver liquidity services.

The period-by-period budget constraint is given by

$$\begin{aligned}&P_tc_t+M_{H,t}+B_{H,t}+B^m_{H,t}+S_tB^*_{H,t}+S_tB^{*m}_{H,t}\\&\quad=P_ty_t+M_{H,t-1}+(1+i_{t-1})B_{H,t-1}+\left(1+i^m_{t-1}\right)B^m_{H,t-1}+S_t\left(1+i^*_{t-1}\right)B^*_{H,t-1}+S_t\left(1+i^{*m}_{t-1}\right)B^{*m}_{H,t-1}\end{aligned} \tag{B.3}$$

Households maximize (B.1) subject to (B.3), and to the constraints $M_{H,t}\geq 0, B_{H,t}\geq 0$ and $B^*_{H,t}\geq 0$. These latter constraints mean that households are unable to issue securities with the same liquidity properties as government securities.

We will assume, for convenience, that as in the New Keynesian model in the article, goods prices in each currency are known one period in advance. The first-order conditions are given by:

$$-\frac{1}{P_t}u'(c_t)+\frac{1}{P_{t+1}}\beta E_t u'(c_{t+1})+\frac{1}{P_t}v_M\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)\leq 0 \tag{B.4}$$

$$-\frac{1}{P_t}u'(c_t)+\frac{1+i_t}{P_{t+1}}\beta E_t u'(c_{t+1})+\frac{1}{P_t}\kappa v_M\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)\leq 0 \tag{B.5}$$

$$-\frac{1}{P_t}u'(c_t)+\frac{1+i^m_t}{P_{t+1}}\beta E_t u'(c_{t+1})=0 \tag{B.6}$$

$$-\frac{S_t}{P_t}u'(c_t)+\frac{1+i^*_t}{P_{t+1}}\beta E_t S_{t+1}u'(c_{t+1})+\frac{S_t}{P_t}v_{B^*}\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)\leq 0 \tag{B.7}$$

$$-\frac{S_t}{P_t}u'(c_t)+\frac{1+i^{*m}_t}{P_{t+1}}\beta E_t S_{t+1}u'(c_{t+1})=0. \tag{B.8}$$

The foreign household's problem is symmetric. The first-order conditions for the foreign household are:

$$-\frac{1}{P^*_t}u'\left(c^*_t\right)+\frac{1}{P^*_{t+1}}\beta E_t u'\left(c^*_{t+1}\right)+\frac{1}{P^*_t}v^*_{M^*}\left(\frac{M^*_{F,t}}{P^*_t},\frac{B^*_{F,t}}{P^*_t},\frac{S_t^{-1}B_{F,t}}{P^*_t}\right)\leq 0 \tag{B.9}$$

$$-\frac{1}{P^*_t}u'\left(c^*_t\right)+\frac{1+i^*_t}{P^*_{t+1}}\beta E_t u'\left(c^*_{t+1}\right)+\frac{1}{P^*_t}\kappa v^*_{M^*}\left(\frac{M^*_{F,t}}{P^*_t},\frac{B^*_{F,t}}{P^*_t},\frac{S_t^{-1}B_{F,t}}{P^*_t}\right)\leq 0 \tag{B.10}$$

$$-\frac{1}{P^*_t}u'\left(c^*_t\right)+\frac{1+i^{*m}_t}{P^*_{t+1}}\beta E_t u'\left(c^*_{t+1}\right)=0 \tag{B.11}$$

$$-\frac{S_t^{-1}}{P^*_t}u'\left(c^*_t\right)+\frac{1+i_t}{P^*_{t+1}}\beta E_t S^{-1}_{t+1}u'\left(c^*_{t+1}\right)+\frac{S_t^{-1}}{P^*_t}v^*_B\left(\frac{M^*_{F,t}}{P^*_t},\frac{B^*_{F,t}}{P^*_t},\frac{S_t^{-1}B_{F,t}}{P^*_t}\right)\leq 0 \tag{B.12}$$

$$-\frac{S_t^{-1}}{P^*_t}u'\left(c^*_t\right)+\frac{1+i^m_t}{P^*_{t+1}}\beta E_t S^{-1}_{t+1}u'\left(c^*_{t+1}\right)=0. \tag{B.13}$$

Equations (B.6) and (B.8) imply the relationship:

$$\left(1+i^m_t\right)E_t u'(c_{t+1})=\left(1+i^{*m}_t\right)E_t\frac{S_{t+1}}{S_t}u'(c_{t+1})=0.$$

If we maintain the assumption of rational expectations and no market frictions, then the assumption that the conditional distribution of exchange rates and consumption is jointly lognormal, we can derive

$$i^{*m}_t+E_t s_{t+1}-s_t-i^m_t=-\tfrac{1}{2}var_t(s_{t+1})-cov_t(m_{t+1},s_{t+1}),$$

where $m_{t+1}=\ln\left(\frac{u'(c_{t+1})}{P_{t+1}}\right)$. If markets are complete, we have $s_{t+1}-s_t=m^*_{t+1}-m_{t+1}$, where $m^*_{t+1}=\ln\left(\frac{u'\left(c^*_{t+1}\right)}{P^*_{t+1}}\right)$. Using this relationship, we can write

$$i^{*m}_t+E_t s_{t+1}-s_t-i^m_t=r_t, \tag{B.14}$$

where $r_t=\frac{1}{2}\left(var_t(m_{t+1})-var_t\left(m^*_{t+1}\right)\right)$. As noted in the text, we do not insist that r_t be interpreted as a time-varying risk premium. It may arise for other reasons as well, such as financial market or expectational frictions, or from the presence of noise traders. While we derived (B.14) from (B.6) and (B.8), it is straightforward to check that equations (B.11) and (B.13) imply the same relationship.[47]

Assume equations (B.4) and (B.5) hold with equality, so that the home agent holds positive amounts of home money and home government bonds. Using (B.5) and (B.6):

$$\frac{1+i_t}{1+i^m_t}+\frac{\kappa v_M\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)}{u'(c_t)}=1. \tag{B.15}$$

47. These relationships are well-known in the literature. See the survey of Engel (2014) for example.

Similarly, using (B.4) and (B.6) gives us:

$$\frac{1}{1+i_t^m}+\frac{v_M\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)}{u'(c_t)}=1.$$

Rearranging these two equations, we find:

$i_t^m-i_t=\frac{\kappa}{1-\kappa}i_t$, which in the text we write as $i_t^m-i_t=\alpha i_t$, $\alpha>0$.

Taking the analogous set of relationships for the foreign country, we arrive at:

$$\left(i_t^m-i_t\right)-\left(i_t^{m*}-i_t^*\right)=\alpha\left(i_t-i_t^*\right).$$

If we had added a shock to liquidity preferences as in Engel (2016), we would then, using (B.14) arrive exactly at the model given in the text, in which

$$i_t^*+E_ts_{t+1}-s_t-i_t=\eta_t+r_t,$$

where $\eta_t\equiv\left(i_t^m-i_t\right)-\left(i_t^{m*}-i_t^*\right)=\alpha\left(i_t-i_t^*\right)+v_t$.

Note that if the home household holds the foreign government bond, we have:

$$\frac{1+i_t^*}{1+i_t^m}E_t\left(\frac{S_{t+1}}{S_t}\right)+\frac{v_{B^*}\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)}{u'(c_t)}=1.$$

Together with equation (B.15), we find:

$$\left(1+i_t^*\right)E_t\left(\frac{S_{t+1}}{S_t}\right)-\left(1+i_t\right)=\left(1+i_t^m\right)\frac{v_B\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)-v_{B^*}\left(\frac{M_{H,t}}{P_t},\frac{B_{H,t}}{P_t},\frac{S_tB^*_{H,t}}{P_t}\right)}{u'(c_t)}$$

Our model implies that if home households hold both government bonds, the difference in the expected rates of return reflects the difference in the liquidity services that the two bonds provide to home households. If the foreign government bond pays a higher monetary return, it must provide a lower liquidity return to the home household in equilibrium.

Appendix A2: Data Source

Generic data source table

Data	**Data source**
Spot exchange rates	Datastream (DS)
1Y forward rates	Datastream (DS)
1M forward rates	Datastream (DS)
1Y Government bond yield	Datastream (DS), Bloomberg (BBG), central banks
1M Government bond yield	Datastream (DS), Bloomberg (BBG), central banks
1Y interest rate swap	Bloomberg (BBG)
1Y credit default swap	Bloomberg (BBG), Markit (MK)
1M LIBOR rates	Datastream (DS)
Consumer price index	IMF IFS
Central bank policy rates	IMF IFS, Norges bank for Norway
Bloomberg exchange rate forecasts	Bloomberg (BBG)

All the variables are created by filling the missing value in the order reported. For exchange rates and forward rates, we do a trilateral cross to get the non-US related exchange rates. For example, the AUD per CAD exchange rate is constructed by log($S_{\mathrm{USD/CAD}}$) - log($S_{\mathrm{USD/AUD}}$). Germany government yield, debt to GDP and CDS are used for EUR.

Specific data ticker table

Data	AUD	CAD	EUR	JPY	NZD	NOK	SEK	CHF	GBP	USD
Spot exchange rates	AUSTDO$	CNDOLL$	USEURSP	JAPAYE$	NZDOLL$	NORKRO$	SWEKRO$	SWISSF$	USDOLLR	-
1Y forward rates	USAUDYF	USCADYF	USEURYF	USJPYYF	USNZDYF	USNOKYF	USSEKYF	USCHFYF	USGBPYF	-
1M forward rates	USAUD1F	USCAD1F	USEUR1F	USJPY1F	USNZD1F	-	USSEK1F	USCHF1F	USGBP1F	-
1Y government bond yield	BBG:GTAUD1Y Govt BBG: C1271Y INDEX	DS:CNTBB1Y BBG: C1011Y INDEX	BBG:GTDEM1Y Govt BBG: C9101Y INDEX	BBG:GTJPY1Y Govt BBG: C1051Y INDEX	BBG:GTNZD1Y Govt DS:NZGBY1Y BBG: C2501Y INDEX	BBG:ST3XY Index BBG: C2661Y INDEX	Sveriges Riksbank website, Treasury bills SE12M BBG:BV010259 Index BBG: C2591Y INDEX	Swiss National Bank, spot interest rate for 1Y govt bond BBG: C2561Y INDEX	BBG:GTGBP1Y Govt	BBG:GB12 Govt FRED BBG: C0821Y INDEX
1M government bond yield	DS:TRAU1MT BBG: AUTE1MYL Index	DS:TRCN1MT BBG: FMSTTB1M Index	DS:TRBD1MT BBG: GETB1M Index	DS:TRJP1MT	DS:TRNZ1MT BBG: NDTB1M Curncy	-	DS:TRSD1MT	DS:TRSW1MT	DS:TRUK1MT BBG: UKGTB1M Index	DS:TRUS1MT BBG: GB1M Index
1Y Interest Rate Swap[a]	BBG:ADSWAP1Q CURNCY BBG:ADSWAP1 CURNCY	BBG:CDSW1 CURNCY	BBG:EUSW1V3 CURNCY BBG:EUSA1 CURNCY	BBG:JYSW1 CURNCY	BBG:NDSWAP1 CURNCY	BBG:NKSW1 CURNCY	BBG:SKSW1 CURNCY	BBG:SFSW1V3 CURNCY BBG:SFSW1 CURNCY	BBG:BPSW1V3 CURNCY BBG:BPSW1 CURNCY	BBG:USSW1 CURNCY
1Y Credit Default Swap	BBG:AUSTLA CDS USD SR 1Y D14 Corp MK:QS973P	BBG:CANPAC CDS USD SR 1Y D14 Corp MK:27CBJG	BBG:GERMAN CDS USD SR 1Y D14 Corp MK:3AB549	BBG:JGB CDS USD SR 1Y D14 Corp MK:4B818G	BBG:NZ CDS USD SR 1Y D14 Corp MK:6B5178	BBG:NORWAY CDS USD SR 1Y D14 Corp MK:6CFB55	BBG:SWED CDS USD SR 1Y D14 Corp MK:8F7220	BBG:SWISS CDS USD SR 1Y D14 Corp MK:HPBCIO	BBG:UK CDS USD SR 1Y D14 Corp MK:9A17DE	BBG:US CDS USD SR 1Y D14 Corp MK:9A3AAA
1M LIBOR rates	ECAUD1M	ECCAD1M	ECEUR1M	EUJPY1M	ECNZD1M	ECNOR1M	ECSWE1M	ECSWF1M	ECUKP1M	ECUSD1M
Bloomberg exchange rate forecasts	FCUSAUQ"XYY" Index	FCUSCAQ"XYY" Index	FCUSEUQ"XYY" Index	FCUSJPQ"XYY" Index	FCUSNZQ"XYY" Index	FCUSNOQ"XYY" Index	FCUSSEQ"XYY" Index	FCUSCHQ"XYY" Index	FCUSGBQ"XYY" Index	X={1,2,3,4} Y=year

[a] See the data appendix of Du *et al.* (2018a) for the detail of the construction, available at: https://sites.google.com/site/wenxindu/data/govt-cip?authuser=0.

Data period

Data	AUD	CAD	EUR	JPY	NZD	NOK	SEK	CHF	GBP	USD
Spot exchange rates	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
1Y forward rates	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
1M forward rates	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
1Y government bond yield	99M11–18M1	99M1–18M1	99M11–18M1	99M11–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M11–18M1	99M1–18M1
1M government bond yield	99M1–00M6, 00M9, 00M12, 01M3, 09M11–13M3	99M1–18M1	10M11–18M1	12M8–18M1	99M1–18M1	NA	99M1–18M1	09M1–18M1	99M1–18M1	99M1–18M1
1Y interest rate swap	99M1–18M1	01M2–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
1Y credit default swap	08M3–18M1	09M4–18M1[a]	08M1–18M1	08M1–18M1	08M2–18M1[b]	08M1–18M1	08M1–18M1	09M1–18M1	08M1–18M1	08M1–18M1
1M LIBOR rates	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
Consumer price index	99Q1–17Q4	99M1–18M1	99M1–18M1	99M1–18M1	99Q1–17Q4	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1	99M1–18M1
Central bank policy rates	99M1–18M1	99M1–17M7	99M1–15M3	NA	99M3–18M1	99M1–18M1	02M7–17M6	00M1–18M1	99M1–16M8	99M1–17M4
Bloomberg exchange rate forecasts	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	06Q1–17Q4	NA

[a]There are multiple missing values in different months.
[b]There are missing values at 2008m3–m4.

Appendix A3: Summary statistics

All the summary statistics of percentage variables scaled by 100 to improve visibility. For example, i of 4.06 represents 4.06% annualized interest rate. Interest rates and forward rates reported are with 1-year tenor. AR1 stands for slope coefficient from an AR1 regression and rmse stands for root mean squared errors.

	Obs	Mean	SD	Min	Max	AR1	rmse	Obs	Mean	SD	Min	Max	AR1	rmse
				AUD							**CAD**			
$s(\ln(S))$	2,052	−0.70	1.56	−4.7	1.1	0.96	0.03	2,052	−0.78	1.55	−4.8	0.9	0.95	0.03
$\Delta s(\%)$	2,052	−0.07	3.09	−13.3	25.5	−0.00	3.09	2,052	−0.04	2.91	−13.8	21.1	−0.06	2.90
$q(\ln(Q))$	2,052	−0.73	1.57	−4.8	1.0	0.96	0.03	2,052	−0.78	1.57	−4.9	1.0	0.96	0.03
$i^R(\%)$	2,052	1.94	1.60	−2.5	6.1	0.97	0.24	2,052	0.05	1.64	−4.1	5.8	0.98	0.20
$\Delta i^R(\%)$	2,052	0.00	0.24	−1.3	1.4	−0.04	0.24	2,052	−0.00	0.20	−1.2	1.0	0.06	0.20
$f-s(\%)$	2,052	2.19	1.76	−2.5	7.0	0.97	0.23	2,052	−0.00	1.82	−5.4	6.1	0.98	0.22
$\tilde{\eta}(\%)$	2,052	0.25	0.31	−1.0	1.6	0.74	0.16	2,052	−0.05	0.30	−1.6	1.1	0.79	0.14
$\Delta\tilde{\eta}(\%)$	2,052	0.00	0.17	−1.3	1.2	−0.30	0.17	2,052	0.00	0.14	−1.3	1.2	−0.25	0.14
$\hat{\eta}(\%)$	2,029	0.09	0.30	−1.1	1.3	0.82	0.15	1,845	−0.12	0.30	−1.5	0.6	0.91	0.12
$\Delta\hat{\eta}(\%)$	2,052	−0.00	0.15	−1.0	1.3	−0.23	0.15	2,052	0.00	0.12	−0.9	1.6	−0.21	0.11
$\tau(\%)$	2,029	0.16	0.20	−0.4	1.1	0.87	0.09	1,845	0.08	0.20	−0.5	1.3	0.85	0.10
$l^R(\%)$	1,121	0.03	0.12	−0.3	0.6	0.84	0.06	530	0.03	0.14	−0.5	0.4	0.84	0.07
$\lambda(\%)$	1,121	0.09	0.31	−1.0	1.3	0.84	0.14	530	−0.03	0.32	−1.3	0.6	0.90	0.12
$i(\%)$	228	4.06	1.50	1.5	6.7	0.99	0.24	228	2.35	1.63	0.4	6.1	0.99	0.20
$IRS(\%)$	228	4.48	1.69	1.6	8.0	0.99	0.25	205	2.23	1.34	0.5	5.7	0.98	0.21
				EUR							**JPY**			
$s(\ln(S))$	2,052	−1.20	1.52	−5.1	0.5	0.87	0.05	2,052	4.18	0.85	2.4	5.5	0.96	0.04
$\Delta s(\%)$	2,052	0.01	2.58	−15.6	17.9	−0.01	2.59	2,052	0.03	3.75	−25.5	14.0	0.02	3.76
$q(\ln(Q))$	2,052	−1.19	1.54	−5.2	0.6	0.89	0.05	2,052	4.24	0.86	2.3	5.6	0.97	0.04
$i^R(\%)$	2,052	−0.64	1.59	−4.2	4.6	0.98	0.19	2,052	−2.42	1.96	−7.0	1.0	0.99	0.21
$\Delta i^R(\%)$	2,052	−0.00	0.19	−1.2	1.1	0.02	0.19	2,052	0.01	0.21	−1.4	1.3	0.17	0.20
$f-s(\%)$	2,052	−0.76	1.75	−4.8	4.8	0.98	0.19	2,052	−2.74	2.02	−7.9	1.1	0.99	0.23
$\tilde{\eta}(\%)$	2,052	−0.11	0.29	−1.5	1.1	0.77	0.13	2,052	−0.32	0.30	−2.0	1.0	0.77	0.15
$\Delta\tilde{\eta}(\%)$	2,052	0.00	0.14	−1.3	1.7	−0.28	0.13	2,052	−0.00	0.16	−1.4	2.5	−0.30	0.15
$\hat{\eta}(\%)$	2,029	−0.01	0.31	−1.3	1.4	0.87	0.13	2,029	−0.23	0.29	−1.8	0.5	0.87	0.13
$\Delta\hat{\eta}(\%)$	2,029	0.00	0.14	−1.2	1.4	−0.23	0.13	2,029	0.00	0.13	−1.3	1.7	−0.20	0.13
$\tau(\%)$	2,029	−0.10	0.21	−1.3	0.5	0.88	0.09	2,029	−0.09	0.23	−1.0	0.7	0.90	0.10
$l^R(\%)$	1,177	−0.04	0.12	−0.7	0.5	0.81	0.06	1,184	0.02	0.12	−0.7	0.5	0.81	0.06
$\lambda(\%)$	1,177	0.03	0.31	−1.4	1.5	0.82	0.15	1,184	−0.19	0.31	−1.9	0.7	0.85	0.13
$i(\%)$	228	1.73	1.75	−0.9	5.0	1.00	0.18	228	0.13	0.23	−0.3	0.8	0.97	0.05
$IRS(\%)$	228	2.06	1.74	−0.3	5.4	1.00	0.19	228	0.26	0.27	−0.1	1.1	0.98	0.06

	Obs	Mean	SD	Min	Max	AR1	rmse	Obs	Mean	SD	Min	Max	AR1	rmse
				EUR							**JPY**			
$s(\ln(S))$	2,052	−0.53	1.56	−4.6	1.3	0.97	0.04	2,052	1.14	1.53	−3.1	2.6	0.97	0.03
$\Delta s(\%)$	2,052	−0.10	3.35	−14.0	21.6	−0.02	3.36	2,052	0.07	2.86	−15.5	20.6	−0.03	2.87
$q(\ln(Q))$	2,052	−0.56	1.58	−4.6	1.2	0.96	0.04	2,052	1.13	1.55	−3.2	2.6	0.98	0.03
$i^R(\%)$	2,052	2.30	1.55	−2.2	7.0	0.97	0.26	2,052	0.88	1.82	−4.1	7.0	0.98	0.23
$\Delta i^R(\%)$	2,052	0.00	0.26	−1.2	1.4	−0.10	0.26	2,052	−0.01	0.22	−1.4	1.1	0.02	0.22
$f-s(\%)$	2,052	2.61	1.74	−2.2	7.9	0.98	0.23	2,052	0.86	2.02	−4.7	7.3	0.98	0.24
$\tilde{\eta}(\%)$	2,052	0.31	0.39	−1.4	2.0	0.76	0.22	2,052	−0.02	0.35	−2.1	1.5	0.78	0.18
$\Delta\tilde{\eta}(\%)$	2,052	0.00	0.24	−1.4	1.4	−0.22	0.23	2,052	−0.00	0.19	−2.5	1.0	−0.22	0.18
$\hat{\eta}(\%)$	2,029	0.17	0.43	−1.3	1.8	0.85	0.21	2,029	0.09	0.35	−1.6	1.8	0.82	0.18
$\Delta\hat{\eta}(\%)$	2,029	−0.00	0.22	−1.2	1.5	−0.22	0.21	2,029	−0.00	0.19	−1.8	1.1	−0.24	0.18
$\tau(\%)$	2,029	0.14	0.23	−0.5	1.4	0.86	0.11	2,029	−0.11	0.19	−1.1	0.5	0.80	0.10
$l^R(\%)$	1,073	0.09	0.14	−0.2	0.8	0.86	0.07	1,054	−0.09	0.13	−0.8	0.1	0.87	0.06
$\lambda(\%)$	1,073	0.09	0.35	−0.9	1.9	0.81	0.17	1,054	0.02	0.36	−1.8	1.6	0.78	0.19
$i(\%)$	228	4.38	1.74	1.8	7.7	0.99	0.26	228	3.10	2.04	0.4	7.0	0.99	0.23
$IRS(\%)$	228	4.87	2.06	2.0	8.9	1.00	0.23	228	3.52	2.17	0.8	7.7	0.99	0.27
				SEK							**CHF**			
$s(\ln(S))$	2,052	1.27	1.52	−2.9	2.8	0.95	0.03	2,052	−0.84	1.55	−4.9	1.0	0.98	0.03
$\Delta s(\%)$	2,052	0.06	2.71	−9.6	18.8	−0.01	2.72	2,052	−0.15	2.91	−19.7	11.9	−0.09	2.90
$q(\ln(Q))$	2,052	1.28	1.53	−2.9	2.8	0.96	0.03	2,052	−0.83	1.58	−4.9	1.1	0.99	0.03
$i^R(\%)$	2,052	−0.44	1.66	−4.7	4.7	0.98	0.20	2,052	−1.58	1.57	−5.7	3.6	0.98	0.20
$\Delta i^R(\%)$	2,052	−0.00	0.20	−1.4	1.0	0.02	0.20	2,052	0.00	0.20	−1.4	0.9	−0.00	0.20
$f-s(\%)$	2,052	−0.42	1.85	−5.3	4.8	0.98	0.20	2,052	−1.97	1.67	−6.2	3.7	0.97	0.22
$\tilde{\eta}(\%)$	2,052	0.02	0.35	−1.5	1.6	0.82	0.16	2,052	−0.38	0.34	−2.1	0.7	0.82	0.17
$\Delta\tilde{\eta}(\%)$	2,052	0.00	0.17	−1.2	1.8	−0.25	0.16	2,052	−0.00	0.18	−1.2	2.2	−0.26	0.17
$\hat{\eta}(\%)$	2,029	0.11	0.33	−1.4	1.4	0.87	0.15	2,029	−0.33	0.28	−1.8	0.5	0.84	0.14
$\Delta\hat{\eta}(\%)$		0.00	0.15	−1.2	1.7	−0.27	0.14		−0.00	0.15	−1.2	1.8	−0.19	0.14
$\tau(\%)$	2,029	−0.09	0.18	−1.2	0.6	0.85	0.08	2,029	−0.06	0.21	−0.9	0.7	0.84	0.10
$l^R(\%)$	1,129	−0.02	0.12	−0.4	0.6	0.81	0.06	905	−0.01	0.13	−0.4	0.8	0.77	0.06
$\lambda(\%)$	1,129	0.18	0.32	−1.5	1.5	0.84	0.15	905	−0.38	0.27	−1.6	0.9	0.85	0.12
$i(\%)$	228	1.92	1.72	−0.9	4.7	1.00	0.19	228	0.89	1.23	−1.0	3.8	0.99	0.17
$IRS(\%)$	228	2.35	1.76	−0.6	5.5	1.00	0.21	228	0.93	1.33	−1.1	4.0	0.99	0.17
				GBP							**USD**			
$s(\ln(S))$	2,052	−1.53	1.50	−5.5	−0.0	0.98	0.03	2,052	−1.00	1.54	−4.9	0.7	0.98	0.03
$\Delta s(\%)$	2,052	0.13	2.92	−14.5	19.7	−0.05	2.93	2,052	0.06	3.14	−17.9	13.2	0.00	3.14
$q(\ln(Q))$	2,052	−1.53	1.52	−5.6	−0.0	0.99	0.03	2,052	−1.01	1.56	−5.1	0.8	0.98	0.03
$i^R(\%)$	2,052	0.32	1.77	−4.2	6.1	0.98	0.21	2,052	−0.40	1.84	−5.3	5.9	0.99	0.21
$\Delta i^R(\%)$	2,052	−0.01	0.21	−1.3	1.2	−0.01	0.21	2,052	0.00	0.21	−1.2	1.2	0.18	0.20
$f-s(\%)$	2,052	0.42	1.92	−4.4	6.4	0.98	0.20	2,052	−0.20	2.03	−6.1	7.1	0.99	0.23
$\tilde{\eta}(\%)$	2,052	0.11	0.31	−1.1	1.6	0.70	0.16	2,052	0.19	0.34	−1.2	2.1	0.81	0.16
$\Delta\tilde{\eta}(\%)$	2,052	0.00	0.18	−1.4	2.1	−0.36	0.17	2,052	−0.00	0.17	−1.3	1.4	−0.22	0.16
$\hat{\eta}(\%)$	2,029	0.14	0.30	−1.1	1.6	0.82	0.15	2,029	0.08	0.32	−1.3	1.5	0.87	0.14
$\Delta\hat{\eta}(\%)$	2,029	−0.00	0.16	−1.3	1.6	−0.31	0.15	2,029	−0.00	0.14	−1.3	1.6	−0.17	0.14
$\tau(\%)$	2,029	−0.03	0.20	−1.4	1.0	0.84	0.09	2,029	0.11	0.18	−0.4	1.2	0.82	0.09
$l^R(\%)$	995	0.03	0.13	−0.5	0.7	0.81	0.07	1,054	−0.02	0.15	−0.6	0.4	0.80	0.09
$\lambda(\%)$	995	0.13	0.32	−0.7	1.8	0.82	0.16	1,054	0.00	0.31	−1.4	1.5	0.81	0.15
$i(\%)$	228	2.60	2.17	−0.0	6.2	0.99	0.22	228	1.95	1.91	0.1	6.1	0.99	0.19
$IRS(\%)$	228	3.06	2.30	0.3	6.8	1.00	0.21	228	2.37	2.10	0.3	7.5	0.99	0.22

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