The term structure of CIP violations*

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Abstract

We show theoretically that persistent deviations from covered interest parity (CIP) across multiple horizons imply simultaneous arbitrage opportunities only if uncollateralized interbank lending rates are riskless. In the absence of observable riskless discount rates, we extract them empirically from interest rate swaps using a simple no-arbitrage framework. They deliver novel quantitative benchmarks that reconcile a zero cross-currency basis with non-zero cross-currency basis swap rates. We quantify that the no-arbitrage benchmark, which is consistent with intermediary-based asset pricing paradigms, accounts for about two thirds of the alleged CIP deviations. The residual pricing errors are associated with the limits-to-arbitrage framework.

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1 Introduction

An important literature provides evidence of persistent violations of covered interest parity (CIP), at least since the 2008 Global Financial Crisis (GFC) (e.g., Ivashina, Scharftstein, and Stein, 2015; Du, Tepper, and Verdelhan, 2018). Figure 1 offers an example of the evidence in the form of a non-zero cross-currency basis, using LIBOR as a proxy for marginal funding costs, and a non-zero cross-currency basis swap (xccy) rate. Both are computed for the Euro to U.S. Dollar exchange rate.[1] Departures from zero for both time series are taken as prima facie evidence that CIP is violated.

The CIP deviations are often interpreted as evidence of intermediary constraints (Du, Tepper, and Verdelhan, 2018; Du, Hebert, and Huber, 2019; Fleckenstein and Longstaff, 2020, among many others). Extant theory implies that intermediary constraints could operate either within the no-arbitrage (NA) environment (e.g., Brunnermeier and Sannikov, 2014, He and Krishnamurthy, 2013) or outside of it. In that case they would be consistent with the limits-to-arbitrage (LTA) paradigm (e.g., Garleanu and Pedersen, 2011, Gromb and Vayanos, 2010, Shleifer and Vishny, 1997).

In this paper, we make a dual contribution. First, we exploit the richness of the term structure of CIP to establish which framework, NA or LTA, represents a more suitable interpretation of the evidence. We find that short horizon departures from zero are fully consistent with NA, and that about two thirds of departures from zero at longer horizons are explained by NA. The remaining one third of the variation in xccy rates is consistent with LTA. Second, we find that the most popular measures of intermediary constraints are related to both NA and LTA components.

We argue that the key to understanding the joint behavior of short- and long-term CIP from the NA perspective is the nature of effective funding rates (EFRs) faced by market participants. A common practice for testing short-term CIP is to use bank funding rates, such as LIBOR, to compute the cross-currency basis. There is significant evidence that bank funding rates may be not appropriate because they can embed default risk, which could deliver a non-zero LIBOR-based forward basis (Du, Im, and Schreger, 2018). An additional complication for obtaining marginal funding rates is that counterparties face differential funding costs depending on their credit risk and bargaining power, especially post-GFC (Rime, Schrimpf, and Syrstad, 2019).

At horizons above one year, CIP is characterized via xccy rates. Xccy are foreign exchange derivative instruments that allow for the exchange of two variable rate loans

[1] We refer to interbank rates as LIBOR, regardless of the currency of denomination. Also, marginal funding costs refer to funding rates of marginal investors, such as intermediaries or sophisticated households, whose stochastic discount factor prices foreign exchange assets (e.g., He and Krishnamurthy, 2018).

denominated in different currencies, typically LIBOR. Because xccy rates are market prices, tests of long-term CIP do not rely on an explicit choice of the appropriate cross-country interest rates. Implicitly, the zero xccy rate benchmark is associated with LIBOR being the correct discount rate.

We show, in a model-free setting, that, under NA, both the basis and xccy rates cannot be simultaneously zero unless the EFR is simultaneously equal to LIBOR and is riskless. That prompts us to depart from the conventional approach and take the view that marginal EFRs are unobservable. That perspective changes the empirical strategy for testing CIP.

Instead of using observable proxies for EFRs to evaluate CIP, we identify the unobserved EFRs from interest rate swaps (IRS) and require that short-term CIP based on forward exchange rates holds in terms of these interest rates. We then assess if xccy rates valued at the unobserved EFRs match their data counterparts. Thus, the valuation of xccy is a relative value exercise, in which we determine their theoretical value using the market values of other instruments.

We implement the valuation exercise by applying a standard stochastic discount factor (SDF) approach to characterize present values of cash flows associated with foreign exchange contracts. To estimate each country's SDF, which implies the EFR, we use an affine no-arbitrage model, and data for G11 countries from January 2000 to December 2019. Upon estimating the associated EFRs, we relate them to various observable candidates. We find that only three variables are significant. The EFRs are closely associated with Treasury yields adjusted for convenience, and sovereign credit default swap (CDS) rates. In addition, there is a modest contribution from interbank rates, i.e., LIBOR. Thus, the estimated EFR is a linear combination of LIBOR, a Treasury rate adjusted for convenience, and a credit risk premium.

We next use the EFRs implied from our model to revisit the evidence on long-term CIP deviations characterized by non-zero cross-currency swap rates. We compute model-implied xccy rates, which were not used in the estimation. While the basis is zero at horizons of one year and less, we find that xccy rates are non-zero, just like in the data. The average 5-year xccy rate across countries during the post-crisis period is 24 basis points (bps) vs. 21 bps in the model. In the cross-section, we find that the departures of xccy rates from zero increase with the departure of EFRs from LIBOR.

Because intermediary asset pricing theories could be both consistent with NA and LTA frameworks, we evaluate how the estimated SDFs and the xccy pricing errors relate to the intermediary factors proposed in the literature. We find that the equity capital of primary dealers (He, Kelly, and Manela, 2017) and the trade-weighted U.S. dollar index (Avdjiev, Du, Koch, and Shin, 2019; Jiang, Krishnamurthy, and Lustig,

2018) are two highly significant variables that are co-moving with the SDFs. Because households cannot trade IRS, which we use to estimate the SDFs, our evidence suggests that intermediaries are the marginal investors in the xccy market.

Next, we show that xccy pricing errors, which are by construction uncorrelated with model-based xccy rates, reflect also market frictions that correlate with financial intermediary health. We find that the same two intermediary variables that affect the SDFs are co-varying significantly with changes in the xccy pricing errors. Cross-sectional regressions show that expected changes in the cross-currency pricing errors line up with their beta sensitivities to both factors. Lastly, we implement the Haddad and Muir (2021) cross-sectional analysis of intermediation costs and assets' exposure to a measure of intermediary risk aversion. We find that swaps are intermediated more than forward contracts, and that their risk aversion sensitivities line up accordingly. The overall evidence is consistent with the role of intermediaries in xccy premiums.

Given that intermediary constraints seem to affect both the SDF and departures from arbitrage, we next evaluate the relative contribution of the two sources of variation in xccy rates. Our variance decomposition of the 5-year xccy rates shows that our no-arbitrage framework explains, on average, about 68% of the variation in their levels. The remaining 32% of their variation that is left unexplained is attributed to LTA.

Taken at face value, our findings lead to a novel and important conclusion. A simple no-arbitrage framework with a minimal set of assumptions is helpful for understanding, to a first order, the dynamics of the forward-spot exchange rate basis and xccy rates. In fact, a key insight is that they can both be zero only under the restrictive assumption that EFRs are equivalent to LIBOR. Thus, our analysis suggests that non-zero and large xccy rates do not necessarily imply arbitrage opportunities.

Our no-arbitrage benchmarks provide a good quantitative account of the levels and dynamics of these rates. At the same time, we associate both the estimated SDFs (NA) and pricing errors (LTA) with factors that are likely correlated with balance sheet costs (Andersen, Duffie, and Song, 2019; Du, Tepper, and Verdelhan, 2018; Fleckenstein and Longstaff, 2020). That result lends credence to the role of intermediary based asset pricing for quantitatively realistic models and calls for additional research for understanding which specific constraints operate through NA or LTA.

Related literature

Discussions and test of CIP deviations go back to Keynes (1923). For pre-GFC analysis, see also Dooley and Isard (1980); Frenkel and Levich (1975); Fletcher and

Taylor (1994, 1996); Pasquariello (2014); Popper (1993); Taylor (1987, 1989). Initial attempts to explain CIP deviations point towards frictions in global intermediation of USD funding, e.g., Baba, Packer, and Nagano (2008), Bottazzi, Luque, Pascoa, and Sundaresan (2012) Coffey, Hrung, and Sarkar (2009), Griffolli and Ranaldo (2011), and McGuire and von Peter (2012). Ivashina, Scharftstein, and Stein (2015) provide a model that rationalizes these effects through foreign banks that cut USD lending more than U.S. banks in crisis times. Coffey, Hrung, and Sarkar (2009) and Bahaj and Reis (2018) support such views by documenting that U.S. swap lines to foreign central banks are effective in reducing CIP deviations. Other explanations relate to an increase in bank counterparty risk associated with the crisis (Baba and Packer, 2009; Csavas, 2016; Levich, 2012; Skinner and Mason, 2011; Tuckman and Porfirio, 2003; Wong, Leung, and Ng, 2016; Wong and Zhang, 2018). Coffey, Hrung, and Sarkar (2009) and Fong, Valente, and Fung (2010) provide evidence in support of both credit and liquidity risk as drivers of CIP deviations.

CIP deviations have persisted even after funding conditions became less strained in the post-crisis period, which is often associated with a reduced ability to conduct arbitrage in response to post-crisis regulation. Du, Tepper, and Verdelhan (2018) document a co-movement in CIP deviations with bank balance sheet constraints. Similarly, Cenedese, Della Corte, and Wang (2020) point towards post-crisis leverage constraints as a source of CIP deviations. Borio, McCauley, McGuire, and Sushko (2016) argue that an increase in hedging demand coupled with reduced balance sheet capacity is responsible for persistent deviations. See also Liao (2016) for arguments of hedging demand. Anderson, Du, and Schlusche (2019) indicate that banks' arbitrage positions reduced in response to a reduction in wholesale funding supply following a regulatory change in the money market mutual fund industry in 2016. Avdjiev, Du, Koch, and Shin (2019) cite banks' leverage constraints as a source of friction, driven by the strength of the dollar. In that same spirit, Du, Hebert, and Huber (2019) argue that CIP deviations may reflect financial intermediaries' shadow costs of capital constraints. Liu (2019) attributes these deviations, among other arbitrage spreads, to LTA.

Rime, Schrimpf, and Syrstad (2019) take a view that LIBOR-based CIP deviations do not necessarily imply arbitrage opportunities, like we do. In contrast to us, they use observable interest rates to estimate feasible transaction costs. Similarly, Kohler and Müller (2018) argue for another set of observable rates, cross-currency repos, which are consistent with CIP. Georgievska (2020) explains CIP deviations with the time-varying spread between risk-free and collateral rates, a.k.a. collateral rental yield, that is estimated using observable proxies. Andersen, Duffie, and Song (2019) question benefits of CIP arbitrage to bank shareholders in the light of required funding value adjustments.

2 Conceptual framework

One of the reasons for the remarkable interest in CIP violations is that they speak to two important paradigms in financial economics. First, CIP is a classic example of the no-arbitrage valuation approach. Second, the persistence of CIP violations likely reflects some form of a constraint in financial markets. Thus, any evidence on CIP violations necessarily speaks to the interaction of the two paradigms. In this section, we discuss the implications of both.

2.1 A no-arbitrage perspective

Our starting point is a setting which precludes arbitrage opportunities in international financial markets. To be clear and perhaps seemingly pedantic, by arbitrage opportunities, we refer to *strict* arbitrage opportunities, that is, implementable trades with non-positive price that generate a non-negative payoff and a positive payoff with positive probability. Our aim is to evaluate empirically to what extent such a framework can rationalize the apparent funding anomalies observed in recent international financial markets.

From the perspective of a U.S. investor, the assumption of the absence of arbitrage opportunities implies the existence of a valuation framework for returns on any assets denominated in U.S. dollars by means of a stochastic discount factor. A stochastic discount factor (SDF) is a stochastic process $M_{0,T}$ such that, for any gross return $R_{0,T}$ between times 0 and T in U.S. dollars, we have

$$E_0 \left[M_{0,T} R_{0,T} \right] = 1. \tag{1}$$

Standard asset pricing theory (see e.g., Duffie, 2001) guarantees that such a stochastic process exists whenever there are no arbitrage opportunities.

By the same token, from the perspective of, say, a Euro area investor, the absence of arbitrage implies the existence of an SDF, say $\widehat{M}_{0,T}$, to value the gross returns on any asset denominated in Euros, $\widehat{R}_{0,T}$ between times 0 and T. That is, we must have

$$E_0 \left[\widehat{M}_{0,T} \widehat{R}_{0,T} \right] = 1.$$

2.2 The role of intermediary constraints

The intermediary constraints considered in the literature take a variety of forms. A unifying theme connecting this work is that intermediaries suffering from low net worth cannot easily exploit profit opportunities. Accordingly, episodes with low intermediary net worth tend to coincide with high risk premia on more intermediated assets. Microfoundations for these limitations come in the form of a variety of constraints imposed on intermediaries, and are of regulatory, institutional, or incentive compatibility nature, among others.

For example, capital requirements may prevent intermediaries from executing trades from a regulatory perspective, while from an institutional viewpoint, margin requirements may apply. Similarly, from an incentive perspective, equity constraints may require intermediaries to have sufficient skin in the game. Under different modeling assumptions, such considerations can be mapped into leverage, collateral, margin, or funding liquidity constraints.

The critical point is that some of these modeling assumptions lead to the existence of a unique SDF and, therefore, rule out arbitrage opportunities. Our no-arbitrage (NA) framework encompasses such models and implicitly reflects the associated constraints. Yet, there exist models in which arbitrage possibilities can arise in equilibrium because of constraints. Thus, our NA approach can provide quantitative guidance on the extent to which CIP violations reflect limits to arbitrage (LTA) associated with the latter perspective.

Table 1 provides an overview of the distinction between NA and LTA with respect to a non-exhaustive list of influential recent contributions. Specifically, the last two columns indicate whether a given approach relies on the emergence of arbitrage opportunities for explaining observed anomalies, especially in derivatives markets. A quick inspection of the table suggests that an important part of the literature operates under NA.

Further, the table enumerates the frictions that are consistent with NA in that they affect the unique SDF. Influential theoretical contributions in this spirit include He and Krishnamurthy (2013), or Brunnermeier and Sannikov (2014), in which intermediaries' stochastic discount factor prices risky assets. These models feature a representative financial intermediary and largely abstract away from heterogeneity in the financial sector. On the empirical side, Adrian, Etula, and Muir (2014), He, Kelly, and Manela (2017), and Haddad and Muir (2021) test asset pricing models based on a stochastic discount factor with intermediary factors.

Another set of papers featured in Table 1 present models in which constraints on intermediaries imply that the law of one price is explicitly violated, in that claims to identical cash flows exhibit different prices in equilibrium. This literature on LTA is also closely related to the literature on segmentation in financial markets in that intermediary constraints prevent cross-market arbitrage. Many models feature heterogeneous agents subject to a variety of constraints. Early theoretical models include

Shleifer and Vishny (1997), who consider mispricing due to the limited capital of arbitrageurs, and Dow and Gorton (1994), who study the impact of holding costs when traders have limited horizons. More recent examples include Garleanu and Pedersen (2011) on the CDS basis, Gabaix, Krishnamurthy, and Vigneron (2007) on the market for mortgage-backed securities, Acharya, Lochstoer, and Ramodarai (2013) on limits to hedging in commodity markets, and Jermann (2020) on negative swap spreads. Liu and Longstaff (2004) analyze portfolio choice for arbitrageurs with collateral constraints. Empirically, Du, Tepper, and Verdelhan (2018) point to the relevance of LTA in explaining CIP violations.

3 CIP in the short and the long run

In this section we develop a no-arbitrage framework that relates derivative contracts written on exchange rates to interest rates at both short and long horizons.

3.1 Forward rates and short-term basis

A currency contract that is struck at time 0 to sell €1 forward at time T for the price $\$F_{0,T}$ has a net USD cash flow of $F_{0,T} - S_T$. A textbook forward valuation strategy of the contract would involve two portfolios that, in the absence of arbitrage, should deliver the same return on investment. One portfolio is long USD by investing \$1 domestically at the annualized domestic $T-$period "risk-free rate" $r_{0,T}$ to receive $e^{Tr_{0,T}}$. The second portfolio is short the USD (long the EUR) by investing the same amount denominated in EUR, €$1/S_0$, at the annualized foreign $T-$period "risk-free rate" $\widehat{r}_{0,T}$, and hedges the currency exposure by selling the EUR proceeds forward at the price $\$F_{0,T}$ per €1. Both investments should have equal payoffs:

$$F_{0,T}/S_0 e^{T\widehat{r}_{0,T}} - e^{Tr_{0,T}} = 0,$$

implying that

$$F_{0,T}/S_0 = e^{T(r_{0,T} - \widehat{r}_{0,T})}.$$

Departing from the textbook treatment, we have to think about appropriate measures of funding rates for the potential arbitrageurs and about handling the counterparty risk. The use of LIBOR was common before the GFC because it was assumed that financial institutions could fund themselves at LIBOR and they were believed to have very low default risk. Both assumptions no longer hold. In particular, post-GFC interbank rates contain non-trivial counterparty credit risk, because they represent uncollateralized borrowing.

One could mitigate that risk by engaging in a CDS position on a borrowing bank. However, that would require knowing the exact credit risk of the panel of LIBOR banks. Alternatively, one could engage in some sort of collateralization. Collateral may take different forms, cash or an asset, USD or EUR-denominated. Each of these variations would impact an effective funding rate (EFR) faced by the parties in a transaction. Indeed, Rime, Schrimpf, and Syrstad (2019) document heterogeneity in funding costs in the post-GFC environment. EFRs would further be affected by fair value adjustments (Andersen, Duffie, and Song, 2019). More generally, as emphasized by Fleckenstein and Longstaff (2018), engaging in a derivatives transaction entails an implicit rental of a dealer's balance sheet. The associated cost should be reflected in the EFRs.

Furthermore, some observers make a distinction between the funding rate and the investment rate. Presumably, an arbitrageur can select a safe counterparty, such as a central bank, and deposit funds at the Treasury rate of that country (e.g., Rime, Schrimpf, and Syrstad, 2019). While that is certainly a possibility for some market players, it is less clear whether a Treasury rate is the marginal investment rate in the foreign exchange markets. Lastly, if CIP does hold for a pair of interest rates, then the representative arbitrageur is indifferent whether to be long USD and short synthetic USD or vice versa, making the distinction between funding and investment rates moot. Thus, we let the data speak regarding this issue, treat both funding and investment rates as unobservable, and do not distinguish them by their economic interpretation.

In order to make accounting for discount rates operational, we rely on the SDF-based pricing relation in Equation (1) to value a forward. Daily marking to market and costly collateral introduce additional cash flows to these contracts. Johannes and Sundaresan (2007) demonstrate that these cash flows represent the opportunity cost of collateral, which can be represented as a dividend yield on an asset.[2] Thus,

$$E_0(M_{0,T}e^{\eta_{0,T}}F_{0,T}) = S_0 \cdot E_0(\widehat{M}_{0,T}e^{\widehat{\eta}_{0,T}}),$$

[2]Here the cost of collateral is a constraint affecting cash flows. On the one hand, the Johannes and Sundaresan (2007) framework is developed in a NA environment. Thus the SDF does not explicitly depend on intermediary frictions. On the other hand, in equilibrium models that we reviewed earlier, frictions are often manifested in the cash flows. Thus, embedding the framework in equilibrium would result in an endogenous SDF that depends on the friction.

where η and $\widehat{\eta}$ represent the domestic and the foreign cost of collateral, respectively, and $\widehat{M}$ is the foreign SDF. Thus,

$$F_{0,T}/S_0 = E_0(\widehat{M}'_{0,T})/E_0(M'_{0,T}), \tag{2}$$

where we use M' as a shorthand for Me^{η}. Therefore M' reflects both discounting as well as the cash flow adjustments arising from collateralization.

The forward price in logs is given by

$$f_{0,T} - s_0 = \log E_0(\widehat{M}'_{0,T}) - \log E_0(M'_{0,T}) = T(r'_{0,T} - \widehat{r}'_{0,T}), \tag{3}$$

where $r'_{0,T} \equiv -T^{-1}\log E_0(M'_{0,T})$ and $\widehat{r}'_{0,T} \equiv -T^{-1}\log E_0(\widehat{M}'_{0,T})$ are the corresponding domestic and foreign rates, respectively. We could relate these interest rates to the marginal EFRs via:

$$r'_{0,T} = r_{0,T} - \eta_{0,T}, \quad \widehat{r}'_{0,T} = \widehat{r}_{0,T} - \widehat{\eta}_{0,T}.$$

Define the forward premium as $\rho_{0,T} = T^{-1}(f_{0,T} - s_0)$. Next, the forward basis is

$$b^r_{0,T} = \rho_{0,T} - (r'_{0,T} - \widehat{r}'_{0,T}) = 0.$$

This first basic step allows us to connect our framework to the literature on cross-currency bases.

Indeed, the literature on CIP violations (e.g., Du, Tepper, and Verdelhan, 2018) explores either the LIBOR or OIS forward basis defined as

$$b^i_{0,T} = \rho_{0,T} - (i_{0,T} - \widehat{i}_{0,T}).$$

where i and $\widehat{i}$ represent LIBOR or OIS and their foreign counterparts. The two bases, b^r and b^i, can be equal to zero simultaneously only if there is no substantive economic difference between r' and i.

Further, the literature on the specialness of U.S. Treasuries (e.g., Du, Im, and Schreger, 2018, Jiang, Krishnamurthy, and Lustig, 2019) evaluates the Treasury forward basis,

$$b^y_{0,T} = \rho_{0,T} - (y_{0,T} - \widehat{y}_{0,T}),$$

where y and $\widehat{y}$ represent U.S. and foreign Treasury yields, respectively. This basis is interpreted as the relative convenience yield of Treasuries. Implicit in this interpretation is the existence of interest rates at which the basis is equal to zero.

In this paper we take a view that the theoretical interest rates r' and $\widehat{r}'$ are not readily observable. Thus, the CIP condition (3) cannot be tested with readily available benchmark interest rates. Indeed, in the post-GFC world, one would need to observe effective funding costs of all players in the forward markets and be able to aggregate them correctly to economically match the reported forward prices. Thus, we treat r' and $\widehat{r}'$ as these unobservable aggregate funding costs implicit in the forward premium. Rime, Schrimpf, and Syrstad (2019) take the complimentary route of estimating the marginal funding rates directly using information about wholesale money market funding from non-bank investors.

Our perspective is consistent with the view of Binsbergen, Diamond, and Grotteria (2019) that prices of risky financial assets reflect risk-free rates stripped of the convenience premium. Relatedly, Fleckenstein and Longstaff (2018) use Treasury note futures contracts to back out intermediaries' shadow funding costs.

Of course, the empirical fact that $b^i \approx b^r \approx 0$ before the crisis, and, subsequently b^i has departed from b^r is still worthy of an explanation even if it does not necessarily violate no-arbitrage conditions. After estimating the relevant discount rates, we study the reasons behind the departure of b^i from b^r.

As a next step, we analyze how our framework speaks to long-term CIP violations as represented by cross currency basis swaps (xccy). That analysis leads us to an empirical strategy allowing us to evaluate the no-arbitrage predictions about forward premiums and xccy rates.

3.2 Cross-currency basis swap rates

Xccy contracts are OTC derivative instruments that allow investors to simultaneously borrow and lend in two different currencies at floating interbank rates such as LIBOR or EURIBOR. Specifically, it involves an exchange of principal in two different currencies both at inception and at the expiration date of the swap, as well as an exchange of floating cash flows linked to interbank rates. The exchange of face values of the domestic and foreign legs of xccy are matched using the spot exchange rate between both currencies. Unpredictable variation in exchange rates thus involves a non-trivial amount of exchange rate risk at the maturity date of the contract. The price of the xccy is usually quoted as a fixed spread X over the floating foreign currency denominated interest rate.

We examine xccy contracts from the perspective of an investor who, at initiation of the contract, is paying S_0 dollars and is receiving one euro. Table 2A illustrates the cash flows associated with such a position. The investor would receive floating dollar interest payments at the rate i_t on the USD leg at each date $t+1$, and make floating euro interest payments at the rate $\widehat{i}_t+X$ on the EUR leg at each date $t+1$. The initial principal payments would have to be reversed at maturity T. The present value of all expected future cash flows on the USD leg of the cross-currency basis swap is given by

$$\phi_{0,T} = S_0 \left(-1 + \sum_{t=1}^{T} E_0 \left[M'_{0,t} i_{t-1} \right] + E_0 \left[M'_{0,T} \right] \right),$$

and the present value of all expected future cash flows on the EUR leg of the cross-currency basis swap is given by

$$\widehat{\phi}_{0,T} = +1 - \sum_{t=1}^{T} E_0 \left[\widehat{M}'_{0,t} \left(\widehat{i}_{t-1} + X_{0,T} \right) \right] - E_0 \left[\widehat{M}'_{0,T} \right].$$

here we use the M' notation again to account for the cost of collateral in swap transactions. The xccy is fairly priced if both the USD and the EUR legs have the same value in USD, i.e., $\phi_{0,T} + S_0 \widehat{\phi}_{0,T} = 0$. The condition yields the formula for the constant maturity xccy swap rate $X_{0,T}$:

$$\begin{aligned} X_{0,T} &= \left(\sum_{t=1}^{T} E_0 \left[\widehat{M}'_{0,t} \right] \right)^{-1} \\ &\times \left[\left(\sum_{t=1}^{T} E_0 \left[M'_{0,t} i_{t-1} \right] + E_0 \left[M'_{0,T} \right] \right) - \left(\sum_{t=1}^{T} E_0 \left[\widehat{M}'_{0,t} \widehat{i}_{t-1} \right] + E_0 \left[\widehat{M}'_{0,T} \right] \right) \right]. \end{aligned} \tag{4}$$

Intuitively, the xccy rate is pinned down by the difference in prices between two floating rate notes tied to LIBOR. Floating rate notes are valued at par at the interest rate reset date provided that the discount rate is equivalent to the floating rate coupon (Duffie and Singleton, 1997; Litzenberger, 1992; Ramaswamy and Sundaresan, 1986). A discount rate other than LIBOR would trivially imply a non-zero X without violating no-arbitrage conditions.

Anecdotally, full collateralization, which was prevalent by the late 1990s, led market participants to use the OIS rates instead of the LIBOR rates for discounting starting in 2007. By the end of 2008, the whole industry had switched to OIS (e.g., Cameron, 2013, Hull and White, 2013, Spears, 2019). That would immediately imply a non-zero X. The advantage of our valuation via the SDF is that we do not have to take a stand on a specific reference rate to obtain the discount factor. The empirical question is whether an estimate of X can quantitatively be similar to the observed one while simultaneously respecting a zero basis.

3.3 Connection between short- and long-term CIP

Our general multi-horizon framework allows us to revisit the original case of the forward rates. We show that if the xccy is a one-period instrument, then it collapses to the short-term basis. To see, this, assume that $T = 1$, where 1 refers to absence of interim payments between 0 and T. Then, using the relation between the forward premium and SDFs in Equation (2), the floating-for-floating xccy expression in Equation (4) simplifies to:

$$
\begin{aligned}
X_{0,T} &= \left(E_0\left[\widehat{M}'_{0,T}\right]\right)^{-1}\left[S_0 E_0\left[M'_{0,T}\right](1+i_{0,T})^T - S_0 E_0\left[\widehat{M}'_{0,T}\right]\left(1+\widehat{i}_{0,T}\right)^T\right] \\
&= F_{0,T}(1+i_{0,T})^T\left[1-\frac{F_{0,T}}{S_0}\frac{(1+\widehat{i}_{0,T})^T}{(1+i_{0,T})^T}\right] = F_{0,T}(1+i_{0,T})^T\left[1-e^{b^i_{0,T}}\right]. \qquad (5)
\end{aligned}
$$

If LIBOR is the effective discount rate, i.e., $E_0[M_{0,T}] = (1+i_{0,T})^{-T}$, then the LIBOR (log LIBOR) basis is equal to 1 (0). Equation (5) shows that, in that case, $X_{0,T}$ must be zero as well. This is consistent with the description in the previous section that xccy rates are trivially zero if discount rates are different from LIBOR.

If LIBOR is not the effective discount rate, then the log LIBOR basis is equal to:

$$
b^i = T^{-1}\log\left(1-\frac{X_{0,T}}{F_{0,T}}(1+i_{0,T})^{-T}\right) \approx c - a[T^{-1}(x_{0,T}-f_{0,T}) - i_{0,T}],
$$

where $a = e^{T\bar{x}}/(1-e^{T\bar{x}})$, $\bar{x} = E[T^{-1}(x_{0,T}-f_{0,T})-i_{0,T}]$, $c = T^{-1}\log(1-e^{T\bar{x}})+a\bar{x}$. Thus, a non-zero LIBOR basis does not necessarily contradict no-arbitrage. Unfortunately, this expression is not testable because liquid xccy and forwards overlap only at a 1-year horizon. Even then, a forward contract has a single cash flow, while xccy has four quarterly payments.

4 Empirical framework

In this section we develop a methodology that allows us to take the concepts from the previous section to the data.

4.1 Empirical strategy

In order to evaluate the quantitative success of our view of forward premiums and xccy rates, we need to obtain estimates of collateral-adjusted SDFs M' and $\widehat{M}'$. Our strategy is to infer these objects from prices of domestic and foreign interest rate swaps (IRS). Our approach is to implement it in the most simple-minded way possible, akin to bootstrapping discount factors from bond prices.

The literal bootstrapping is not feasible because, as emphasized in the previous section, there is a potential dichotomy between interest rates related to cash flows and the ones affecting discount factors. Thus, we use a simple no-arbitrage dynamic term structure model, which allows to account for these differences. We estimate this model on a country-by-country basis to fit the interbank rates and the corresponding IRS curve of a given country.

Having obtained a collection of M' and $\widehat{M}'$, we can evaluate expressions in Equations (3) and (4) and compare them to the observed forward premiums and xccy rates, respectively. Importantly, we are not using the data on $X_{0,T}$ for the estimation. We exploit the identity $b^r_{0,T} = 0$ to identify foreign rates $\widehat{r}$ via the observed forward premiums.

4.2 Pricing xccy in terms of interest rate swaps

We infer domestic and foreign SDFs from IRS rates because the xccy cash flows are, by no-arbitrage, linked to those of IRS contracts. Specifically, we swap both the USD and the EUR interest rates into fixed rates using an IRS in each currency, at prices CMS and $\widehat{CMS}$, respectively (CMS stands for "constant maturity swaps"). We illustrate these cash flows in Table 2B.

The net cash flows of the USD leg $\pi_{0,T}$ of the fixed-for-fixed xccy are given by

$$\pi_{0,T} = S_0 \left(-1 + \sum_{t=1}^{T} CMS_{0,T} E_0 \left[M'_{0,t} \right] + E_0 \left[M'_{0,T} \right] \right)$$

and the present value of expected future cash flows on the EUR leg is given by

$$\widehat{\pi}_{0,T} = \left(+1 - \sum_{t=1}^{T} \left(\widehat{CMS}_{0,T} + X_{0,T} \right) E_0 \left[\widehat{M}'_{0,t} \right] - E_0 \left[\widehat{M}'_{0,T} \right] \right).$$

The xccy is priced fairly if $\pi_{0,T} + S_0 \widehat{\pi}_{0,T} = 0$, which leads to the alternative expression for $X_{0,T}$:

$$
\begin{aligned}
X_{0,T} \;&=\; \left(\sum_{t=1}^{T} E_0 \left[\widehat{M}'_{0,t} \right] \right)^{-1} \qquad (6) \\
&\times \; \left(CMS_{0,T} \sum_{t=1}^{T} E_0 \left[M'_{0,t} \right] - \widehat{CMS}_{0,T} \sum_{t=1}^{T} E_0 \left[\widehat{M}'_{0,t} \right] + E_0 \left[M'_{0,T} \right] - E_0 \left[\widehat{M}'_{0,T} \right] \right).
\end{aligned}
$$

Thus, we express the xccy rate in terms of (observable) interest swap rates and (unobserved) discount factors, $M'_{0,t}$ and $\widehat{M}'_{0,t}$.

4.3 A model

We describe our model for the U.S. only. All other countries have the same notations augmented with hats, $\widehat{\cdot}$. We assume that the unobservable state is captured by a vector z_t that follows a VAR(1):

$$z_{t+1} = \Phi z_t + \Sigma \varepsilon_{t+1}.$$

The spot interest rate is $r_t = \delta_{r,0} + \delta_r^\top z_t$, and the SDF is

$$-\log M_{t,t+1} = r_t + \nu_t^\top \nu_t / 2 + \nu_t \varepsilon_{t+1},$$

where the conditional volatility of the log SDF, $\nu_t = \Sigma^{-1}(\nu_0 + \nu \cdot z_t)$, is often referred to as the price of risk. We assume that the opportunity cost of collateral is $\eta_t = \delta_{\eta,0} + \delta_\eta^\top z_t$. As a result we can construct the discount factor:

$$E_0 \left[M'_{0,T} \right] = E_0 \left[M_{0,T} e^{\eta_{0,T}} \right] = E_0 \left[\prod_{t=0}^{T-1} M_{t,t+1} e^{\eta_t} \right].$$

While we are using the interest rate r_t in the same way as a risk-free rate appears in a classical framework, we have to be careful with its interpretation. We cannot estimate the true risk-free rate using the data on OTC interest rate derivatives alone. What we estimate is the EFR in these markets. A different EFR could be associated with different markets. See, for example, Binsbergen, Diamond, and Grotteria (2019) for equity options, and Fleckenstein and Longstaff (2018) for Treasury note futures.

As highlighted earlier, this is the simplest model one could entertain. The model lacks various forms of heteroscedasticity (regimes, stochastic volatility). It also does not account for various regulatory changes that took place in the money markets during our sample. All these features would help us fit the data better, although with loss of parsimony. Our objective is not to provide the best possible fit, but to examine how much of the xccy valuation we can explain using the most minimal set of assumptions.

We further assume that the observable one-month LIBOR rate is given by $i_t = r_t + \delta_{i,0} + \delta_i^\top z_t$. This assumption is consistent with the intensity-based approach to modeling credit risk (e.g., Duffie and Singleton, 1999). We connect i_t to LIBOR rates corresponding to longer horizons via hypothetical LIBOR bonds $L_{0,T}$ discounted at the continuously compounded yield $i_{0,T} = T^{-1} \log(1 + i^q_{0,T} \cdot T \cdot 30/360)$ where $i^q_{0,T}$ denotes a quoted LIBOR rate and $T \leq 12$ corresponds to maturities of up to 12 months.[3] As a result,

$$L_{0,T} \equiv \exp\left(-i_{0,T} \cdot T\right) = E_0 \left[M_{0,T} e^{-\sum_{t=0}^{T-1} (\delta_{i,0} + \delta_i^\top z_t)} \right].$$

We are not using the cost of collateral η here because LIBOR represents uncollateralized lending.

Now we can use the 3-month LIBOR rates for the computation of the IRS. Here we discount all cash flows accounting for the cost of collateral η_t. The standard argument then implies:

$$CMS_{0,T} \;=\; \frac{\sum\limits_{t=1}^{T} E_0 \left[M'_{0,t} \left(e^{i_{t-1,t}} - 1 \right) \right]}{\sum\limits_{t=1}^{T} E_0 \left[M'_{0,t} \right]}. \tag{7}$$

This representation of the IRS is stylized to conserve on notation. In the implementation, we account for the actual payment frequencies of the contracts. We discuss institutional details in the online appendix.

4.4 Identification

As specified, the model is under-identified. We adopt the canonical form used by Joslin, Le, and Singleton (2013) and choose the latent state z_t so that the matrix $\Phi - \nu$ governing the dynamics under the risk-adjusted distribution is diagonal. Further, because both loadings δ_r and covariance matrix Σ control the scale of r_t, we set the former to unity. All other parameters are free.

[3] The day count convention for LIBOR rates is act/360. We use 30/360 as the daycount convention given that it is numerically close to act/360, and it simplifies the implementation.

We have an unusual situation in that we have two reference interest rates in the model (i and r), and one of them (r) is not observable. Furthermore, the cost of collateral (η) is not observable either. We rely on three observations to identify r and η. First, as highlighted earlier, $b^i \approx 0$ before the crisis. Second, by assumption, $b^r = 0$ throughout the whole sample. Third, the cost of collateral appears in the valuation of IRS, but not LIBOR.

We rely on the first observation and assume that $r_t = i_t + u_t$, $u_t \sim (0, \sigma_u^2)$ before the crisis (December 2007). The variance of the observation noise u_t is selected to be 1% of the variance of 1-month LIBOR. This approach is also consistent with the widespread view, both in academia and industry, that i is a better proxy for r than a Treasury yield y, because of the convenience premium present in Treasuries and the "refreshed AA" quality of banks in the LIBOR panel. This assumption allows us to pin down Σ (because δ_r is set to 1 for identification purposes). Thus, once the scale of the state variables is fixed, we can identify r and i separately in the post-crisis period.

We rely on the second observation to identify the level of foreign EFRs by setting them to $\widehat{r}'_{0,T} = r'_{0,T} - \rho_{0,T}$. We pick $T = 3$ months although the relation should hold for any T. We use $T = 6$ months as an external model validation. Finally, the third observation, in combination with the other two assumptions, identifies r and η separately.

5 Evidence

We first discuss the data, and then present the model's implications for the forward basis and xccy rates.

5.1 Data

We use a panel data set on interest and exchange rates for G11 countries from January 2000 to December 2019. G11 currencies include the USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, NOK.[4] Specifically, we obtain information on spot and forward exchange rates with maturities of 1, 3, 6, and 12 months. We adopt the convention of measuring exchange rates as the USD price per unit of foreign currency.

[4]DKK is pegged to EUR, but we are not duplicating the analysis because of our focus on the valuation of forward and xccy contracts rather than their realized payoffs. As we have shown, the valuation primarily depends on the local interest rates.

We also source closing prices for cross-currency basis swap rates with maturities of 1, 3, 5, 7, 10, 15, 20, and 30 years. In addition to data on exchange rates, we source country-specific information on Treasury yields, interbank rates (LIBOR), and interest swap rates with matching maturities. For comparability, our data set is similar to that in Du, Tepper, and Verdelhan (2018). All data are sourced from Bloomberg. Details about data sources are discussed in the online appendix.

The black lines in Figure 2 display the 3-month and 6-month LIBOR bases, $b^i_{0,T}$, and cross-currency basis swap rates $X_{0,T}$ for the 5-year and 20-year contracts, for selected currencies, NZD, EUR, and JPY. The full set is provided in the online appendix. Consistent with previous evidence, we observe negative rates, during and post crisis, for all countries except for AUD, CAD, and NZD, which become positive during the same time period. The set of left columns in Figure 3 provide the corresponding summary statistics. Tables supporting this figure are provided in the online appendix. The magnitudes are largely consistent with Du, Tepper, and Verdelhan (2018) with a proviso that we have a longer sample, and a slightly different delineation between the pre-, during, and post-crisis periods. Table 3A displays the results from a principal component analysis (PCA) of xccy rates by currency. The rates exhibit a clear factor structure with three factors explaining most of the variation in their term structure.

5.2 Results

Fitting a term structure model to a swap curve is a standard exercise that is not expected to yield any surprises. Table 3B shows that the LIBOR-IRS curves exhibit a three-factor structure. Thus, we choose the dimension of the state vector z_t to be 3 in our model. We set the cost of collateral to a constant $\delta_{\eta,0}$ in each country, as we faced difficulties in detecting statistically significant variation in this variable. It ranges between 2 and 5 bps (annualized) for most countries, and clusters around 15-20 bps for AUD, CAD, and NOK. Each country-specific model fits the respective IRS curves well.

5.2.1 Forward bases

As one measure of fit, we report a dimension of the model that is particularly relevant for us. The first row of Figure 2 displays the time-series of the theoretical 3-month basis $b^r_{0,0.25}$ (blue line). The column labeled 'Model' in Panel A of Figure 3 shows the summary statistics. Overall, the basis is close to zero in contrast to the LIBOR basis.

The second row of Figure 2 and the column labeled 'Model' in Panel B of Figure 3 report similar information for the 6-month basis $b^r_{0,0.5}$. The 6-month forward rates were not used for estimation, so this is a first glimpse of our model's extrapolation capacity. While the fit is not as good as at the 3-month horizon, b^r is much closer to zero and less volatile than the companion LIBOR basis b^i.

5.2.2 Xccy rates

We use the estimated SDFs M and $\widehat{M}$ to construct xccy rates using Equation (6). The third and fourth rows of Figure 2 and the column labeled 'Model' in Panels C and D of Figure 3 display the results for 5-year and 20-year contracts, respectively. The averages implied by the model are consistent with the evidence. Nevertheless, there are departures between the model and the data in the time-series.

Before the crisis, the observed and the theoretical xccy rates are visually similar. During the crisis, we see a broad switch in the level of X. For some currencies, like CHF, DKK, or EUR, the switch is broadly consistent with the evidence. In some cases, like AUD, or CAD, it is more muddled. Mechanically, the model generates the change because our identifying assumptions allow for departures between r and i. The economic interpretation of the specific quantitative effect is straightforward: LIBOR has adjusted to reflect the riskiness of the banking sector. After the crisis, the relation between the observed and the theoretical X is weaker, reflecting the fact that the model can match the general trend in xccy rates, but not the local deviations. These local deviations probably reflect various constraints faced by the market participants that are not accounted for in our framework. We investigate this possibility in the subsequent analysis.

6 Interpretation of the evidence

We first relate the estimated EFRs to observable variables. Second, we evaluate the economic sources of the variation in the SDFs and in the xccy pricing errors.

The first exercise has a dual purpose: investigating the source of the empirical success of our model and developing economic intuition for the estimated EFRs. One might worry that our results are driven by the latent cost of collateral η, which mechanically adjusts LIBOR, as a proxy for r, so that $r' = r - \eta$ prices assets correctly. That we set η to a constant should alleviate this concern as r, rather than r', is doing all the work in our model. More broadly, it is useful to understand the structure of EFRs in the fixed-income swap markets. Thus, one would want to evaluate which variables r is related to. The second exercise allows us to quantify the contribution of the NA and LTA components and to examine their relation to measures of intermediary constraints.

6.1 Effective funding rates

What would be an appropriate observable proxy for the EFR r? We use two approaches to address this question. First, we construct such a proxy by theorizing about the relation between various observable rates. Second, we implement a panel regression that allows us to consider a large number of possibly relevant variables, and select the ones that co-move with r in a significant fashion.

Yields on Treasury bonds, $y_{0,T}$, continue to serve as a natural starting point to think about EFRs. We know three reasons for why that may not be a good proxy. Dealers cannot fund themselves at government rates. Next, Treasury yields reflect a convenience premium (e.g., Krishnamurthy and Vissing-Jorgensen, 2012). Lastly, in the post-crisis environment, Treasury yields reflect credit risk (e.g., Chernov, Schmid, and Schneider, 2020).

With these considerations in mind, we study the following proxy for the EFR:

$$\widetilde{r}_{0,T} \equiv y_{0,T} + \lambda_{0,T} - CDS_{0,T},$$

where λ is the convenience premium and CDS is a premium on a sovereign credit default swap. The yield and CDS information are readily available. We use the U.S. Refcorp - Treasury spread to estimate λ in the U.S. (Longstaff, 2004; Li and Song, 2019).[5] Having obtained the U.S. convenience premium λ, we obtain the foreign $\widehat{\lambda}$ from the Treasury basis via

$$\widehat{\lambda}_{0,T} = \lambda_{0,T} - b^y_{0,T} + (\widehat{CDS}_{0,T} - CDS_{0,T}) + (\widehat{\eta}_{0,T} - \eta_{0,T}).$$

As mentioned earlier, the last term is small and constant in our model. Du, Im, and Schreger (2018) and Jiang, Krishnamurthy, and Lustig (2018, 2019) work through similar computations in their empirical work. The key difference is that they do not estimate country-specific λ separately.

Because reliable CDS information is available only at maturities starting at 1 year, the shortest interest rate that we can evaluate is for $T = 1$ year. Figure 4 plots $r_{0,T}$ and

[5]The bonds of the Resolution Funding Corporation (Refcorp) are as safe as U.S. Treasuries because its debt is effectively guaranteed by the U.S. government. The Refcorp bonds also have the same tax treatment.

its proxy $\widetilde{r}_{0,T}$. We see that the proxy is tracking the EFR quite well, but there are also evident departures. Japan is the strongest example of large discrepancies, and Sweden is one of the better fitting ones. While there is a reasonably close association between $\widetilde{r}$ and the model-implied interest rate, differences between them are not surprising. Even if there is no noise associated with the ingredients of $\widetilde{r}$, it does not account for risk associated with the interbank market, and so it may not be capturing the EFR of dealers.

As the relation between the observable, the theorized, and the estimated EFRs is not perfect, we investigate whether other variables are worth considering. Our candidates are the ingredients of $\widehat{r}$ taken separately: Treasury yields, CDS premiums, and liquidity proxies. We also consider their combinations: $y + \lambda$ (convenience-adjusted Treasury), $y - CDS$ (credit-risk-adjusted Treasury), and $\widetilde{r}$ itself. Furthermore, we consider rates at which banks can fund themselves on an uncollateralized basis. This includes LIBOR as a pre-GFC reference rate, and OIS as a post-GFC reference rate for swap contracts. Finally, we consider a set of U.S - only variables: the effective Federal Funds rate (EFFR) as another measure of near-money rates, the certificate of deposit - Treasury spread as a measure of the opportunity cost of collateral (Nagel, 2016), and the interest rates implicit in S&P 500 option box spreads (Binsbergen, Diamond, and Grotteria, 2019). We provide a detailed overview of all data sources in the online appendix.

Table 4 provides evidence regarding the relation between changes in r and changes in candidate variables by regressing the former on the latter at a monthly frequency. We run regressions for individual variables and for all of them taken together. Not all of them are available at each horizon. We focus on tenors T of 3 months and 1 year. The row MAT reflects which horizon is used for a specific regression. The two multivariate regressions in columns (12) and (13) include all the variables that are available at the two horizons, respectively. We run panel regressions and add currency fixed effects to focus on the within currency variation. We add month fixed effects to absorb common variation across currencies. In particular, these fixed effects account for the role of USD swap lines extended to foreign central banks in stress periods (Bahaj and Reis, 2018; Coffey, Hrung, and Sarkar, 2009). The common U.S. variables are not compatible with month fixed effects as they are absorbed by them. Thus, U.S. variables do not appear in the multivariate regressions, and we do not use month fixed effects in the corresponding univariate regressions.

When evaluating the univariate regressions, we focus on the magnitude of the estimated coefficient (the closer to 1 the better) and the within R^2. The leading variables here are LIBOR and the convenience-adjusted Treasury with coefficients around 0.6, and R^2 around 0.5. The weakest variables are the U.S.-only ones: EFFR, CD-Treasury spread, and the option box spread with coefficients below 0.2 and R^2 below

0.06. Our initial proxy for the EFR $\widetilde{r}$ occupies an intermediate position with the coefficient of 0.48 and R^2 of 0.33.

Moving to multivariate regressions, we find that, at the 3-month horizon, LIBOR and convenience-adjusted Treasury rates are the two variables that remain significant. This finding is supportive of our prior that interbank funding costs should also be related to EFRs in swap markets. Initially, we have allowed y and λ to appear separately, but the estimated coefficients were nearly identical, so we have combined them into one intuitive variable with no loss in R^2. We had also included the other candidate variables in the multivariate regression, but we subsequently removed them because they turned out to be statistically insignificant.

At the 1-year horizon, the CDS premium emerges as a variable that is statistically important in addition to LIBOR and the convenience-adjusted Treasury rates. The negative coefficient is intuitive, as it implicitly adjusts Treasuries for credit risk. Recall, that we did not use CDS information in the 3-month regression because 3-month CDS rates are not available. Thus, the best fit uses the same ingredients as our theoretical proxy $\widetilde{r}$, but with somewhat different weights.

It is interesting that OIS is not significant in multivariate regressions. Some might view this as surprising in the context of common wisdom that the right discount rate for swaps must be OIS because of collateralization. Our evidence is consistent with Rime, Schrimpf, and Syrstad (2019) who argue that OIS contracts, being derivatives, are not well suited for raising funds.

Figures 4 and 5 compare the estimated r with the best prediction according to the multivariate regressions presented in columns (12) and (13) in Table 4. The predictions are for the changes, so we obtain predictions for levels by cumulating the changes. At the 1-year horizon, the predicted r is more accurate than $\widetilde{r}$ and, in fact, is very close to r. At the 3-month horizon, the prediction tracks r almost perfectly.

Evidently, the credit risk adjustment via CDS should not be one-for-one with the Treasury yield itself. Furthermore, LIBOR is the missing ingredient in the theoretical proxy $\widetilde{r}$. One explanation for that is that the proxy is noisy and the adjustments in CDS and LIBOR happen to soak up these errors. Alternatively, a smaller adjustment via CDS is consistent with a view that sovereign credit risk is not the only risk reflected in CDS premiums. Consistent with the notion of dealers' EFR, mixing in some LIBOR risk could indicate that sovereign CDS premiums also reflect bank risk due to the two-way feedback effects between sovereign and financial risk (Acharya, Drechsler, and Schnabl, 2014).

6.2 Is the EFR different from LIBOR?

As mentioned earlier, one concern could be that LIBOR is in fact a good proxy for the EFR and, thus, all the explanatory power in the model is driven by the latent cost of collateral. First, we can see from column (3) of Table 4 that the regression coefficient of Δr on Δi is significantly different from one. Next, Figure 5 compares explicitly our EFR with LIBOR. It is evident that i is substantively different from r during the post-crisis period (we set them to be similar before the crisis as part of our identification strategy). Thus, the difference in forward basis and xccy valuation comes from a different funding rate.

As a further characterization of the difference between the two rates, we consider the theoretical connection between this difference and swap rates X. The no-arbitrage framework suggests that xccy rates are zero only under the strong assumption that the EFR is identical to LIBOR. Thus, under the null of our model, xccy rate deviations from zero should be positively related to the differences between observed LIBOR rates and our model-implied EFRs.

We test this hypothesis by projecting, in a pooled cross-section, the absolute values of the observed 5-year xccy rates on the 3-month $i-r$ spread. We cluster standard errors by month to account for cross-sectional dependence in the residuals. The results are reported in Table 5.

In column (1), we find that xccy rates deviate on average about 5 bps more from the zero benchmark when the $i - r$ spread is greater by one percentage point. In column (2), we add monthly time fixed effects to provide a fairer comparison across time periods. That specification suggests a 21 bps xccy rate in absolute value for a 100 bps spread between LIBOR and r. This is economically significant, as 21 bps corresponds approximately to the average cross-country xccy rate in the post-crisis period.

In columns (3) and (4), we add those variables that are significant in explaining the dynamics of model-implied interest rates, the convenience-adjusted Treasury rates and the CDS premium. As we do not have CDS rates with a 3-month maturity, we use the 6-month rate instead. Neither of those two variables significantly changes the magnitude or the significance of the relation between xccy rates and $i-r$ spreads. In the specification in column (5), we further add currency fixed effects to soak up the average difference in cross-country xccy rates. Even in that case, we find a positive and statistically significant relation between xccy rates and $i - r$ spreads.

6.3 Stochastic discount factors

In this section we evaluate whether the variables that measure intermediary constraints co-move with the estimated SDFs. We consider two broad groups of variables. The first group uses the intermediary factors advocated in the literature. The motivation is that if a variable is significant in an excess return regression, it is implicitly associated with a linear SDF. Second, as a control, we consider various measures of uncertainty as many equilibrium SDFs depend on it, even in the absence of intermediary constraints.

Specifically, we measure intermediary constraints using the leverage of security broker-dealers (Adrian, Etula, and Muir, 2014, AEM) and the capital ratios of bank holding companies (He, Kelly, and Manela, 2017, HKM), the trade-weighted U.S. dollar index, which proxies for the limited willingness of intermediaries to provide USD funding and demand for USD associated with the convenience of USD assets (Avdjiev, Du, Koch, and Shin, 2019; Jiang, Krishnamurthy, and Lustig, 2018), and total dealers' cash balances with the Federal Reserve. We do not report results for the latter variable because it turns out that their coefficients are statistically insignificant.

Next, we consider the Jurado, Ludvigson, and Ng (2015) real uncertainty, macroeconomic uncertainty, and financial uncertainty measures; the Bekaert and Hoerova (2014) uncertainty and risk aversion measures; and the CBOE VIX index. We also examine country-specific measures of uncertainty, including the Baker, Bloom, and Davis (2016) economic policy uncertainty indices, implied volatility from 5x10-year swaptions, and xccy bid-ask spreads. As these country-specific measures are not significant in multivariate regressions, we do not report their results. Details on all data sources are available in the online appendix.

Table 6 summarizes the relation between SDFs and candidate variables by regressing the former on the latter at the monthly frequency. The AEM is an exception as it is available at the quarterly frequency only. We run panel regressions and examine the connection of SDFs to the individual variables, and to all of them together in a multivariate setting (with the exception of AEM). We add currency fixed effects to focus on the within variation at the exchange rate level.

We find that all the variables are significant individually. However, only four of them are jointly significant in multivariate regressions. On the intermediary side, we have the HKM intermediary capital ratio (ICR) and the U.S. dollar index (USD). On the uncertainty side, we have macroeconomic and financial uncertainty measures of Jurado, Ludvigson, and Ng (2015).

The signs of the regression coefficients for AEM and HKM are consistent with economic intuition. Marginal utility is high when dealer leverage is high or when intermediary capital is low. Relatedly, the expensiveness of the USD is associated with times of high marginal utility.

Usually asset-pricing theory associates periods of high uncertainty with a high price of risk (e.g., Bansal and Yaron, 2004). Thus, the empirical finding of a negative relation between the SDFs and uncertainty measures is puzzling. There is a group of models, however, that finds a positive univariate relation between the price of risk and consumption variance (e.g., Backus, Routledge, and Zin, 2010; Segal, Shaliastovich, and Yaron, 2015; Zviadadze, 2017). For instance, if periods with high consumption variance are associated with high expected future consumption growth, this relationship can arise.

Further, the results in column (11), which are for a shorter sample period that is determined by the availability of the USD factor, suggest opposite signs for macroeconomic and financial uncertainty, in line with the notions of 'good' and 'bad' uncertainty (Segal, Shaliastovich, and Yaron, 2015). The evidence is also consistent with Ludvigson, Ma, and Ng (2021) who argue that financial uncertainty causes a sharp decline in real activity whereas macro uncertainty is initially associated with increase in real activity in most estimations.

Haddad and Muir (2021) (HM) point out that a positive correlation between the SDF and the intermediary variables could simply indicate that risk-bearing capacity of households lines up with that of intermediaries. Thus, the evidence does not establish that the SDF of intermediaries is the one that prices xccy. HM's identification strategy is not applicable here as it relies on the cross section of assets with different degrees of intermediation. Here we have SDFs instead. Nevertheless, given that we have estimated these SDFs using interest rate swaps, which households cannot trade, it seems plausible to surmise that these are intermediary SDFs.

6.4 Xccy pricing errors

While our fit of xccy rates is reasonable, there are pricing errors. We interpret the xccy pricing errors as manifestation of LTA. As discussed previously, LTA may reflect potential components associated with intermediary constraints or other frictions that could be reflected in asset prices. This motivates us to investigate whether there is a covariation between xccy pricing errors and various measures of financial frictions that have been developed in the literature.

We consider three broad groups of variables. The first two are exactly the same as in the previous section. In the third group we use indicators of distress in the banking sector via the U.S. Treasury over Eurodollar (TED) spread, and the LIBOR-OIS spread. The latter is available for all countries. See online appendix for details.

Table 7 reports the results in the style of Table 6. Specifically, we examine the relation between changes in xccy pricing errors and changes in candidate variables by regressing the former on the latter at the monthly frequency (except for AEM, which is only available at the quarterly frequency). Many of the variables are individually significant. Only three variables remain significant in the multivariate regression: ICR, USD, and LIBOR-OIS spread. We report our final specification of the multivariate regression where we exclude the insignificant variables.

We check if the significant variables capture risk premiums by implementing cross-sectional regressions of expected changes in the xccy pricing errors on the beta exposures to the candidate risk factors. Specifically, we first estimate for each currency the exposure (i.e., betas) of changes in xccy pricing errors to changes in ICR and USD. We then relate the average xccy pricing errors to these betas and estimate, with some abuse of language, the "price of risk" associated with the pricing errors.

We plot the average realized changes in pricing errors against their predicted counterparts based on factor exposures and risk prices in Figure 6. The predicted pricing errors line up with xccy "returns" quite well. The estimated risk prices are statistically significant. Our cross-section is small, so the evidence is merely suggestive. The evidence is consistent with the view that our model omits important sources of risk premiums in these markets.

As we highlighted earlier, HM caution that a cross-sectional relation between excess returns and factor exposures to intermediary health may simply reflect high excess returns in times when dealers also happen to be constrained. They suggest overcoming this interpretation by focusing on a cross-section of asset classes. Evidence in favor of intermediary-based asset pricing is tied to a positive cross-sectional relation between the cost of intermediation for a given asset class and its exposure to intermediary risk aversion. Thus, in a last step, we conduct similar cross-sectional tests for the pricing errors of both 5-year xccy swap rates and 6-month forward premiums.

We regress changes in pricing errors on the HM intermediary risk aversion factor to estimate the exposure to intermediary risk. We then relate these beta exposures to the proportion of turnover that is intermediated through dealers in each corresponding market. In its 2019 triennial survey on OTC derivative products, the Bank for International Settlement reports, by currency, how much dealers account for the turnover in forward and swap markets, respectively.

The results in Figure 7 convey two messages. First, for all currencies, FX swaps are on average more intermediated through dealers than FX forwards. Second, there appears to be a positive link between the amount of dealer activity and exposure to intermediary risk aversion for xccy swaps, while that relation is much noisier for forward premiums. Remember that our fit of forward premiums was tight, and that we matched the broad pattern of xccy rates. However, our model is not able to eliminate all pricing errors for xccy.

Overall, this evidence further supports that both the NA and the LTA paradigms are interacting with each other. A common set of intermediary variables is associated with both. We next quantify the relative contribution of each channel.

6.5 Variance decomposition

Our objective is to better understand the relative contribution of the NA and LTA frameworks to the variation in xccy rates. We have established that intermediary variables are correlated with both the SDFs and the xccy pricing errors. That suggests that some amalgamation of exisiting frameworks is required for explaining CIP behaviour. Next, we resort to a simple variance decomposition to quantify how much of xccy rates is explained by NA versus LTA. We treat model-based xccy rates as associated with the fomer, and the model pricing errors as associated with the latter.

The pricing errors $xccy^e$ are defined as the difference between the observed, $xccy^d$, and model-implied, $xccy^m$, xccy rates. The fact that both $xccy^m$ and $xccy^e$ depend on ICR and on USD could complicate the variance decomposition. We exploit the fact that pricing errors are orthogonal to the model-based swap rates, by the model estimation design. Thus, even though both are related to the two intermediary factors, the specific linear combination of ICR and USD implied by our analysis of $xccy^e$ in Table 7 is not correlated with $xccy^m$, in population. In the online appendix, we explicitly factor out sample covariances between the relevant terms, and they are small.

According to column (12) in Table 7, changes in the xccy pricing errors are approximately a linear combination of changes in ICR and changes in the level of USD, with loadings of 180.49 and -0.44, respectively. We use these loadings to decompose the variance of xccy rates into contributions from $xccy^m$, ICR, and USD. We follow Campbell and Ammer (1993), and split (small) pairwise covariances between the respective factors equally.

We report in the first row of Table 8 the fraction of the variance of the observed xccy rate levels explained by the total variance derived from all components. On average the three factors capture about 95.21% of the observed xccy rate variance, with some variation across currencies. The last three rows in Table 7 provide a decomposition of the total variance into the respective components. The contribution to the total variance from the model-implied xccy rate is 68.33%, on average. The remaining 31.67% is, therefore, attributed to LTA. Thus, although NA plays a major role in understanding the behaviour of CIP, LTA contributes with a quantitatively important component as well.

7 Conclusion

In the era following the GFC, prices in exchange rate markets have exhibited patterns that are unusual from the perspective of classical textbook theories, and are, therefore, considered to be anomalies. CIP has been violated at both short and long horizons, as suggested by a non-zero basis and cross-currency basis swaps that have traded at non-zero prices. We examine the dynamics of the term structure of CIP violations across G11 currencies in a unifying way using a no-arbitrage framework.

First, we assume that true EFRs are unobserved and latent. Second, we assume the existence of a stochastic discount factor, implying that traded prices are consistent with no-arbitrage. Third, we assume that OTC derivatives transactions are fully collateralized.

Under these assumptions, we proceed and back out the true unobserved discount rates from plain vanilla interest rate swap contracts. We show that the discount rates implied from our model consistently price forward exchange rates and cross-currency basis swaps across all eleven currencies. Thus, true discount rates consistent with no-arbitrage can go a long way towards reconciling short-term and long-term CIP deviations. There remains one third of variation in xccy rates unexplained by the no-arbitrage model. We attribute this variation to limits to arbitrage.

We show that both variations in the estimated SDF and xccy pricing errors are related to measures of intermediaries' health and their willingness to provide USD funding. The evidence is consistent with exisiting theories of intermediary constraints.

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Figure 1: **CIP deviations for the Euro.** We display the log three-month LIBOR basis, defined as the difference between the forward-spot exchange rate premium and the LIBOR interest rate differential in the corresponding currencies, $f - s - (i^{\$} - i^{€})$, and the 5-year cross-currency basis swap rate for the Euro vs. the U.S. dollar. The swap exchanges interest payments reflecting LIBOR rates in the two countries. The swap rate is quoted as the spread over the EURIBOR-based interest payments. The sample period is January 2000 to December 2019. Source: Bloomberg.

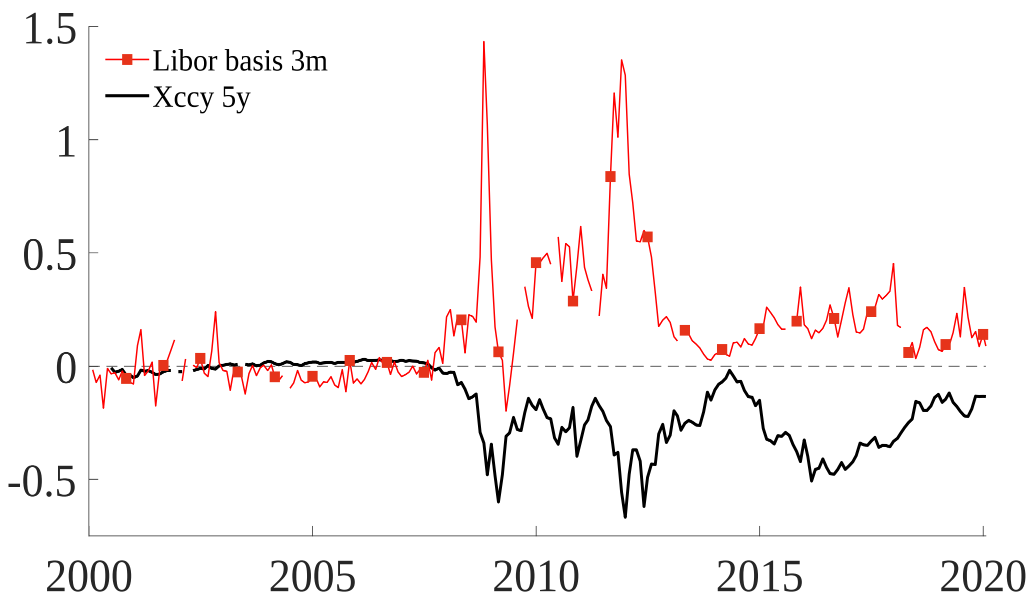
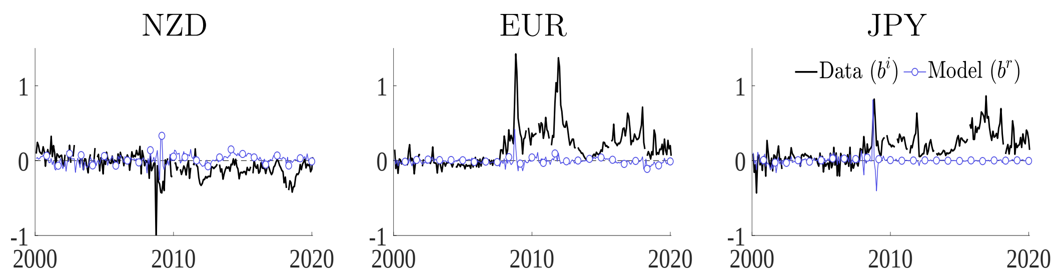
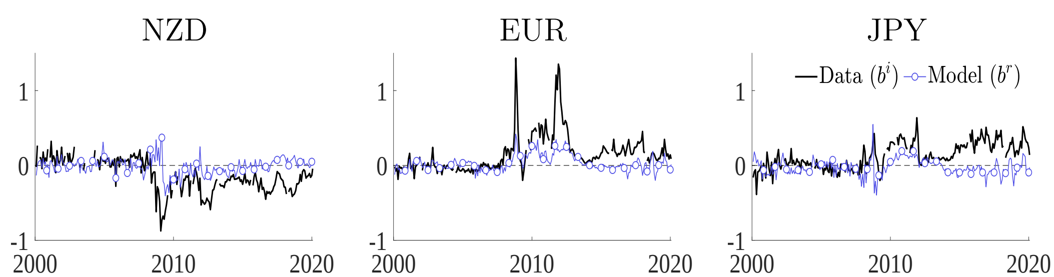


Figure 2: **Time-series of forward basis and xccy for NZD, EUR, and JPY.** In these figures, we report the time series of the forward basis (3 and 6 months, based on LIBOR in the data and on EFR in the model) or xccy rates (5 and 20 years) implied from the model and compare it with the data. The sample period is January 2000 to December 2019. Source: Bloomberg. Similar results for other G11 currencies are provided in the online appendix.

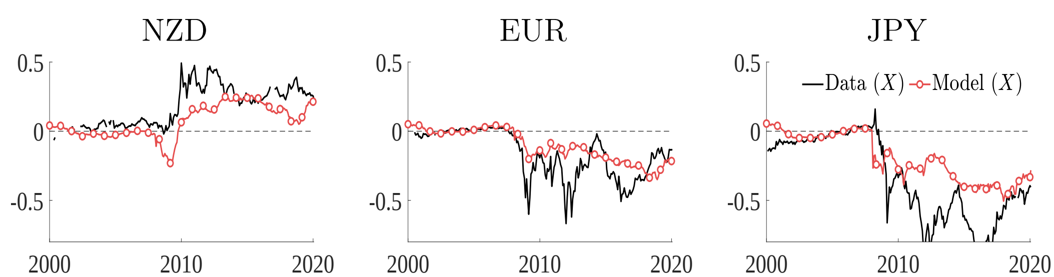
(A) Forward basis 3-month maturity



(B) Forward basis 6-month maturity



(C) xccy 5-year maturity



(D) xccy 20-year maturity

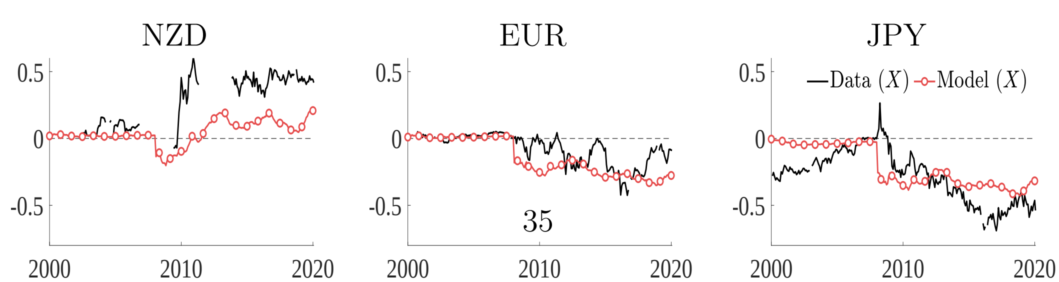
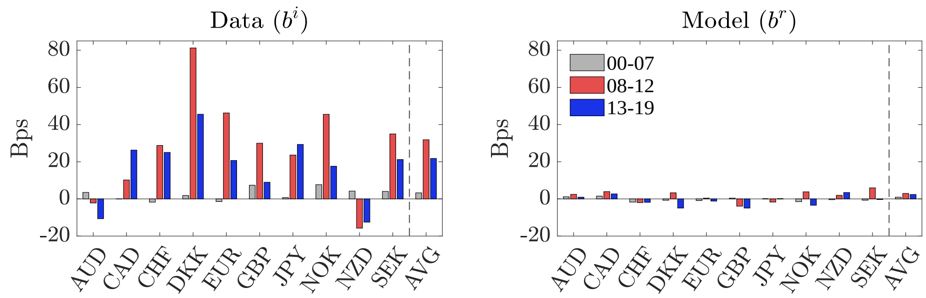
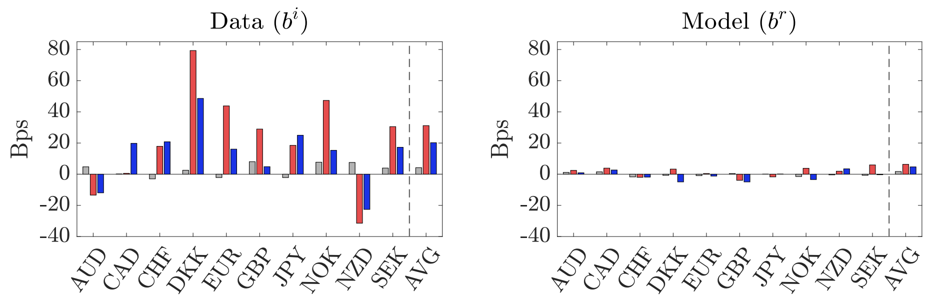


Figure 3: **Forward basis and xccy rate.** We report the mean of the forward basis (based on LIBOR in the data and on EFR in the model) and the cross-currency basis swap rate (in bps). We also report the cross-sectional average of absolute rates, AVG. All exchange rates are expressed as the USD price per unit of foreign currency. We report statistics for the G10 currencies. The countries and currencies are denoted by their usual abbreviations: Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), Danish krone (DKK), Euro (EUR), British pound (GBP), Japanese yen (JPY), Norwegian krone (NOK), New Zealand dollar (NZD), and Swedish krona (SEK). The sample period is January 2000 to December 2019. Source: Bloomberg. Tables with supporting numbers are provided in the online appendix.

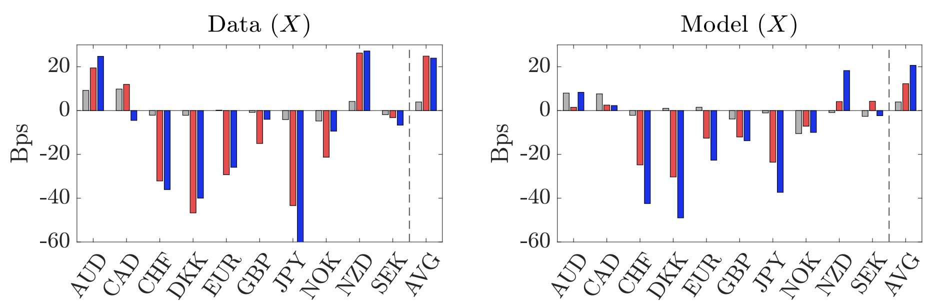
(A) Forward basis 3-month maturity



(B) Forward basis 6-month maturity



(C) xccy 5-year maturity



(D) xccy 20-year maturity

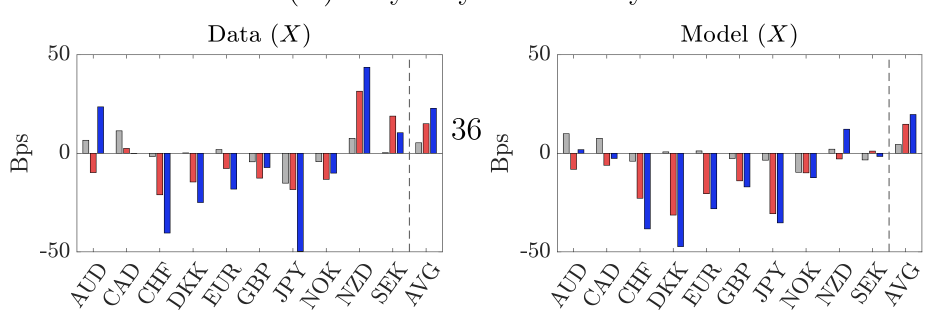


Figure 4: **Comparison of 1Y Interest Rate Proxies.** Each figure compares the model-implied 1-year interest rate to the predicted one and given by $\Delta r = -0.01 + 0.17 \cdot \Delta\text{LIBOR} + 0.40 \cdot \Delta(\text{Treasury} + \lambda) - 0.24 \cdot \Delta\text{CDS}$, where LIBOR corresponds to the country-specific Libor/interbank rate, Treasury corresponds to the country-specific Treasury rate, and λ refers to the country-specific convenience yield, computed as the Treasury basis plus the U.S. Refcorp-Treasury spread, and CDS corresponds to the country-specific 1-year local currency denominated CDS premium (we use the USD denomination if the local currency CDS is not available). We use G11 currencies, i.e., USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. We use Libor rates for USD, JPY, GBP, CHF, Cdor rates for Canada, Euribor rates for EUR, BBSW rates for AUD, BKBM rates for NZD, Stibor rates for SEK, Cibor rates for DKK, Nibor rates for NOK. The sample period is January 2000 to December 2019. Source: Bloomberg.

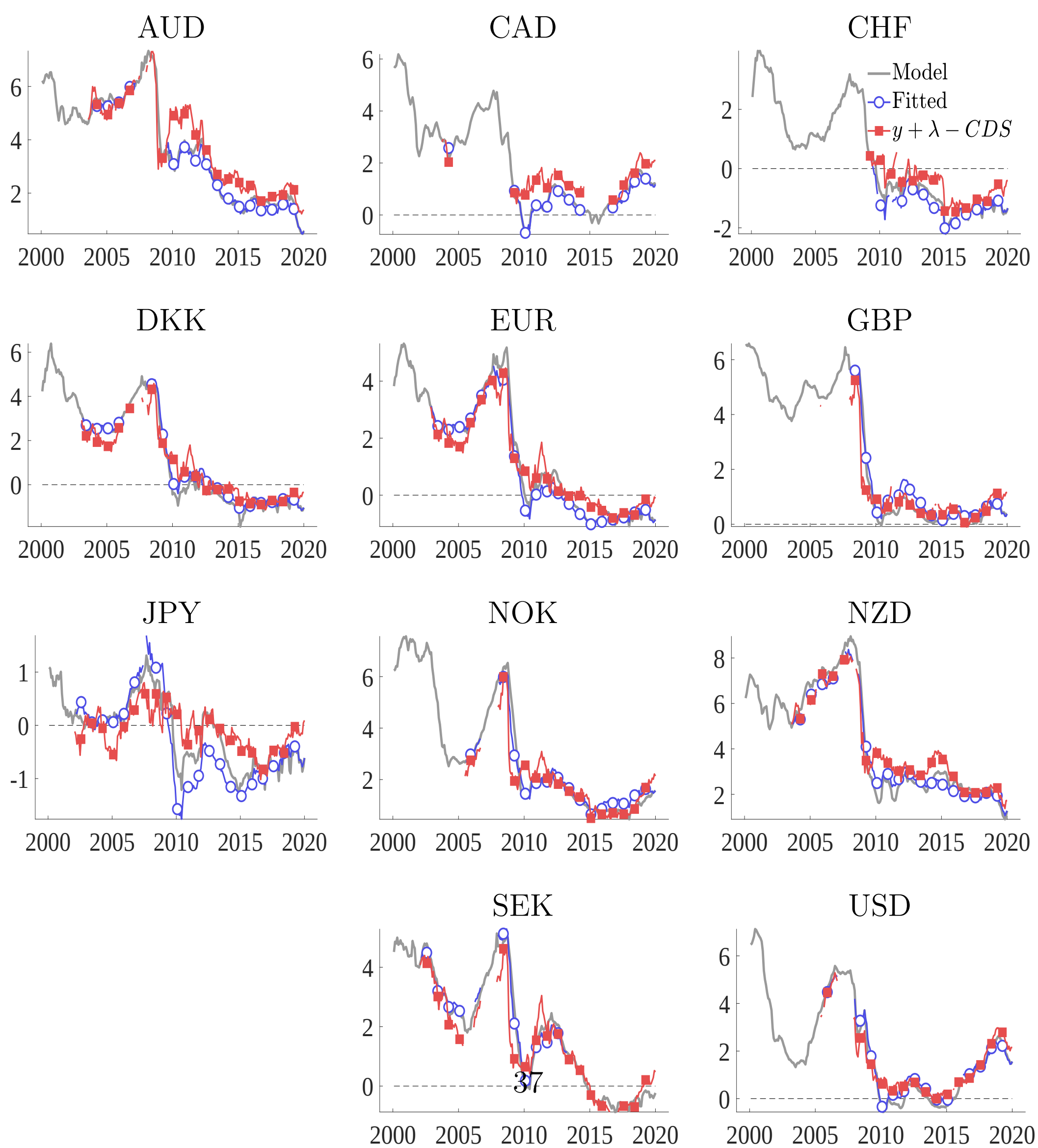


Figure 5: **Comparison of 3M Interest Rate Proxies.** Each figure compares the model-implied 3-month interest rate to the predicted one and given by $\Delta r = -0.01 + 0.31 \cdot \Delta\text{LIBOR}+0.38 \cdot \Delta(\text{Treasury}+\lambda)$, where LIBOR corresponds to the country-specific Libor/interbank rate, Treasury corresponds to the country-specific Treasury rate, and λ refers to the country-specific convenience yield, computed as the Treasury basis plus the U.S. Refcorp-Treasury spread. We use G11 currencies, i.e., USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. We use Libor rates for USD, JPY, GBP, CHF, Cdor rates for Canada, Euribor rates for EUR, BBSW rates for AUD, BKBM rates for NZD, Stibor rates for SEK, Cibor rates for DKK, Nibor rates for NOK. The sample period is January 2000 to December 2019. Source: Bloomberg.

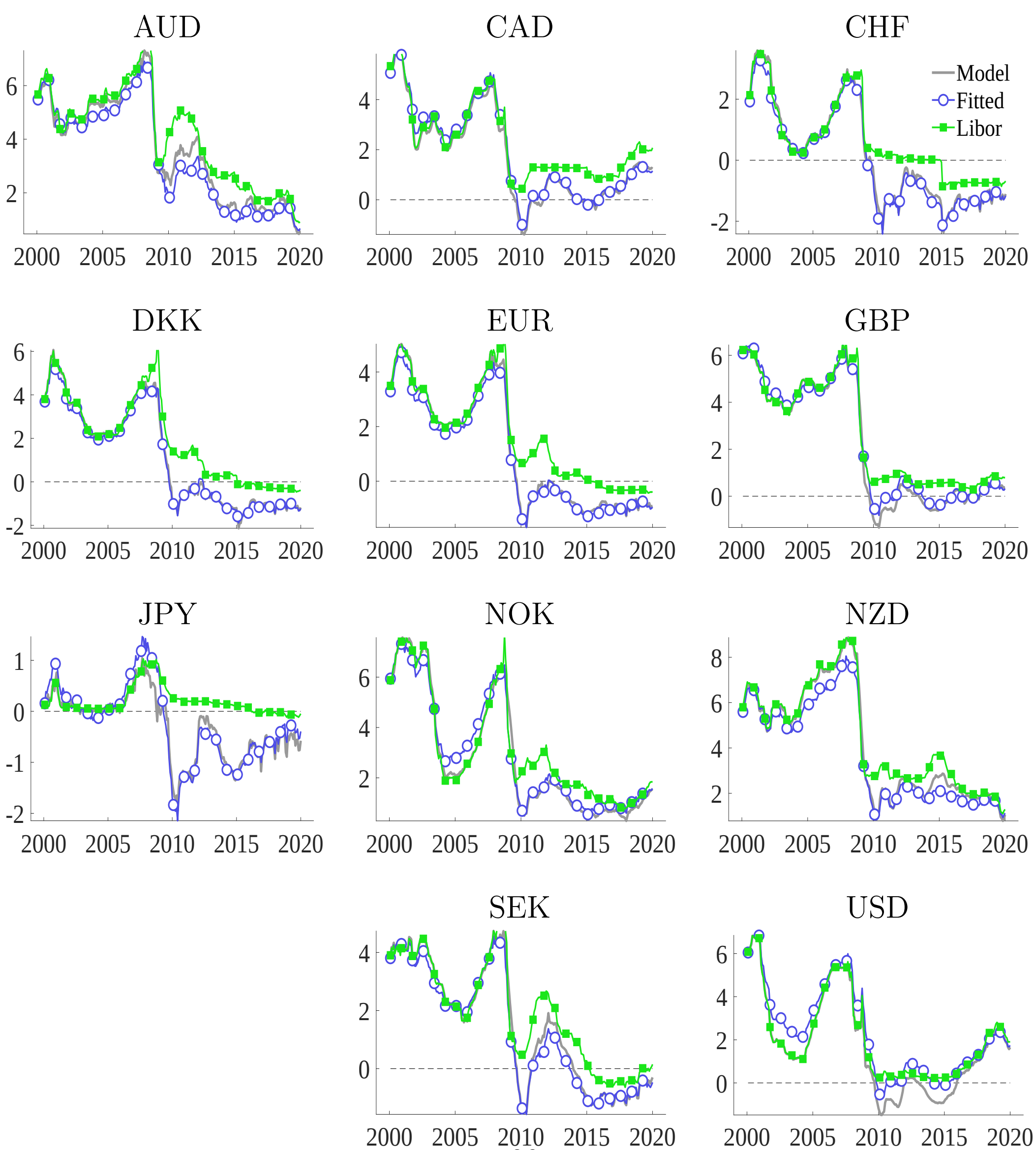
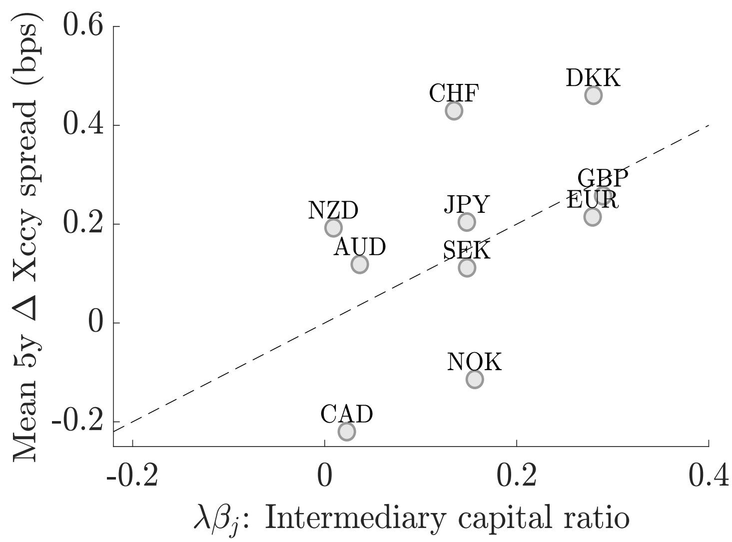


Figure 6: **Factor Exposure of Xccy Basis Swap Spread Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied 5-year xccy basis swap rate on a risk factor, i.e., $\Delta xccy_t^e = \alpha + \beta \cdot RF_t + \varepsilon_t$. We then project the average level of the xccy spread on the estimated betas $\widehat{\beta}$. We use two risk factors: changes in the He, Kelly, and Manela (2017) intermediary capital ratio (Δ HKM-ICR); changes in the trade-weighted U.S. dollar index (Δ USD Factor). The sample period is January 2000 to December 2019.



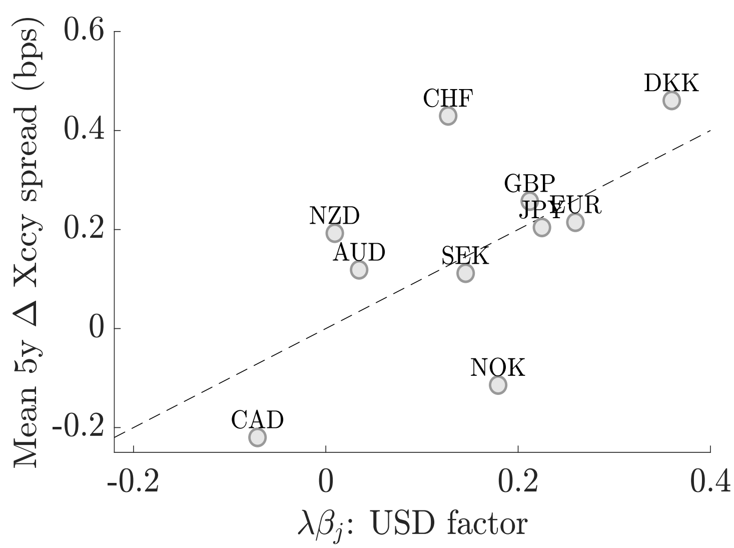


Figure 7: **Factor Exposure of Xccy Basis Swap Spread and Forward Premium Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied (i) 5-year xccy basis swap rate and (ii) the 6-month forward premium on the Haddad and Muir (2021) intermediary risk aversion factor, i.e., $\Delta xccy^e_{t+1} = \alpha + \beta \cdot RF_t + \varepsilon_t$. We then project the estimated raw betas $\hat{\beta}$ on the fraction of foreign exchange turnover accounted for by intermediaries. In its 2019 triennial Central Bank survey on foreign exchange turnover, the BIS provides information on the fraction of turnover acccounted for by intermediaries for FX forwards and FX swaps, respectively. The sample period is 2000Q1 to 2017Q3.

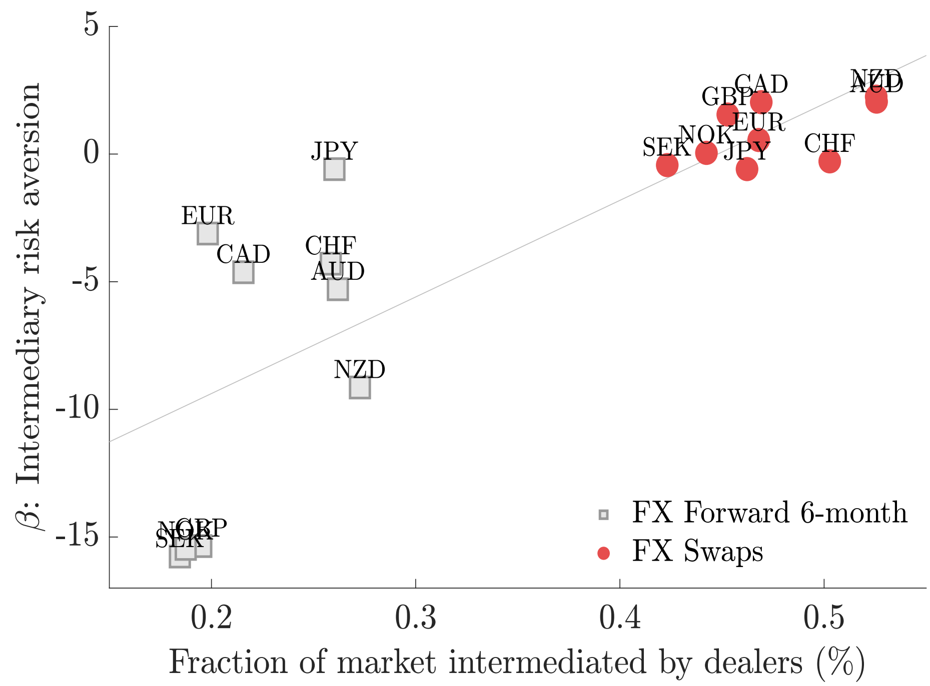


Table 1: The Role of Frictions and Constraints for CIP Violations

This table summarizes a set of papers that consider how frictions or constraints may contribute to violations of covered interest rate parity. These frictions and constraints can be common (intermediary capital, intermediary risk aversion, intermediary leverage, regulatory capital ratio, funding liquidity) or asset-specific (margins, collateral, transaction costs, idiosyncratic risk, asset liquidity). For each reference, we indicate whether it is primary theoretical, empirical, or both. We also highlight whether the frictions impact asset prices through their impact on the stochastic discount factor, and therefore don't imply arbitrage violations, or whether they imply limits to arbitrage.

	Analysis		Frictions/Constraints										Arb. Viol.	
Study	Theory	Empirics	Intermediary Capital	Intermediary Risk Aversion	Intermediary Leverage	Regulatory Capital Ratio	Funding Liquidity	Margin	Collateral	Transaction Costs	Idiosyncratic Risk	Asset Liquidity	Yes	No
Dow and Gorton (1994)	✓								✓				✓	
Pontiff (1996)		✓									✓		✓	
Shleifer and Vishny (1997)	✓						✓						✓	
Liu and Longstaff (2004)	✓								✓				✓	
Gabaix, Krishnamurthy, and Vigneron (2007)	✓	✓					✓						✓	
Garleanu and Pedersen (2011)	✓							✓					✓	
Acharya, Lochstoer, and Ramodarai (2013)	✓	✓					✓						✓	
He and Krishnamurthy (2013)	✓		✓			✓								✓
Brunnermeier and Sannikov (2014)	✓				✓							✓		✓
Adrian, Etula, and Muir (2014)		✓			✓									✓
He, Kelly, and Manela (2017)		✓	✓											✓
Haddad and Muir (2021)	✓	✓		✓										✓
Du, Tepper, and Verdelhan (2018)		✓				✓							✓	
Jermann (2020)	✓	✓								✓			✓	

Table 2: Cash flows from a plain vanilla and fixed-for-fixed cross-currency basis swap

Panel A in this table illustrates the cash flows generated by a stylized cross-currency basis swap that receives the floating interest rate of i_t on the USD leg at each date $t+1$, and pays the floating interest rates $\widehat{i}_t + X$ on the EUR leg at each date $t+1$. The price of the cross-currency basis swap is given by X. S indicates the USD value per unit of foreign currency. Panel B transforms the plain vanilla cross-currency basis swap into a stylized fixed-for-fixed cross-currency basis swap, constructed as a package of a standard cross-currency basis swap that receives the floating interest rate of i_t on the USD leg at each date $t+1$, and pays the floating interest rates $\widehat{i}_t + X$ on the EUR leg at each date $t+1$. The notional face values of the domestic and foreign legs are matched using the spot exchange rate S_0, where S indicates the USD value per unit of foreign currency. The floating payments in each currency are converted into fixed payments using plain vanilla interest rate swaps at prices CMS and $\widehat{CMS}$ respectively.

		$S = \$1/€1$		Cash flows at time	
		XC Basis Swap	0	t	T
Panel A	XC Swap	EUR Leg	$+\,€1$	$-\,€\left(\widehat{i}_{t-1} + X\right)$	$-\,€\left(\widehat{i}_{T-1} + X\right) - €1$
		USD Leg	$-\,\$S_0$	$+\,\$S_0 i_{t-1}$	$+\,\$S_0 i_{T-1} + \$S_0$
Panel B	€ IRS	Fix Leg		$-\,€\widehat{CMS}$	$-\,€\widehat{CMS}$
		Float Leg		$+\,€\widehat{i}_{t-1}$	$+\,€\widehat{i}_{T-1}$
	\$ IRS	Float Leg		$-\,\$S_0 i_{t-1}$	$-\,\$S_0 i_{T-1}$
		Fix Leg		$+\,\$S_0 CMS$	$+\,\$S_0 CMS$

Table 3: Factor Structure in International Anomalies - By Currency

This table reports the results from a principal component analysis (PCA). We report the cumulative proportion of variance explained by the five first principal components (PC1 to PC5). We use G11 currencies, i.e., USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. In Panel A, we focus on the term structure of cross-currency basis swaps using maturities of 1y, 3y, 5y, 7y, 10y, 15y, 30y, except for NZD, which omits 30y. In Panel B, we examine the factor structure across all interbank (LIBOR) and IRS rates. For the former we use maturities of 1m, 3m, 6m, and 1y, except for NOK, which omits 1y. We use Libor rates for USD, JPY, GBP, CHF, Cdor rates for Canada, Euribor rates for EUR, BBSW rates for AUD, BKBM rates for NZD, Stibor rates for SEK, Cibor rates for DKK, Nibor rates for NOK. For the latter we use maturities of 1y, 3y, 5y, 7y, 10y, 15y, 30y. The sample period is January 2000 to December 2019. Source: Bloomberg

(A) XCCY	USD	JPY	GBP	CAD	EUR	AUD	CHF	NZD	SEK	DKK	NOK
PC1	–	87.78	79.40	72.07	87.43	81.56	89.60	93.72	69.45	88.44	88.49
PC2	–	96.30	95.94	91.13	98.01	95.17	97.74	98.12	95.13	98.29	96.45
PC3	–	99.71	98.56	97.93	99.37	98.88	99.44	99.72	98.62	99.44	99.07
PC4	–	99.95	99.51	99.35	99.89	99.61	99.87	99.86	99.70	99.79	99.50
PC5	–	99.99	99.89	99.76	99.97	99.88	99.94	99.95	99.84	99.90	99.72
(B) LIBOR+IRS	USD	JPY	GBP	CAD	EUR	AUD	CHF	NZD	SEK	DKK	NOK
PC1	88.89	79.30	94.24	87.65	96.26	96.31	94.52	88.85	82.17	95.22	93.03
PC2	99.31	96.41	99.45	99.24	99.62	99.52	99.53	99.60	98.84	99.60	99.30
PC3	99.81	98.44	99.84	99.78	99.87	99.82	99.83	99.83	99.72	99.84	99.83
PC4	99.93	99.66	99.92	99.91	99.96	99.91	99.92	99.90	99.88	99.95	99.93
PC5	99.98	99.83	99.98	99.96	99.99	99.97	99.98	99.95	99.96	99.98	99.97

Table 4: Model-implied Interest Rates and Candidate Proxies

In this table, we report results from the panel regressions where we project changes in the model-implied interest rates on changes in proxy candidates at matching maturities. At the country level, we use the Treasury yield, the OIS rate, the interbank rate (LIBOR), the Treasury convenience yield λ (computed as the Treasury basis plus the U.S. Refcorp-Treasury spread), the sum of Treasury and convenience yield, the IRS rate, the CDS premium, the difference between the Treasury yield and CDS premium (Treasury-CDS), a linear combination of the Treasury yield, convenience yield, CDS premium (Treasury+λ-CDS). We also use common variables, namely the effective federal funds rate (EFFR), the certificate of deposit rate over Treasury yield spread (CD-Treasury), and the option-implied box spread (BOX). We use either the 3-month or the 1-year maturity. The data frequency is monthly based on the last available monthly information. Standard errors are robust and adjusted for heteroscedasticity. All regressions contain currency fixed effects and time fixed effects, and we report the within and adjusted R^2 values from the panel regressions. We use the G11 currencies: USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. For NOK, we lack data on OIS rates and 1y NIBOR rates. The sample period is January 2000 to December 2019.

VARIABLES	(1) Δr	(2) Δr	(3) Δr	(4) Δr	(5) Δr	(6) Δr	(7) Δr	(8) Δr	(9) Δr	(10) Δr	(11) Δr	(12) Δr	(13) Δr
Treasury	0.38***												
	(0.04)												
OIS		0.50***											
		(0.11)											
LIBOR			0.65***									0.31***	0.17**
			(0.05)									(0.06)	(0.07)
λ				0.25***									
				(0.03)									
Treasury+λ					0.57***							0.38***	0.40***
					(0.03)							(0.04)	(0.05)
EFFR						0.14***							
						(0.03)							
CD-Treasury							0.13***						
							(0.04)						
CDS								0.03					-0.24**
								(0.09)					(0.10)
Treasury-CDS									0.31***				
									(0.04)				
Treasury+λ-CDS										0.47***			
										(0.05)			
BOX											0.19***		
											(0.06)		
Constant	-0.01***	-0.01***	-0.01***	-0.02***	-0.01***	-0.02***	-0.02***	-0.02***	-0.02***	-0.02***	-0.02***	-0.01***	-0.01***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
OBSERVATIONS	2,640	2,024	2,640	2,640	2,640	2,640	2,640	1,812	1,812	1,812	1,870	2,640	1,550
CCY Groups	11	10	11	11	11	11	11	11	11	11	11	11	10
CCY FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
MONTH FE	YES	YES	YES	YES	YES	NO	NO	YES	YES	YES	NO	YES	YES
MAT	3M	3M	3M	3M	3M	3M	3M	1Y	1Y	1Y	1Y	3M	1Y
w.R^2	0.218	0.261	0.466	0.089	0.527	0.049	0.040	-0.001	0.143	0.333	0.051	0.575	0.391
adj.R^2	0.607	0.646	0.732	0.543	0.763	0.048	0.040	0.495	0.568	0.663	0.051	0.787	0.679

Table 5: Xccy Rates and Spreads between LIBOR and Model-Implied Interest Rates

In this table, we report results from the panel regressions where we project the absolute values of the observed 5-year xccy rates ($|xccy5y^D|$) on the spread between LIBOR and model-implied interest rates at the 3-month maturity ($3\text{M}(i-r)$). At the country level, we control for the Treasury yield adjusted for the convenience premium (Treasury $+\lambda$), and the CDS premium. The data frequency is monthly based on the last available monthly information. Standard errors are clustered by month. We indicate whether regressions contain currency or monthly time fixed effects, and we report the adjusted R^2 values from the panel regressions. We use the G11 currencies except for the USD: JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. For NOK, we lack data on OIS rates and 1y NIBOR rates. The sample period is January 2008 to December 2019.

VARIABLES	(1) $\lvert xccy5y^D \rvert$	(2) $\lvert xccy5y^D \rvert$	(3) $\lvert xccy5y^D \rvert$	(4) $\lvert xccy5y^D \rvert$	(5) $\lvert xccy5y^D \rvert$
$3\text{M}(i-r)$	4.75***	21.10***	17.89***	18.66***	10.08***
	(0.85)	(1.70)	(1.54)	(1.70)	(2.65)
ΔTreasury+λ			-1.65***	-2.04***	2.16***
			(0.31)	(0.33)	(0.80)
CDS				10.70	7.90
				(6.51)	(4.98)
Constant	20.81***	7.84***	12.33***	12.66***	14.82***
	(0.81)	(1.35)	(1.38)	(1.66)	(2.33)
OBSERVATIONS	1,405	1,405	1,405	1,269	1,269
CCY GROUPS	10	10	10	10	10
CCY FE	NO	NO	NO	NO	YES
MONTH FE	NO	YES	YES	YES	YES
adj.R^2	0.017	0.139	0.151	0.174	0.721

Table 6: Relation of model-implied stochastic discount factors to frictions/constraints

In this table, we report results from the panel regressions where we project the level of the model-implied log SDFs on common risk factors. We use the following common factors: the Adrian, Etula, and Muir (2014) dealer leverage factor (AEM-LV2); the He, Kelly, and Manela (2017) intermediary capital ratio factor (HKM-ICR); the trade-weighted U.S. dollar index (USD Factor); the Jurado, Ludvigson, and Ng (2015) real uncertainty (JNL-RU12), macroeconomic uncertainty (JNL-MU12), and financial uncertainty (JNL-FU12); the Bekaert-Horeova uncertainty (BH-UC) and risk aversion (BH-RA) measures; the CBOE VIX index (VIX). The data frequency is monthly based on the last available monthly information, except for the regression in column (1), which uses quarterly data. Standard errors are robust and adjusted for heteroscedasticity. All regressions contain currency fixed effects, specifications with currency-specific variables contain time fixed effects, and we report the within and adjusted R^2 values from the panel regressions. We use the G11 currencies excluding the USD: JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. The sample period is January 2000 to December 2019.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
VARIABLES	m	m	m	m	m	m	m	m	m	m	m
AEM-LV2	0.08***										
	(0.03)										
HKM-ICR		-8.49**								-24.70***	-17.93**
		(4.09)								(2.63)	(7.99)
USD FACTOR			0.10***								0.08***
			(0.01)								(0.01)
JNL-RU12				-69.50***							
				(4.79)							
JNL-MU12					-38.74***					-37.65***	-44.46***
					(1.78)					(1.81)	(2.25)
JNL-FU12						-27.49***				-3.54***	9.66***
						(1.47)				(1.05)	(2.75)
BH-UC							-0.06***				
							(0.00)				
BH-RA								-0.09***			
								(0.01)			
VIX									-0.18***		
									(0.01)		
Constant	-2.75***	-2.05***	-11.99***	57.96***	32.77***	24.45***	-1.28***	-1.03***	0.94***	36.91***	21.89***
	(0.12)	(0.32)	(0.63)	(4.15)	(1.60)	(1.42)	(0.09)	(0.11)	(0.21)	(1.51)	(2.28)
OBSERVATIONS	772	2,463	1,813	2,602	2,602	2,602	2,591	2,591	2,602	2,463	1,674
CCY FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
MONTH FE	NO	NO	NO	NO	NO	NO	NO	NO	NO	NO	NO
$w.R^2$	0.019	0.003	0.076	0.194	0.375	0.197	0.166	0.146	0.180	0.399	0.487
$adj.R^2$	0.132	0.130	0.171	0.294	0.452	0.296	0.269	0.251	0.281	0.475	0.542

Table 7: Spread between Model-implied and Observed 5y-Xccy Basis Swap Rates

In this table, we report results from the panel regressions where we project changes in the spread between the model-implied and the observed 5-year xccy basis swap rates ($\Delta xccy^e$) on changes in proxy candidates for explanatory variables. We use the following common factors: the Adrian, Etula, and Muir (2014) dealer leverage factor (AEM-LV2); the He, Kelly, and Manela (2017) intermediary capital ratio factor (HKM-ICR); the trade-weighted U.S. dollar index (USD Factor); the Jurado, Ludvigson, and Ng (2015) real uncertainty (JNL-RU12), macroeconomic uncertainty (JNL-MU12), and financial uncertainty (JNL-FU12); the Bekaert-Horeova uncertainty (BH-UC) and risk aversion (BH-RA) measures; the CBOE VIX index (VIX); the Ted rate (TED). At the country level, we use the Libor-Ois spreads. All tenors are 5 year contracts. The data frequency is monthly based on the last available monthly information, except for the regression in column (1), which uses quarterly data. Standard errors are robust and adjusted for heteroscedasticity. All regressions contain currency fixed effects and column (11) contains time fixed effects, and we report the within and adjusted R^2 values from the panel regressions. We use the G11 currencies excluding the USD: JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. The sample period is January 2000 to December 2019.

VARIABLES	(1) $\Delta xccy^e$	(2) $\Delta xccy^e$	(3) $\Delta xccy^e$	(4) $\Delta xccy^e$	(5) $\Delta xccy^e$	(6) $\Delta xccy^e$	(7) $\Delta xccy^e$	(8) $\Delta xccy^e$	(9) $\Delta xccy^e$	(10) $\Delta xccy^e$	(11) $\Delta xccy^e$	(12) $\Delta xccy^e$
AEM-LV2	-0.01											
	(0.08)											
HKM-ICR		137.98***										180.49***
		(24.92)										(54.24)
USD FACTOR			-0.64***									-0.44***
			(0.10)									(0.12)
JNL-RU12				-115.03***								
				(40.25)								
JNL-FMU12					-61.75***							
					(21.16)							
JNL-FU12						-39.92**						
						(18.39)						
BH-UC							-0.05***					
							(0.01)					
BH-RA								-0.03**				
								(0.01)				
VIX									-0.13***			
									(0.03)			
TED										-1.77***		
										(0.66)		
LIBOR-OIS											-4.67*	-7.02***
											(2.61)	(2.65)
Constant	-0.01	0.12	0.11	0.06	0.05	0.04	0.05	0.06	0.05	0.06	0.04	0.16
	(0.26)	(0.09)	(0.12)	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)	(0.10)	(0.14)
OBSERVATIONS	691	2,228	1,635	2,354	2,354	2,354	2,344	2,344	2,354	2,354	1,773	1,301
CCY Groups	10	10	10	10	10	10	10	10	10	10	10	10
CCY FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
MONTH FE	NO	NO	NO	NO	NO	NO	NO	NO	NO	NO	YES	NO
MAT	5Y	5Y	5Y	5Y	5Y	5Y	5Y	5Y	5Y	5Y	5Y	5Y
w.R^2	-0.001	0.015	0.047	0.009	0.009	0.003	0.021	0.005	0.016	0.008	0.010	0.084
adj.R^2	-0.014	0.012	0.042	0.006	0.006	-0.000	0.018	0.001	0.012	0.004	0.227	0.079

Table 8: Model-implied Variance Decomposition

This table reports the model-implied variance decomposition for the levels of 5-year xccy swaps. Define $xccy^d$ to be the observed 5-year xccy rate in the data, $xccy^m$ to be the 5-year xccy rate implied by the no-arbitrage model, ICR to be the He, Kelly, and Manela intermediary capital ratio, USD to be the USD factor in levels, and $var(\cdot)$ refers to their variances. According to column (12) in Table 7, we have $\Delta xccy^d \approx \Delta xccy^m + 180.49 \cdot \Delta ICR - 0.44 \cdot \Delta USD$. We use this expression to infer a variance decomposition for the level of 5-year xccy rates. We equally split pairwise covariances between the corresponding factors. All ratios are reported in %. We report the average ratios across currencies and all ratios at the currency level. We use G11 currencies excluding USD, i.e., JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. The sample period is January 2000 to December 2019.

	MEAN	JPY	GBP	CAD	EUR	AUD	CHF	NZD	SEK	DKK	NOK
$var(xccy^m, ICR, USD)/var(xccy^d)$	95.21	55.25	91.59	109.53	90.50	46.41	146.37	43.76	171.11	119.24	78.37
$var(xccy^m)/var(xccy^m, ICR, USD)$	68.33	79.78	57.34	55.62	71.28	48.48	79.91	100.50	53.89	83.10	53.36
$var(ICR)/var(xccy^m, ICR, USD)$	11.40	9.55	18.07	14.73	11.90	39.79	8.52	-2.14	6.22	7.66	-0.32
$var(USD)/var(xccy^m, ICR, USD)$	20.28	10.67	24.58	29.64	16.82	11.73	11.56	1.65	39.89	9.23	46.96