

The term structure of covered interest rate parity 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 accounts for about two thirds of the alleged CIP deviations. The residual pricing errors are associated with intermediary constraints.

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1 Introduction

An important literature provides evidence of persistent violations of covered interest parity (CIP) since the 2008 Global Financial Crisis (GFC) (e.g., [Ivashina, Scharfstein, 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.¹ Both are computed for the Euro to U.S. Dollar (USD) exchange rate. Departures from zero for both time series are taken as prima facie evidence that CIP is violated. In turn, CIP violation is equivalent to the failure of no-arbitrage (NA).

In this paper we present a new quantitative benchmark for CIP violations. We establish, using this benchmark, that departures from CIP are not as large as one might think on the basis of the presented evidence. Two main reasons drive this conclusion.

First, it is important to account for effective funding rates (EFRs) in over-the-counter markets when counterparty risk is effectively removed via collateralization. Using LIBOR to compute the cross-currency basis may not be appropriate because it reflects counterparty default risk ([Du, Im, and Schreger, 2018](#)).² More generally, funding heterogeneity makes it difficult to use a single observable measure of costs ([Rime, Schrimpf, and Syrstad, 2019](#)).

Second, xccy contracts feature recurring cash flows that are contractually linked to LIBOR. Therefore, both the basis and xccy rates are simultaneously zero only when the EFR coincides with LIBOR. Otherwise, xccy rates may differ from zero even when the basis is zero. This perspective is consistent with the view that LIBOR was a good proxy for the EFR before the GFC when both the basis and xccy rates were close to zero (e.g., [Feldhutter and Lando, 2008](#), [Hull and White, 2013](#)).

To implement the new benchmark, we depart from prior work that uses observable proxies for EFRs and treat them as unobserved. We turn to another market to infer EFRs. Specifically, we ask how an interest rate swap (IRS) dealer would value currency derivatives. That is a natural benchmark because IRS are, by NA, closely related contracts, are also traded over-the-counter among a similar set of market participants, and have similar collateralization requirements and cash flows that are

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

²OIS rates, which also represent uncollateralized interbank lending, are not appropriate either because they may reflect market segmentation due to unequal access to interest on excess reserves paid by the Federal Reserve ([Bech and Klee, 2011](#)). Likewise, collateralized funding rates like term repos and other money market rates are not reliably available for horizons beyond 3 months.

linked to international interest rates. Moreover, the IRS market is highly liquid with a weekly traded notional of one trillion USD.

We use a workhorse affine no-arbitrage model to estimate the EFRs from IRS in each country for horizons up to 30 years. We compare the inferred EFRs to various plausible proxies for funding rates and funding frictions and show that they are a linear combination of LIBOR, Treasury rates adjusted for convenience, and sovereign credit default swap (CDS) rates. Our first empirical contribution is to show that, using these rates, the cross-currency basis is zero at horizons up to one year.

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 denominated in different currencies, typically LIBOR. As such, xccy are linked to portfolios of IRS by NA. 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. The zero xccy rate benchmark is implicitly associated, however, with LIBOR being the correct discount rate.

Our second empirical contribution is to assess if xccy rates valued using EFRs extracted from IRS match their data counterparts. We find that the average 5-year xccy rate across countries during the post-crisis period is 21 basis points (bps) in the model vs. 24 bps in the data. Our variance decomposition of the 5-year xccy rates shows that our NA framework explains, on average, about 68% of the variation in their levels.

These findings leave us with a question of what explains xccy pricing errors. It is natural to link them to features of derivatives markets and the relevant intermediaries that are unaccounted for by our model. Such features may suggest some misspecification of our NA model, or imply departures from NA outright. Both interpretations point towards various intermediary constraints as the missing pieces in our account. Indeed, extant theory implies that intermediary constraints could operate both within the NA environment (e.g., [Brunnermeier and Sannikov, 2014](#), [He and Krishnamurthy, 2013](#)) or outside of it, consistent with the limits to arbitrage (LTA) paradigm (e.g., [Garleanu and Pedersen, 2011](#), [Gromb and Vayanos, 2010](#), [Shleifer and Vishny, 1997](#)).

Specifically, our model accounts for collateralization via contractual cash flows and appropriate discounting, but it ignores limited availability of collateral. It also ignores other intermediary constraints that are consistent with NA. Of course, swap pricing errors could be reflective of LTA as well. For instance, in contrast to IRS, xccy are not centrally cleared. Thus, similar cash flows might be priced differently across the two markets due to different margin requirements. Moreover, the IRS market is an order of magnitude larger than the xccy market, and likely more competitive. Thus,

dealers could extract larger rents when trading xccy (Wallen, 2020). More generally, we do not account for constraints leading to violations of the law of one price.

Rather than attempting a detailed account of the institutional features of currency derivatives markets, we provide an empirical assessment of the determinants of the uncovered pricing errors. Our third empirical contribution is thus to evaluate whether the xccy pricing errors relate to measures of intermediary factors constraints proposed in the literature. We find that the leverage of bank holding companies (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 co-varying 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 provide further evidence for the relevance of intermediaries for pricing errors with an analysis along the lines of Haddad and Muir (2021). They establish the importance of intermediaries in pricing by showing that returns of assets that are more intermediated are more sensitive to intermediary risk aversion. We show that swaps are intermediated more than forward contracts, and that their risk aversion sensitivities line up accordingly. The overall evidence is consistent with intermediaries' constraints playing a role in the determination of xccy premiums (He and Krishnamurthy, 2013; Du, Tepper, and Verdelhan, 2018; Andersen, Duffie, and Song, 2019; Fleckenstein and Longstaff, 2020). That evidence is also consistent with the view in Du, Hebert, and Huber (2019) that CIP deviations are a priced source of risk that reflects the shadow cost of intermediary constraints.

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. 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 LIBOR-based cross-currency bases and xccy rates do not necessarily imply arbitrage opportunities. To evaluate whether the insight is economically significant, one has to depart from the standard paradigm of testing CIP violations using observable rates when discounting all contractual cash flows associated with over-the-counter (OTC) transactions.

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 Sundareshan (2012) Coffey, Hrungr, and Sarkar (2009), Griffolli and Rinaldo (2011), and McGuire and von Peter (2012). Ivashina, Scharfstein, 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, Hrungr, 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, Hrungr, 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 to LTA empirically. Fang and Liu (2021) develop a theory of CIP deviations due 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. We further differ from these papers by explicitly connecting new no-arbitrage benchmarks across the term structure of CIP conditions.

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 constraints 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 asset pricing anomalies.

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

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

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

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

2.2 The role of intermediary constraints

The NA perspective is helpful as it codifies how asset prices should relate to each other when agents are unconstrained both in terms of access to information and ability to trade on that information. In this context, the documented departures of cross-currency bases or xccy rates from zero are striking because of their persistence and apparent information transparency in the currency markets. Thus, the leading interpretation attributes the evidence to intermediary constraints.

The intermediary constraints considered in the literature take a variety of forms. A unifying theme connecting that 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 (e.g., [Brunnermeier and Sannikov, 2014](#), [He and Krishnamurthy, 2013](#)). Yet, there are 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, leading to limits to arbitrage (LTA) (e.g., [Dow and Gorton, 1994](#), [Garleanu and Pedersen, 2011](#), [Shleifer and Vishny, 1997](#)). Thus, while documented NA violations suggest the effects of intermediary constraints, the implicit constraints on intermediaries must be of a certain kind.

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, $\text{€}1/S_0$, at the annualized foreign T -period “risk-free rate” $\hat{r}_{0,T}$, 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:

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

implying that

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

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 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 Syrdad, 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 Sundareshan \(2007\)](#) demonstrate that these cash flows represent the opportunity cost of collateral, which can be represented as a dividend yield on an asset. Thus,

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

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

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

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

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

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

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

where i and \hat{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 . That is unlikely, however, because collateralized borrowing costs are lower than uncollateralized borrowing costs, unless counterparty risk is perceived to be trivial. This was the case for interbank borrowing before the GFC.

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,

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

where y and \hat{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 \hat{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 match the reported forward prices. Thus, we treat r' and \hat{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 help identify the funding costs of major investors. 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 1A 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 $\hat{i}_t + X$ on the EUR leg at each date $t + 1$.³ 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

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

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

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

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., $\phi_{0,T} + S_0 \hat{\phi}_{0,T} = 0$. The condition yields the formula for the constant maturity xccy swap rate $X_{0,T}$:

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

³To save on notation, we drop the second subscript for interest rates applicable to subsequent time periods such that $i_{t-1,t} = i_{t-1}$.

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 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.

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 us 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 $b_{0,T}^r = 0$ at the 3-month horizon to identify foreign rates \hat{r} via the observed forward premiums. In other words, we ask whether there exists a set of interest rates that is consistent with a zero basis while jointly explaining the term structure of IRS across countries.

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 1B.

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

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

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

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

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

$$\begin{aligned} X_{0,T} &= \left(\sum_{t=1}^T E_0 [\widehat{M}'_{0,t}] \right)^{-1} \\ &\times \left(CMS_{0,T} \sum_{t=1}^T E_0 [M'_{0,t}] - \widehat{CMS}_{0,T} \sum_{t=1}^T E_0 [\widehat{M}'_{0,t}] + E_0 [M'_{0,T}] - E_0 [\widehat{M}'_{0,T}] \right). \end{aligned} \quad (5)$$

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, $\hat{\cdot}$. 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 $\eta_t = \delta_{\eta,0} + \delta_\eta^\top z_t$.

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 $i_{0,T} = T^{-1} \log(1 + i_{0,T}^q \cdot T \cdot 30/360)$ where $i_{0,T}^q$ denotes a quoted LIBOR rate and $T \leq 12$ corresponds to maturities of up to 12 months.⁴ As a result,

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

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

⁴The 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.

then implies:

$$CMS_{0,T} = \frac{\sum_{t=1}^T E_0 [M'_{0,t} i_{t-1}]}{\sum_{t=1}^T E_0 [M'_{0,t}]} \quad (6)$$

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.

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 $\hat{r}'_{0,T} = r'_{0,T} - \rho_{0,T}$. 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.⁵ 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. 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, $b_{0,T}^i$, 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 2A 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 2B shows that the LIBOR-IRS curves exhibit

⁵DKK 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.

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 $b_{0,0.25}^r$ (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 $b_{0,0.5}^r$. 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 (5). 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 pricing errors $xccy^e$ are defined as the difference between the observed, $xccy^d$, and model-implied, $xccy^m$, xccy rates. The average $xccy^m$ are consistent with the evidence. For instance, the average 5-year $xccy^e$ is 3 bps while the average $xccy^d$ is 21 bps.

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.

Next, we resort to a simple variance decomposition to quantify how much variation in xccy rates is explained by our NA model. We exploit the fact that pricing errors are orthogonal to the model-based swap rates, by the model estimation design, in population. We report in the first row of Table 3 the fraction of the variance of the observed xccy rate levels explained by the model. On average, it is 68.33%. The remaining 31.67% is, therefore, attributed to the pricing errors.

Thus, although the NA framework plays a major role in understanding the behaviour of CIP, there is room left to be explained. As discussed in section 2.2, these deviations from the NA model 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 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 characterize the role of xccy pricing errors 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:

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

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](#)).⁶ Having obtained the U.S. convenience premium λ , we obtain the foreign $\hat{\lambda}$ from the Treasury basis via

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

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 its proxy $\tilde{r}_{0,T}$. 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 \tilde{r} and the model-implied interest rate, differences between them are not surprising. Even if there is no noise associated with the ingredients of \tilde{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 \tilde{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 \tilde{r} itself. Furthermore, we

⁶The 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.

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, Hrungr, 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 \tilde{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 \tilde{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 \tilde{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 \tilde{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, up to a noise term). The EFR is lower than LIBOR, consistent with the economic difference between collateralized and uncollateralized funding rates, respectively.⁷

⁷The EFR is lower than LIBOR and negative during the post-GFC period for some countries, in particular for the Euro area (and Denmark), Switzerland and Japan. This evidence is consistent

As a further characterization of the difference between the EFR and LIBOR, we consider the theoretical connection between this difference and xccy 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 Xccy pricing errors

While our fit of xccy rates is reasonable, there are pricing errors. Our descriptive model accounts for contractual cash flows of the swap contracts. That includes interest earned on posted collateral and the counterparty-risk free discount rate associated with collateralization. However, our framework does not account for any frictions associated with these transactions. For instance, we do not account for limited collateral that dealers might possess or other constraints highlighted in Section 2.2.

with the corresponding Central Banks adopting a negative interest rate policy (NIRP) in 2012, which is manifested via negative LIBOR rates in Figure 5. We provide in the online appendix confidence bands to the estimated EFRs. Accounting for statistical uncertainty implies less dramatic departures from zero.

Thus, the pricing errors may reflect a NA component omitted by our simple model, or a LTA manifestation. Disentangling these two components is beyond the scope of our paper. In this section we investigate a more narrow question of whether one can interpret the xccy pricing errors as potential components of risk premiums associated with intermediary constraints. 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 group uses the intermediary factors advocated in the literature. 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.

Second, we consider various measures of uncertainty: 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.

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 6 reports the results. 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.

[Haddad and Muir \(2021\)](#) 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. Their empirical proxies for intermediary risk aversion can be more broadly interpreted as intermediary health that may limit intermediaries’ ability to take advantage of arbitrage opportunities. 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 [Haddad and Muir \(2021\)](#) 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.

Overall, this evidence further supports the view that NA violations in the form of xccy pricing errors are reflective of priced intermediary risk, consistent with [Du, Hebert, and Huber \(2019\)](#). Future literature could explore incorporating an explicit and empirically implementable link between intermediary constraints and derivative contracts.

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 LIBOR 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 relate the cross-section of these pricing errors to measures of intermediaries' leverage and their willingness to provide USD funding. The evidence is consistent with existing 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, $f - s - (i^{\$} - i^{\text{€}})$, 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.

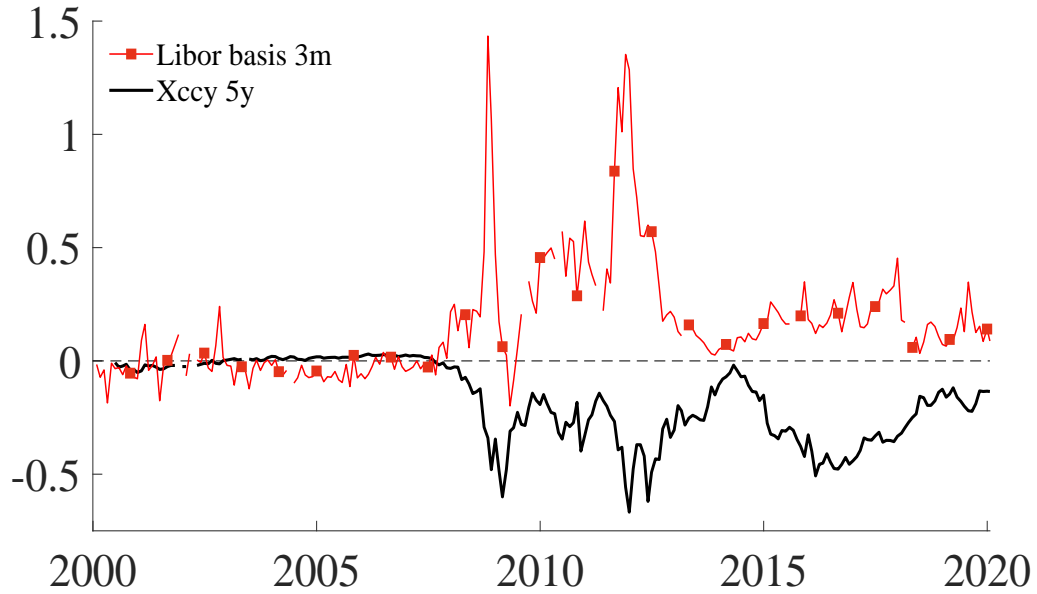
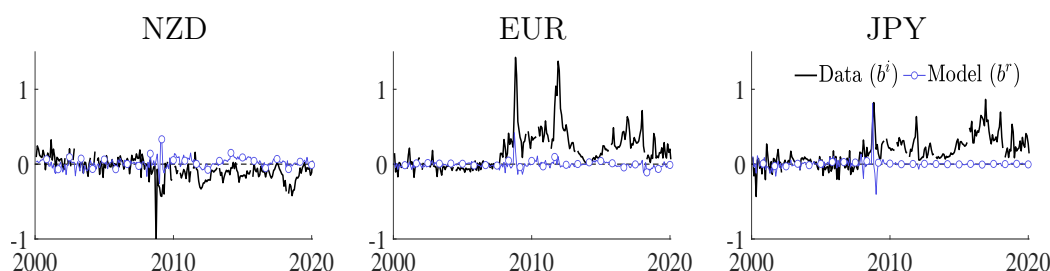
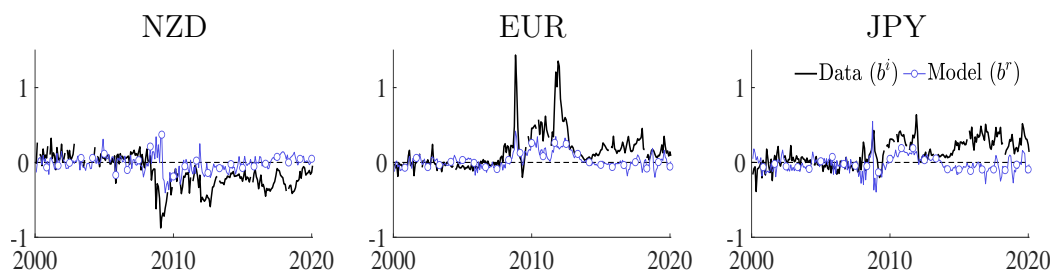


Figure 2: **Time-series of forward basis and xccy for NZD, EUR, and JPY.** In these figures, we report the time series of the forward basis (3 and 6 months, based on LIBOR in the data and on EFR in the model) or xccy rates (5 and 20 years) implied from the model and compare it with the data. The sample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

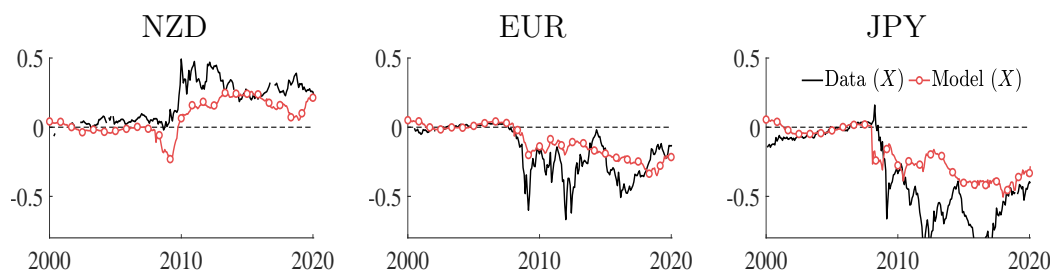
(A) Forward basis 3-month maturity



(B) Forward basis 6-month maturity



(C) xccy 5-year maturity



(D) xccy 20-year maturity

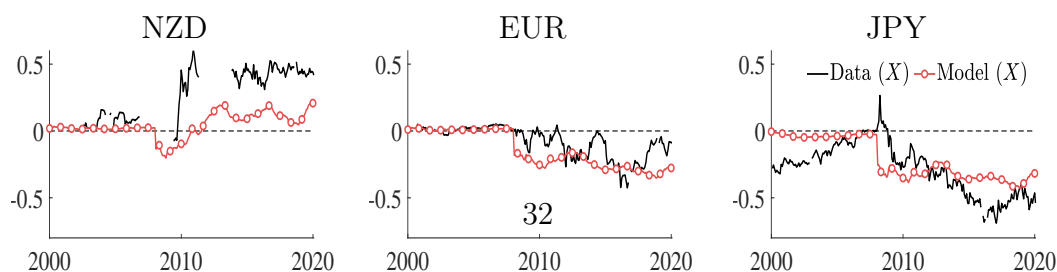


Figure 3: **Forward basis and xccy rate.** We report the mean of the forward basis (based on LIBOR in the data and on EFR in the model) and the cross-currency basis swap rate (in bps). We also report the cross-sectional average of absolute rates, AVG. All exchange rates are expressed as the USD price 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.

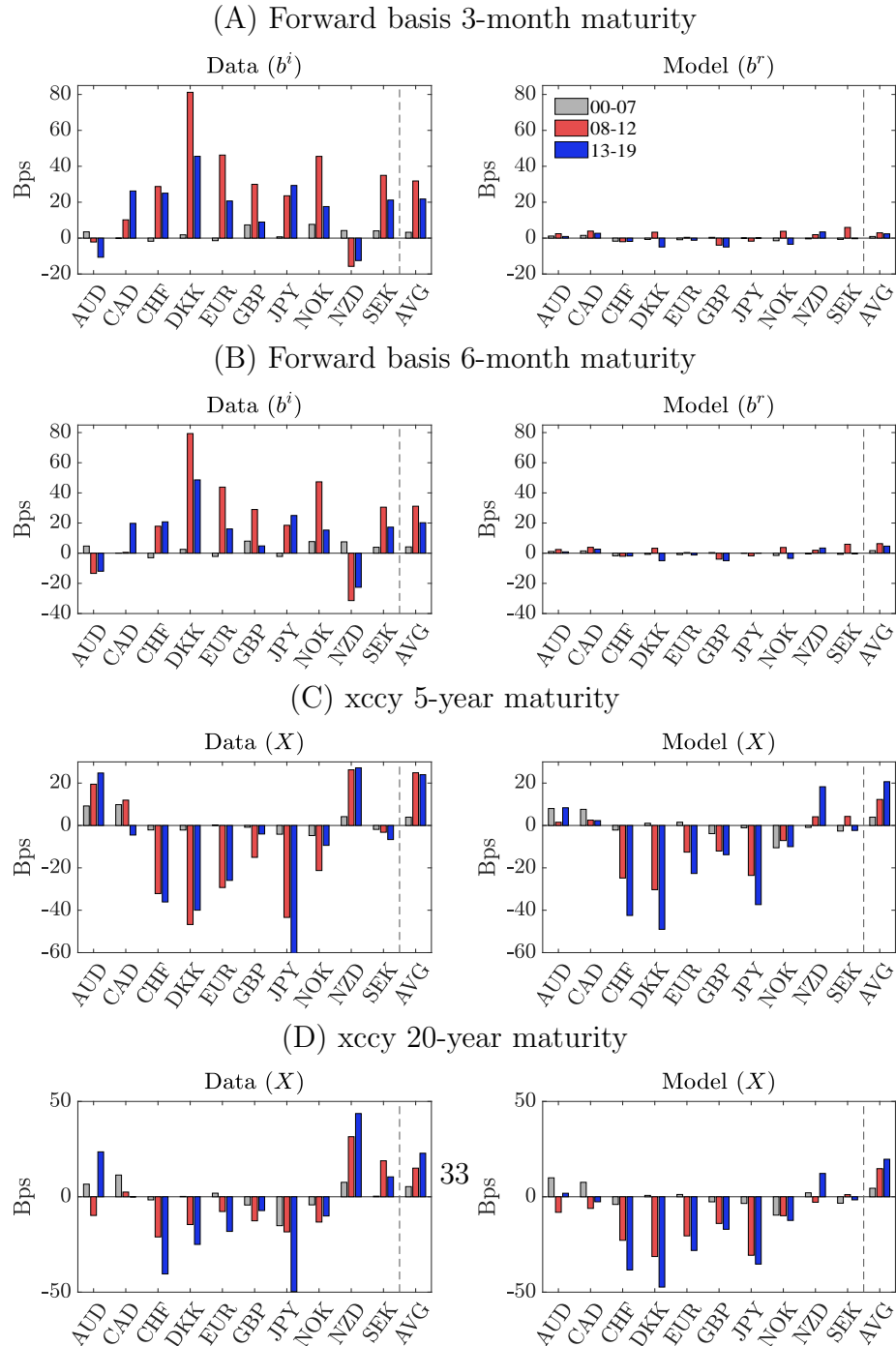


Figure 4: **Comparison of 1Y Interest Rate Proxies.** Each figure compares the model-implied 1-year interest rate to the predicted one and given by $\Delta r = -0.01 + 0.17 \cdot \Delta \text{LIBOR} + 0.40 \cdot \Delta(\text{Treasury} + \lambda) - 0.24 \cdot \Delta \text{CDS}$, where LIBOR corresponds 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.

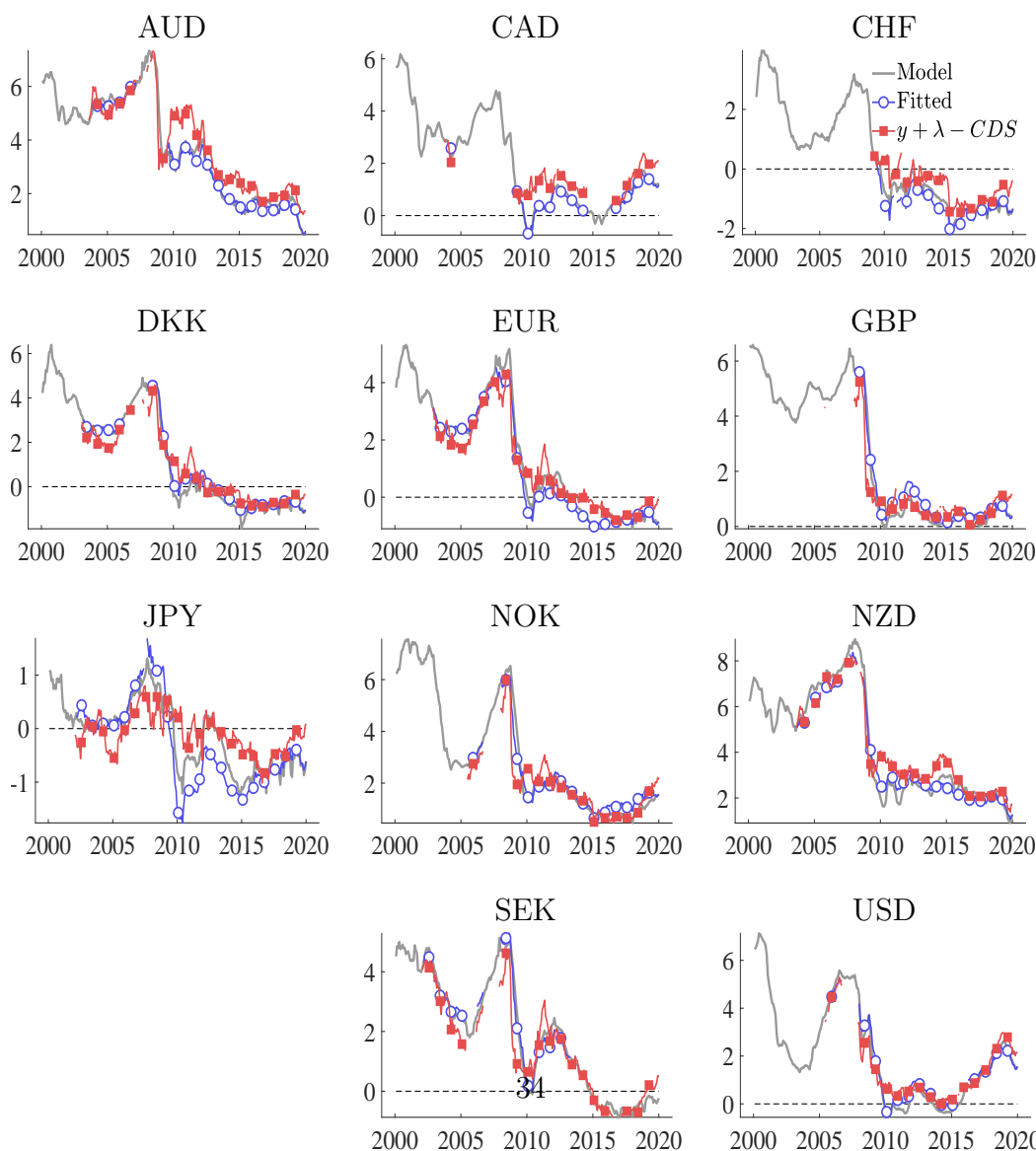


Figure 5: **Comparison of 3M Interest Rate Proxies.** Each figure compares the model-implied 3-month interest rate to the predicted one and given by $\Delta r = -0.01 + 0.31 \cdot \Delta \text{LIBOR} + 0.38 \cdot \Delta (\text{Treasury} + \lambda)$, where LIBOR corresponds to the country-specific Libor/interbank 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.

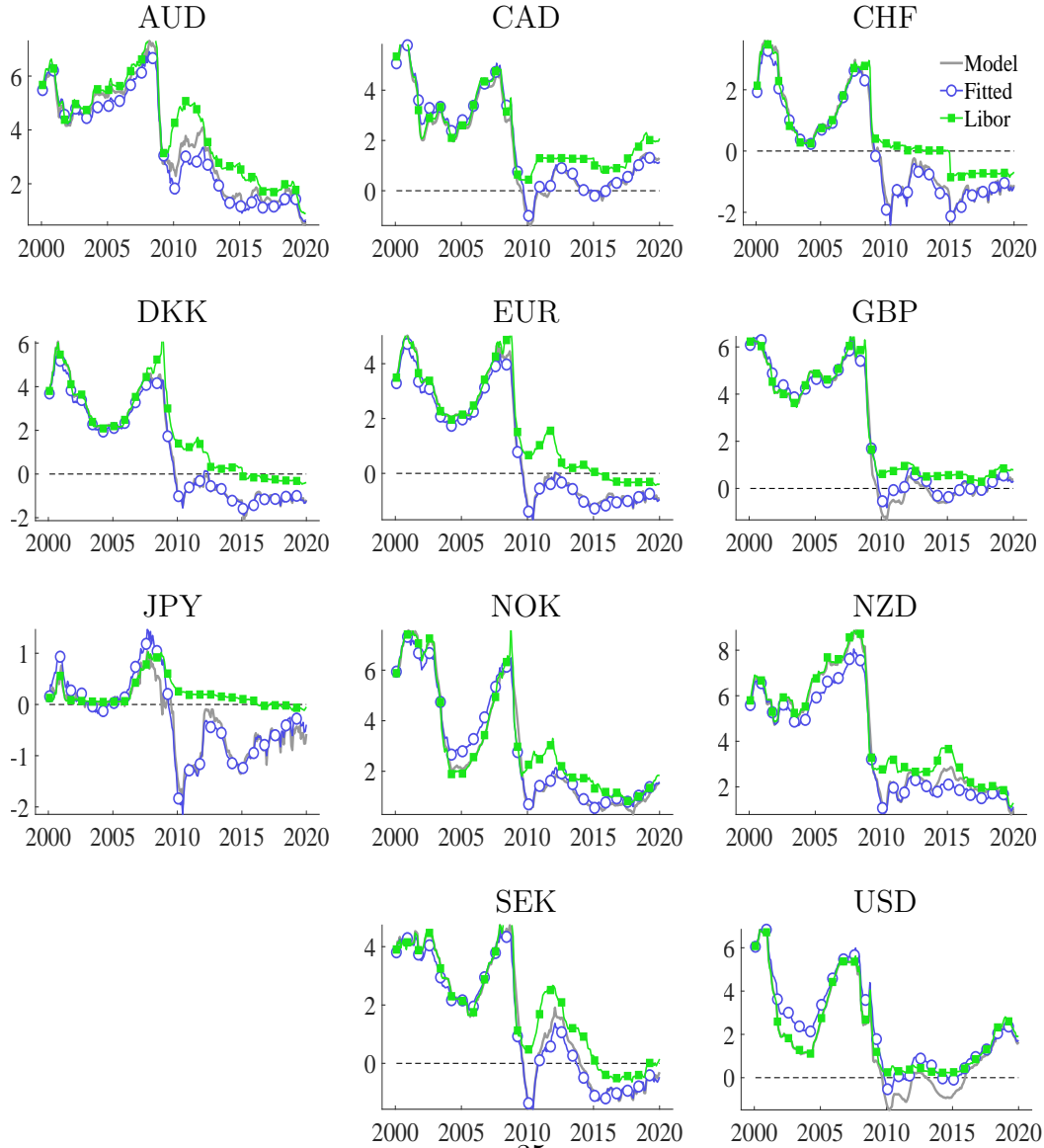


Figure 6: **Factor Exposure of Xccy Basis Swap Spread Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied 5-year xccy basis swap rate on a risk factor, i.e., $\Delta xccy_t^e = \alpha + \beta \cdot RF_t + \varepsilon_t$. We then project the average level of the xccy spread on the estimated betas $\hat{\beta}$. 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.

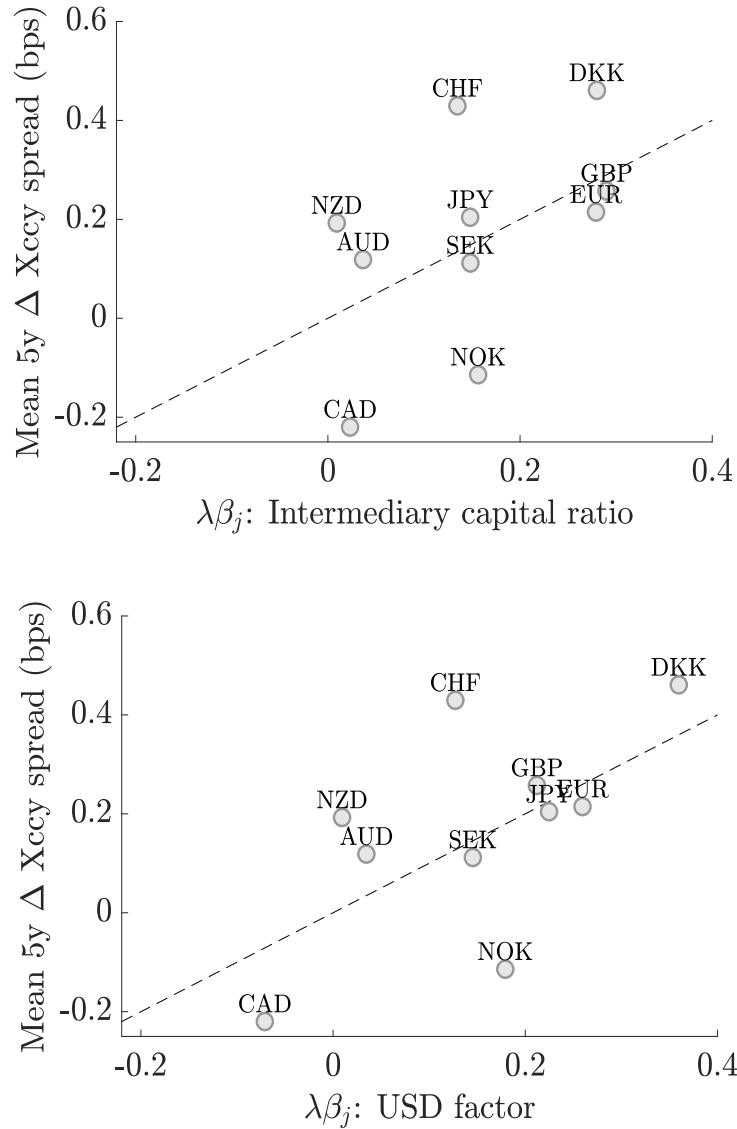


Figure 7: **Factor Exposure of Xccy Basis Swap Spread and Forward Premium Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied (i) 5-year xccy basis swap rate and (ii) the 6-month forward premium on the [Haddad and Muir \(2021\)](#) intermediary risk aversion factor, i.e., $\Delta xccy_{t+1}^e = \alpha + \beta \cdot RF_t + \varepsilon_t$. 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 accounted for by intermediaries for FX forwards and FX swaps, respectively. The sample period is 2000Q1 to 2017Q3.

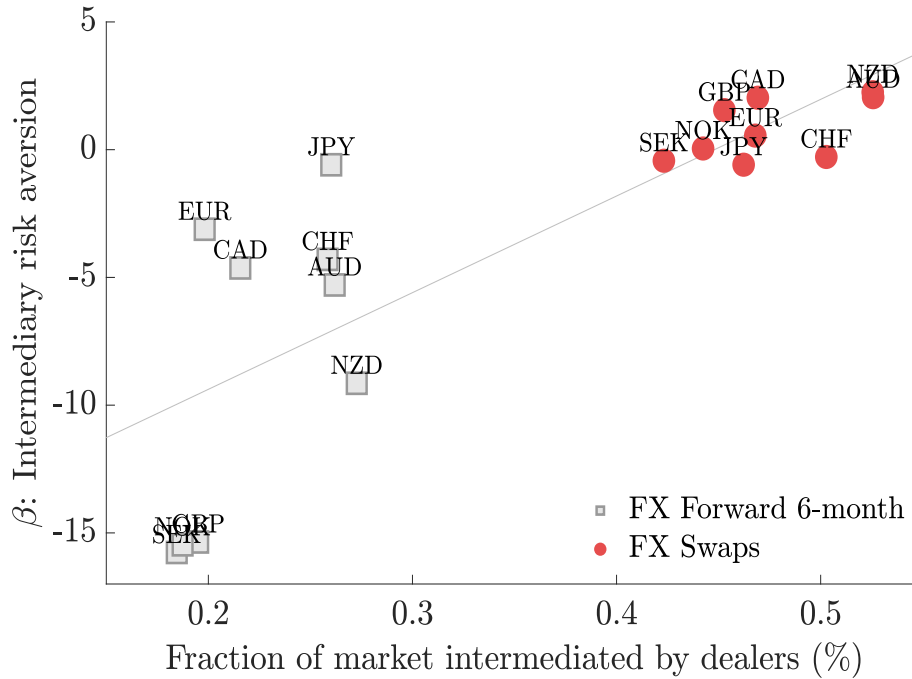


Table 1: Cash flows from a plain vanilla and fixed-for-fixed cross-currency basis swap

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 $\hat{i}_t + X$ 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 $\hat{i}_t + X$ 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.

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

Table 2: 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 3: 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. The residual $xccy^e = xccy^d - xccy^m$ is orthogonal to $xccy^m$ by construction (in population). That property lends natural variance decomposition of $xccy^d$. 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
$var(xccy^m)/var(xccy^d)$	68.33	79.78	57.34	55.62	71.28	48.48	79.91	100.50	53.89	83.10	53.36
$var(xccy^e)/var(xccy^d)$	31.68	20.22	42.65	44.37	28.72	51.52	20.08	-0.49	46.11	16.89	46.64

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*** (0.05)									0.31*** (0.06)	0.17** (0.07)
λ				0.25*** (0.03)									
Treasury+ λ					0.57*** (0.03)							0.38*** (0.04)	0.40*** (0.05)
EFFR						0.14*** (0.03)							
CD-Treasury							0.13*** (0.04)						
CDS								0.03 (0.09)					-0.24** (0.10)
Treasury-CDS									0.31*** (0.04)				
Treasury+ λ -CDS										0.47*** (0.05)			
BOX											0.19*** (0.06)		
Constant	-0.01*** (0.00)	-0.01*** (0.00)	-0.01*** (0.00)	-0.02*** (0.00)	-0.01*** (0.00)	-0.02*** (0.00)	-0.02*** (0.00)	-0.02*** (0.00)	-0.02*** (0.00)	-0.02*** (0.00)	-0.02*** (0.00)	-0.01*** (0.00)	-0.01*** (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 ($|x_{ccy5y}^D|$) on the spread between LIBOR and model-implied interest rates at the 3-month maturity ($3M(i - r)$). At the country level, we control for the Treasury yield adjusted for the convenience premium (Treasury + λ), 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) $ x_{ccy5y}^D $	(2) $ x_{ccy5y}^D $	(3) $ x_{ccy5y}^D $	(4) $ x_{ccy5y}^D $	(5) $ x_{ccy5y}^D $
$3M(i - r)$	4.75*** (0.85)	21.10*** (1.70)	17.89*** (1.54)	18.66*** (1.70)	10.08*** (2.65)
$\Delta\text{Treasury} + \lambda$			-1.65*** (0.31)	-2.04*** (0.33)	2.16*** (0.80)
CDS				10.70 (6.51)	7.90 (4.98)
Constant	20.81*** (0.81)	7.84*** (1.35)	12.33*** (1.38)	12.66*** (1.66)	14.82*** (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: 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 xcy 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)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$	$\Delta xccy^e$
AEM-LV2	-0.01 (0.08)											
HKM-ICR		137.98*** (24.92)										180.49*** (54.24)
USD FACTOR			-0.64*** (0.10)									-0.44*** (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* (2.61)	-7.02*** (2.65)
Constant	-0.01 (0.26)	0.12 (0.09)	0.11 (0.12)	0.06 (0.09)	0.05 (0.09)	0.04 (0.09)	0.05 (0.09)	0.06 (0.09)	0.05 (0.09)	0.06 (0.09)	0.04 (0.10)	0.16 (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