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THE TERM STRUCTURE OF COVERED INTEREST RATE PARITY VIOLATIONS

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ABSTRACT

We quantify the impact of risk-based and non-risk-based intermediary constraints (IC) on the term structure of CIP violations. Using a stochastic discount factor (SDF) inferred from interest rate swaps, we value currency derivatives. The wedge between model-implied and observed derivative prices reflects the impact of non-risk-based IC because our SDF incorporates risk-based IC. There is no wedge at short horizons, while the wedge accounts for 40% of long-term CIP violations. Consistent with IC theory, the wedge correlates with the shadow cost of intermediary capital, and the SDF-implied interest rate is a weighted average of collateralized and uncollateralized interest rates.

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1 Introduction

CIP violations are viewed as prima facie evidence that intermediary constraints (IC) matter for asset valuation (e.g., Ivashina, Scharfstein, and Stein, 2015; Du, Tepper, and Verdelhan, 2018). These violations are measured using currency forwards at short, and using cross-currency basis swaps (XCCY) at long horizons. In this paper we quantify the impact of IC on CIP violations across its term structure.

IC come in a variety of forms, most notably in form of risk based constraints, such as those implied by an asset's volatility or Value-at-Risk, and non-risk based constraints, such as leverage caps or margin requirements. Measuring the impact of IC constraints on CIP violations involves delineating the impact of these two types of constraints as they operate via different channels. The former operate similarly to and jointly with conventional risk factors via covariance with a stochastic discount factor (SDF). The latter constraints result in departures from the standard covariance-based framework and manifest themselves as anomalies.

Separating the two constraints is impossible to accomplish by observing CIP violations alone. We make progress by introducing a model of risk. By taking a stand on such a model, we are able to identify the non-risk-based IC component of these violations. Quantifying this component is our primary focus because it is difficult to disentangle the effects of risk factors and risk-based IC.

Specifically, we extract an SDF by estimating an affine term structure model that accurately prices interest rate swaps (IRS) across G10 countries. This model-implied SDF is particularly relevant because IRS are closely related to XCCY contracts. Institutional arrangements are similar since both contracts trade over-the-counter among a similar set of market participants, they have similar collateralization requirements and cash flows that are contractually linked to LIBOR. Moreover, the IRS market is highly liquid with a weekly traded notional of one trillion USD and subject to central clearing, suggesting a relatively small role for non-risk IC.

We show more formally below that if non-risk-based IC do not matter or if they have a similar impact across the two markets, the IRS-implied SDF would correctly price

XCCY as well. In contrast, if non-risk-based IC do matter, they should manifest themselves as departures between observed XCCY premiums and the ones implied by our model. We refer to such departures as 'IC wedge'.

For example, the traditional measure of the IC wedge used in the literature is the cross-currency basis (measured as the spread between the forward currency premium and the difference between LIBOR rates in two countries) and the XCCY premium at short and long horizons, respectively. Figure 1 offers an example, using the 3-month basis and the 5-year XCCY premium for the Euro to U.S. Dollar (USD) exchange rate. Departures from zero for both time series are attributed by the literature to IC.

We use the estimated SDF model to gauge the magnitude and structure of the IC wedge. That is, we value FX forwards and XCCY. What we discover is that the traditional benchmark used in the literature to measure CIP violations is an inaccurate measure of the IC wedge. In fact, the IC wedge is much smaller than the traditional measures of CIP violations would suggest. For instance, the average 5-year XCCY rate across countries during the post-crisis period is 22 basis points (bps) in the risk-based model vs. 25 bps in the data. Our variance decomposition of the 5-year XCCY rates shows that the risk-based model explains, on average, 60% of the variation in their levels. Thus, one conclusion is that XCCY and IRS markets are, to a large degree, priced consistently.

Two elements of our approach drive the risk-based valuation of XCCY close to that in in the data. 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).[1] We associate the mean of the IRS-based SDF with the EFR. Second, XCCY contracts feature recurring cash flows that are 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 or in the absence of IC.

[1] OIS 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)

We next assess whether the differences between risk-based and observed prices are associated with non-risk-based IC. Figure 1 displays the timing of several regulatory capital and macroprudential regulations that were enacted during our sample period. These include the Basel III Capital Requirements Regulation as a framework to tighten banks' capital and liquidity requirements (12/2010), the start by major dealer banks to formally show Fair Value Adjustments on their balance sheets (01/2011), and the Money Market Fund Reform (10/2016). The various constraints were introduced at different times and, often, long after the emergence of CIP deviations. Therefore, any given regulation is unlikely to be the sole IC driver of the evidence. Similarly, Fleckenstein and Longstaff (2020) emphasize the importance of ICs long before the GFC even in the absence of apparent CIP violations. Thus, we are looking for a more generic mechanism for interpreting our evidence.

We rely on the margin-based pricing framework of Gârleanu and Pedersen (2011) as margin constraints existed both before and after the GFC. One prediction of this framework is that the EFR is a weighted average of the collateralized and uncollateralized interest rates, where the weight is equal to the margin in the fixed-income derivatives market. To be precise, we interpret margins of the Gârleanu and Pedersen (2011) model more broadly as all non-risk-based IC and do not only refer to the literal definition of asset-specific margins. If one approximates the collateralized rate by the Treasury yield adjusted for convenience and credit risk, and the uncollaterlized rate by LIBOR, we find indeed that the estimated EFR is closely approximated by the weighted average of the two.

The model further predicts that the difference between uncollateralized and collateralized interest rates is equal to the shadow cost of capital, which should be proportional to departures of asset risk premiums from the risk-based framework. We find that our measures of IC wedges align with the model-implied measure of the shadow cost of capital. Thus, we attribute the 40% of variation in XCCY not captured by the risk-based model to the quantitative effects of non-risk-based IC constraints. Furthermore, the shadow cost of capital has only a weak empirical connection to other proxies of IC used in the literature, such as the leverage of bank holding companies (He, Kelly, and Manela, 2017) or the trade-weighted U.S. dollar index (Avdjiev, Du, Koch, and Shin, 2019; Jiang, Krishnamurthy, and Lustig, 2018).

Haddad and Muir (2021) establish the importance of intermediaries in pricing by showing that returns of assets that are more intermediated are more sensitive to intermediary risk aversion. Similarly, we relate IC wedges to intermediary risk aversion and examine how their sensitivity to risk aversion relates to the degree of cross-asset intermediation. 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 non-trivial role (40%) in the determination of variation in XCCY premiums.

In summary, our paper offers a novel quantitative framework that allows gauging the impact of IC on asset prices. We apply a risk-based framework that prices a set of relatively less intermediated assets (IRS) to a set of relatively more intermediated assets (XCCY). The gap between risk-based and observed XCCY valuations is labeled as IC wedges. Guided by the margin-based asset pricing framework of Gârleanu and Pedersen (2011), we confirm that IC wedges are IC-related. As a result, we attribute 60% of variation in XCCY to risk-based valuation and 40% to non-risk-based IC.

Related literature

An important literature ascribes post-GFC CIP deviation to IC.[2] The existing literature disagrees, however, about the type of frictions that matter for explaining CIP deviations. See Du and Schreger (2021) for a review.

Initial explanations point towards frictions in global intermediation of USD funding (Baba, Packer, and Nagano, 2008; Bottazzi, Luque, Pascoa, and Sundaresan, 2012; Coffey, Hrung, and Sarkar, 2009; Griffolli and Ranaldo, 2011; McGuire and von Peter, 2012; Ivashina, Scharfstein, and Stein, 2015; Bahaj and Reis, 2018) and risk-based IC (Baba and Packer, 2009; Coffey, Hrung, and Sarkar, 2009; Csavas, 2016; Levich, 2012;

[2]Discussions and test of CIP deviations go back to Keynes (1923). For pre-GFC analysis, when CIP was not violated, see Dooley and Isard (1980); Frenkel and Levich (1975); Fletcher and Taylor (1994, 1996); Pasquariello (2014); Popper (1993); Taylor (1987, 1989).

Fong, Valente, and Fung, 2010; Skinner and Mason, 2011; Tuckman and Porfirio, 2003; Wong, Leung, and Ng, 2016).

Another prominent explanation for large and persistent CIP deviations is linked to non-risk-based IC (Du, Tepper, and Verdelhan, 2018; Avdjiev, Du, Koch, and Shin, 2019; Cenedese, Della Corte, and Wang, 2021; Borio, McCauley, McGuire, and Sushko, 2016; Boyarchenko, Eisenbach, Gupta, Shachar, and Tassel, 2020), which may be amplified in the presence of hedging demand (Borio, McCauley, McGuire, and Sushko, 2016; Liao, 2020) or funding shocks (Anderson, Du, and Schlusche, 2019; Liu, 2019). Fang and Liu (2021) highlight interactions between risk-based and non-risk based constraints.

As mentioned earlier, the introduction of various regulations over time depicted in Figure 1 suggests that these mechanisms are unlikely to rationalize both the pre- and post-GFC evidence. Moreover, the evidence does not convey the quantitative impact of IC on asset valuations. We offer a unified framework that provides a quantitative assessment of how much non-risk-based IC matter. That is a departure from the literature since we theoretically show that non-zero XCCY rates may naturally arise even in the absence of IC if LIBOR is an imperfect benchmark for the funding rate. This perspective differs from the literature that unanimously considers non-zero XCCY rates as a manifestation of IC.

Like us, Du, Hebert, and Huber (2022) target quantitative implications using a different empirical strategy. They estimate the shadow cost of constraints using returns of forward CIP trading strategies. Thus, they agree with the rest of the literature in treating all departures of the cross-currency basis from zero as a manifestation of IC. While they show that their IC measure is priced in the cross-section of assets, we identify the pricing of IC as a residual from risk-based CIP valuation that we link to non-risk-based IC.

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 approach the problem from a fundamentally different perspective by inferring EFR from related markets with identical cash flows, identical market participants, and similar institutional arrangements.

Our approach has broader implications for how we think about and interpret other asset pricing puzzles. Du, Hebert, and Li (2022) adopt a similar approach for explaining negative swap spreads and CIP violations, Binsbergen, Diamond, and Grotteria (2022) for extracting the EFRs of option market makers, and Fleckenstein and Longstaff (2020) for explaining the Treasury cash-futures basis.

2 Conceptual framework

In this section we outline two key ideas that we use in this paper to develop our empirical approach and to interpret our findings. First, we review the SDF-based valuation approach that we use to develop our evidence and quantify the impact of IC on CIP violations. We term the difference between the SDF-based and traded CIP valuations the IC wedge. Second, we introduce a variant of the Gârleanu and Pedersen (2011) model as an equilibrium underpinning of IC wedges in order to interpret the evidence. In particular, we use that framework to measure the intermediary cost of capital and to test whether the IC wedges are consistent with the advocated economic mechanism.

2.1 SDF-based valuation

We follow Andersen, Duffie, and Song (2019) and distinguish between traded prices of currency-linked derivatives and their fair market valuation. The latter is represented by a function of payoffs that satisfies two coherency assumptions: linearity and increase in payoffs. These two assumptions imply the existence, but not necessarily the uniqueness, of an SDF, $M > 0$, with the property that the value of any payoff Y is $E(MY)$ (we are omitting conditioning and timing of payoffs for brevity).

As Andersen, Duffie, and Song (2019) emphasize, although this description appears to be similar to that of arbitrage-free valuation, the absence of arbitrage is not assumed because the SDF is not unique. It is just a representation of how market valuations are assigned via the SDF. We interpret the difference between market valuations and traded prices as the quantitative effect of non-risk IC and refer to it as 'IC wedge'.

In particular, we associate a rate $r = -\log E(M)$ with an asset whose payoff is $Y = 1$. This rate does not have to coincide with the risk-free rate because of the potential multiplicity of M. Thus, we interpret r as the EFR of a bank that engages in collateralized and uncollateralized borrowing and lending.

2.2 The role of intermediary constraints

We present a simple equilibrium framework with constrained financial institutions in the spirit of Gârleanu and Pedersen (2011). In our setup, these institutions face both risk and non-risk-based constraints. Time is discrete and there are two dates, $t = 0, 1$.

We consider two types of financial institutions. Institution A represents constrained buy side investors, such as pension, mutual, or endowment funds. Institution B represents sell side investors acting as dealers, e.g., a major investment bank.

Each institution j is endowed with initial wealth $W_{0,j}$ and invests into two types of risky assets (IRS and XCCY) with returns r_k, $k = 1, \ldots, K$. Each type of swap has multiple maturities, thus $K \geq 2$.

Both institutions fund their investments in money markets using either collateralized loans at the rate r^c or uncollateralized loans at the rate r^u. While r^c may intuitively be thought of as repo rate, we argue in the Online Appendix that repo rates are imperfect proxies for r^c in practice. Hence we treat r^c as unobservable. The rate r^u may be represented by LIBOR.[3]

We denote by θ_j^u and θ_j^c institution j's (dollar) position in uncollateralized and collateralized loans, respectively, and by θ_j^k the positions in the two risky assets. The positions satisfy the requirement:

$$\theta_j^c + \theta_j^u + \sum_{k=1}^{K} \theta_j^k = W_{0,j}. \tag{1}$$

Each institution's budget constraint becomes:

$$W_{1,j} = \theta_j^c(1 + r^c) + \theta_j^u(1 + r^u) + \sum_{k=1}^{K} \theta_j^k(1 + r^k).$$

We assume that institutions face several constraints. On one hand, institutions manage their overall risk exposure using Value-at-Risk or similar risk management techniques. We capture such constraints by modeling institutions as risk-averse with absolute risk aversion parameters α_j. Thus, they exhibit mean-variance preferences, as in, e.g., Acharya, Lochstoer, and Ramadorai (2013) or Greenwood, Hanson, Stein, and Sundaram (2020). As in Gârleanu and Pedersen (2011), institution A is more risk-averse than B, does not participate in the market for uncollateralized loans, and may face limits for derivatives positions.

On the other hand, institutions face non-risk-based constraints manifested by margins m_i which specify a fraction of the investment that must be financed by an agent's own capital. We interpret m as capturing all non-risk-based IC (e.g., leverage constraints, funding value adjustments, liquidity coverage ratio) and not just the asset-specific margin requirements.

We thus require, as in Gârleanu and Pedersen (2011), that the total margin capital

[3] OIS may serve as a more modern version of r^u. We focus on LIBOR because all derivative contracts we consider are explicitly linked to LIBOR. See Augustin, Chernov, Schmid, and Song (2021) for a comparison of the two rates.

cannot exceed the initial endowment, so that

$$\theta_j^u + \sum_{k=1}^{K} m_k |\theta_j^k| \leq W_{0,j}.$$

Denoting the shadow cost of institution B's balance sheet constraint by ψ, we can write the institution's Lagrangian as:

$$\begin{aligned} \mathcal{L}_B &= \theta_B^c(1+r^c) + \theta_B^u(1+r^u) + \sum_{k=1}^{K} \theta_B^k E(1+r^k) \\ &- \frac{\alpha_B}{2} Var\left(\sum_{k=1}^{K} \theta_B^k (1+r^k)\right) - \psi\left(\theta_B^u + \sum_{k=1}^{K} m_k |\theta_B^k| - W_{0,B}\right), \end{aligned}$$

where the first and second terms in the second row represent risk- and non-risk-based constraints, respectively.

From the first order conditions for uncollateralized and collateralized loans and accounting for (1), we obtain:

$$r^u - r^c = \psi, \tag{2}$$

allowing us to identify the shadow cost of capital from proxies for r^u and r^c. From the first order condition with respect to the risky assets, we get, for asset i:

$$E(r^i - r^c) = \alpha_B \cdot Cov\left(r^i - r^c, \sum_{k=1}^{K} \theta_B^k (r^k - r^c)\right) + \psi \cdot m_i \cdot sgn(\theta_B^i). \tag{3}$$

Thus, risk-based constraints impact the first traditional covariance term. Non-risk-based constraints appear as departures from the risk premium in the second term.

As a consequence, all else equal, if two assets have the same payoffs, the asset with the higher margin requirement commands a higher expected return. This is intuitive, as such an asset ties up more capital. For concreteness, for two risky assets, suppose $m_2 > m_1$, then we have for long positions that $E(r^2 - r^1) = \psi(m_2 - m_1)$.

Now, suppose that m_i's for IRS of all maturities are the same and equal to m. Also, assume that B holds net positive positions in an IRS of each maturity, i.e., $\theta_B^i \geq 0$ for all i corresponding to IRS. Imagine constructing an SDF that values IRS contracts as described in Section 2.1. Then the rate associated with the mean of such an SDF is equal to:

$$r = r^c + \psi \cdot m. \tag{4}$$

We refer to this rate as effective funding rate (EFR). Combining this equation with the shadow cost of capital in Equation (2), we obtain that the EFR is a weighted average of the collateralized and uncollateralized rates:

$$r = m \cdot r^u + (1 - m) \cdot r^c. \tag{5}$$

Lastly, applying Equation (3) to value XCCY with r as reference rate, one obtains:

$$E(r^{i'} - r) = \alpha_B \cdot Cov\left(r^{i'} - r^c, \sum_{k=1}^{K} \theta_B^k (r^k - r^c)\right) + \psi \cdot \left(m' \cdot sgn(\theta_B^{i'}) - m\right), \tag{6}$$

where the second term reflects the IC wedge. The IC wedge can be zero under two scenarios. First, non-risk-based constraints are zero in all markets that we consider, in which case $\psi = 0$. Second, non-risk-based constraints have the same impact across all markets and, therefore, cancel out, because $m' = m$. These observations underscore that we consider the relative rather than the absolute impact of non-risk IC. For instance, if the XCCY market is not affected by non-risk IC, we could still obtain a non-zero IC wedge as long as the IRS market is subject to non-risk IC. Thus, to have a sensible measure of the IC impact in our empirical work, we use the arguably less constrained market (IRS) as a reference and the more constrained market as a target.

We use predictions of this stylized model to interpret the evidence developed in the rest of the paper. Subsequently, we use continuously compounded rates, which are approximately equal to simple compounded rates characterized in this section.

3 CIP in the short and the long run

To separate the effect of risk-based and non-risk-based constraints on asset prices, one must take a stand on the model of risk. First, we review risk-based valuation for currency forwards and XCCY. Second, we develop such a model to extract EFRs on the basis of IRS. This is conceptually aligned with Binsbergen, Diamond, and Grotteria (2022) and Fleckenstein and Longstaff (2020), who argue that prices of risky financial assets help identify the funding costs of major investors. Third, we apply the model to value currency forwards and XCCY.

3.1 Forward rates and short-term cross-currency 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 $F_{0,T} - S_T$. These contracts, being overcollateralized derivatives, generate additional cash flows associated with daily marking to market and posting of collateral. 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. We rely on the SDF-based pricing relation to value a forward contract and account for these additional cash flows. Thus,

$$E_0(M_{0,T} e^{\eta_{0,T}} F_{0,T}) = S_0 \cdot E_0(M^*_{0,T} e^{\eta^*_{0,T}}), \tag{7}$$

where η and η^* represent the domestic and the foreign cost of collateral, respectively, and M^* is the foreign SDF. Rearranging terms in Equation (7), we thus have that:

$$F_{0,T}/S_0 = E_0(M^*_{0,T} e^{\eta^*_{0,T}})/E_0(M_{0,T} e^{\eta_{0,T}}). \tag{8}$$

The forward premium in logs is given by:

$$f_{0,T} - s_0 = \log E_0(M^*_{0,T} e^{\eta^*_{0,T}}) - \log E_0(M_{0,T} e^{\eta_{0,T}}) = T(r'_{0,T} - r^{*\prime}_{0,T}), \tag{9}$$

where $r'_{0,T} \equiv -T^{-1} \log E_0(M_{0,T} e^{\eta_{0,T}})$ and $r^{*\prime}_{0,T} \equiv -T^{-1} \log E_0(M^*_{0,T} e^{\eta^*_{0,T}})$ are the corresponding domestic and foreign rates (combined with the extra cash flows for brevity). The domestic and foreign EFRs, r and r^*, are obtained by setting ηs to zero.

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

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

which 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:

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

where i and i^* represent LIBOR or OIS and their foreign counterparts. The 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:

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

where y and 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, which is consistent with our perspective.

As highlighted in Equation (5), the EFRs, r and r^* are equal to a weighted average of secured and unsecured rates, where the former is not easily approximated with an observable interest rate. The corresponding weight is equal to the "representative" margin and is not observable either. These observations compel us to treat EFRs as unobservable.[4] Likewise, the "representative" extra contractual cashflows η are not observable either.

3.2 Long-term 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. 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 contract initiation, pays S_0 dollars and receives 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 $i_t^* + X$ on the EUR leg at each date $t+1$.[5] 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 XCCY is:

$$\phi_{0,T} = S_0 \left(-1 + \sum_{t=1}^{T} E_0 \left[M_{0,t} e^{\eta_{0,t}} i_{t-1} \right] + E_0 \left[M_{0,T} e^{\eta_{0,T}} \right] \right),$$

and the present value of all expected future cash flows on the EUR leg is:

$$\phi_{0,T}^* = 1 - \sum_{t=1}^{T} E_0 \left[M_{0,t}^* e^{\eta_{0,t}^*} \left(i_{t-1}^* + X_{0,T} \right) \right] - E_0 \left[M_{0,T}^* e^{\eta_{0,T}^*} \right].$$

[4] Rime, Schrimpf, and Syrstad (2019) take the complementary route of estimating marginal funding rates using information about wholesale money market funding from non-bank investors.

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

Here we assume that the opportunity cost of collateral, expressed in the form of a dividend yield, is the same as that for forward contracts.

The XCCY is fairly priced if both the USD and the EUR legs have the same value in USD, i.e., $\phi_{0,T} + S_0\phi^*_{0,T} = 0$. The condition yields the formula for the constant maturity XCCY swap rate $X_{0,T}$:

$$X_{0,T} = \left(\sum_{t=1}^{T} E_0 \left[M^*_{0,t} e^{\eta^*_{0,t}} \right] \right)^{-1} \times \tag{10}$$
$$\left[\left(\sum_{t=1}^{T} E_0 \left[M_{0,t} e^{\eta_{0,t}} i_{t-1} \right] + E_0 \left[M_{0,T} e^{\eta_{0,T}} \right] \right) - \left(\sum_{t=1}^{T} E_0 \left[M^*_{0,t} e^{\eta^*_{0,t}} i^*_{t-1} \right] + E_0 \left[M^*_{0,T} e^{\eta^*_{0,T}} \right] \right) \right].$$

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.

3.3 A risk-based valuation framework

If the collection of M and M^* were observable, we could evaluate expressions in Equations (9) and (10) and compare them to the observed forward premiums and

XCCY rates, respectively. The difference would tell us about the quantitative effects of non-risk-based constraints. In practice, we have to estimate M and M^*.

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 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:

$$\pi_{0,T} = S_0 \left(-1 + \sum_{t=1}^{T} CMS_{0,T} E_0 \left[M_{0,t} e^{\eta_{0,t}} \right] + E_0 \left[M_{0,T} e^{\eta_{0,T}} \right] \right)$$

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

$$\pi^*_{0,T} = \left(+1 - \sum_{t=1}^{T} \left(CMS^*_{0,T} + X_{0,T} \right) E_0 \left[M^*_{0,t} e^{\eta^*_{0,t}} \right] - E_0 \left[M^*_{0,T} e^{\eta^*_{0,T}} \right] \right).$$

Since XCCY is priced fairly if $\pi_{0,T} + S_0 \pi^*_{0,T} = 0$, we get another expression for $X_{0,T}$:

$$X_{0,T} = \left(\sum_{t=1}^{T} E_0 \left[M^*_{0,t} e^{\eta^*_{0,t}} \right] \right)^{-1} \times \tag{11}$$
$$\left(CMS_{0,T} \sum_{t=1}^{T} E_0 \left[M_{0,t} e^{\eta_{0,t}} \right] - CMS^*_{0,T} \sum_{t=1}^{T} E_0 \left[M^*_{0,t} e^{\eta^*_{0,t}} \right] + E_0 \left[M_{0,T} e^{\eta_{0,T}} \right] - E_0 \left[M^*_{0,T} e^{\eta^*_{0,T}} \right] \right).$$

Thus, we express the XCCY rate in terms of (observable) interest swap rates and (unobserved) discount factors, $M_{0,t}$ and $M^*_{0,t}$.

Now we can develop a model of the SDFs. We describe our model for the U.S. only. All other countries have the same notation augmented with asterisks. We assume that the unobservable state is captured by a vector z_t that follows a VAR(1):

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

and the "dividend yield" associated with costly collateral is $\eta_t = \delta_{\eta,0} + \delta_\eta^\top z_t$.

Next we have to specify how the contract reference rate LIBOR, i_t, and the EFR, r_t depend on the state. In doing so we have to accommodate essentially a structural break in the data. Before the crisis, LIBOR was treated as de-facto funding rate, i.e., $b^i \approx 0$. This perspective agrees with the widespread pre-crisis view 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. Clearly, this is no longer the case in the post-crisis environment.

Thus, we posit 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. After the crisis, there is a credit spread between LIBOR and the EFR. Specifically, the one-month LIBOR rate is:

$$i_t = \delta_{i,0} + \delta_i^\top z_t.$$

The EFR is:

$$r_t = i_t - \delta_{r,0} - \delta_r^\top z_t.$$

This assumption is consistent with the intensity-based approach to modeling credit risk (e.g., Duffie and Singleton, 1999).

This specification raises a concern that we potentially mix the effect of the onset of the zero-lower bound (ZLB) environment with the increase in the credit risk of the banking sector. We use the "reverse" formulation of riskless and credit-risky rates, where the credit spread is subtracted from the credit-risky rate, to highlight that the model of the observed LIBOR/swap curve is the same throughout our sample. Thus, the structural break assumption does not affect the model performance due to the ZLB with respect to the observables. We offer additional analysis of the interaction with the ZLB issue in the Online Appendix.

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. The conditional mean of the log SDF is related to the EFR r_t. This is consistent with the frameworks laid out in Section 2: the constructed SDF is set up to value IRS only; the mean of that SDF reflects the shadow cost of capital and margin in this market as per Equation (4).[6]

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.[7] As a result,

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

modeled without dividend yield η because LIBOR represents uncollateralized lending.

Now we can use the 3-month LIBOR rates for computing the IRS. Here we discount all cash flows accounting for the cost of collateral η_t. Standard arguments then imply:

$$CMS_{0,T} = \frac{\sum_{t=1}^{T} E_0\left[M_{0,t} e^{\eta_{0,t}} i_{t-1}\right]}{\sum_{t=1}^{T} E_0\left[M_{0,t} e^{\eta_{0,T}}\right]}. \tag{12}$$

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.

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

[6]That we do not estimate the 'true' risk-free rate is not a limitation of our framework, but a manifestation of the lack of uncontroversial risk-free assets. If such asset existed, we could have included its price into our estimation and the estimated r would be interpreted as the risk-free rate.

[7]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.

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 can be captured using a set of baseline assumptions.

3.4 Empirical strategy

We estimate the model specified in the previous section using the standard state-space framework. We treat the state variables as latent. IRS prices and 3-month cross-currency bases are used in observation equations. Because IRS rates depend non-linearly on the state, we use Bayesian MCMC methods to estimate the model parameters and the state realizations. Details are provided in the Online Appendix.

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_i and covariance matrix Σ control the scale of i_t, we set the former to unity. All other parameters are free.

We need extra consideration in addition to the standard identification restrictions. That is because 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 treat the U.S. different from all other countries in that we first estimate the model using the U.S. data, and subsequently estimate the remaining countries one by one conditional on the output from the U.S. estimation. In the U.S. data estimation, we rely on the non-linearity of the IRS rates to identify the variation in r controlled by δ_r. That is because r contributes differently to the numerator and denominator of CMS in Equation (12) (the constant $\delta_{r,0}$ cancels out). To identify $\delta_{r,0}$, we bound the post-GFC sample average of the wedge to be $E\left[\delta_{r,0} + \delta_r^\top z_t\right] \leq 100$ bps (annualized). This interval is fairly wide, incorporating a wide range of different views about likely deviation of EFR with respect to LIBOR. Lastly, the cost of collateral appears in the valuation of IRS, but not LIBOR. That helps with identifying η separately from r. In practice, because identification comes from only few observations at the short end of the curve, it is difficult to pin down the time-varying component of η, so we treat it as a constant.

Moving on to other countries, the identification strategy for r that we use for the U.S. has practical problems. On one hand, the short-maturity IRS rates are insensitive to variations of δ_r, and so the identification mostly comes from the very long end of the IRS curve (e.g., 20 or 30 years). On the other hand, some countries have missing data precisely on that very long end of the swap curve. Lastly, while set identification, which we use for $\delta_{r,0}$ in the U.S., can be appealing as it does not require any strong parameter restrictions, we cannot continue using this strategy for other countries as it defeats our purpose of gauging the magnitude and structure of the IC wedge. To tackle this issue, we assume that at the 3-month horizon b^r is close to zero throughout the whole sample (up to the usual measurement error used in state-space models). This assumption is similar to that of Binsbergen, Diamond, and Grotteria (2022) and Fleckenstein and Longstaff (2020) who identify "risk-free rates" from derivative prices. Similar to the U.S., η is identified off from the difference between IRS and LIBOR.

4 Evidence

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

4.1 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.[8] 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 XCCY 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 XCCY 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. 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.

4.2 Results

Fitting an affine term structure model to a swap curve is a standard exercise that is not expected to yield many surprises. Table 2B shows that the LIBOR-IRS curves exhibit a two- to three-factor structure. Pricing errors are rather large when one uses three factors (e.g., Collin-Dufresne, Goldstein, and Jones, 2008, Dai and Singleton, 2000). The primary reason is that it is hard to capture jointly the short end of the swap curve that is driven by LIBOR rates and the long end that is driven by actual swaps. Thus, we ultimately choose the dimension of z_t to be 4 in our model.

[8] 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.

4.2.1 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 3-month basis computed using the EFR, $b^r_{0,0.25}$, (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.

Interpreting this evidence through the lens of Equation (6), we conclude that the impact of non-risk-based IC is similar across the IRS and currency forward markets. As a consequence, returns in the excess of the EFR reflect only risk-based compensation. Since we used the 3-month forward premium to identify the EFR parameters, we also report the model-implied results at the 6-month horizon, which was not used for estimation at all.

4.2.2 XCCY rates

We use the estimated SDFs M and M^* to construct XCCY rates using Equation (11). 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.

We observe that the average market valuations implied by our model of the SDFs are quite close to the traded prices. For instance, the average difference in the case of 5-year instruments is 3 bps with the average traded price at 25 bps.

The model mechanically generates the change around the crisis because our modeling assumptions allow for departures between r and i. The economic interpretation of the specific quantitative effect is straightforward: the spread reflects the riskiness of the banking sector implicit in LIBOR. After the crisis, the relation between the traded and the theoretical X is weaker, reflecting the fact that the SDF model can match the general trend in XCCY rates, but not the local deviations.

As a result, there is meaningful time-series variation in the differences between the traded and theoretical valuations. Before the crisis, the two are visually similar. During the crisis, we see a broad switch in the level of X. For some currencies, like CHF, EUR, or GBP the switch in market valuation is broadly consistent with the observed one. In some cases, like CAD or SEK, the change is less dramatic. Still, it is consistent with the evidence.

Next, we resort to a simple variance decomposition to quantify how much variation in observed XCCY rates is explained by our SDF model. We exploit the fact that IC wedges 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 60%. The remaining 40% is, therefore, attributed to the IC wedges.

Thus, although the SDF-based valuation framework plays a major role in understanding the behaviour of CIP, the contribution of IC is important as well. As discussed in Section 2.2 in the context of Equation (6), the IC wedges reflect the relative effect of non-risk-based constraints that are not accounted for by the SDF-based framework. An alternative interpretation is that the difference between the theoretical and observed values of X reflect a misspecification of the SDF model. In that case the wedges would be unrelated to various measures of IC. We disentangle these two possibilities in the subsequent analysis.

5 Interpretation of the evidence

We first relate the estimated EFRs to observable variables. Second, we assess whether variation in the IC wedges correlates with non-risk-based constraints.

The first exercise has a dual purpose: investigate the source of the empirical success of our model and develop economic intuition for the estimated EFRs. One might worry that our results are driven by the dividend yield 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. Also, η is small ranging between 10 and 19 bps (annualized) across countries. More broadly, it is useful to understand the drivers of EFRs in fixed-income swap markets. In particular, it would be helpful to know if it can be interpreted as the risk-free rate rather than the representative lending rate of dealers in this market. Thus, one would want to understand the relation of r to other variables.

The second exercise allows us to to verify the IC origins of the gaps between traded and theoretical prices of derivatives and to quantify the impact of IC on valuation.

5.1 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 that could yield an approximation of the risk-free rate. The margin-based asset pricing theory tells us, however, that the EFR should not be equal to risk-free rate unless margins are zero. Therefore, as our second approach, 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 when thinking about risk-free rates. 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 risk-free rate:

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

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; Joslin, Li, and Song, 2019).[9] Having obtained the U.S. convenience premium λ, we obtain the foreign λ^* from the Treasury basis via:

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

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 $\widetilde{r}_{0,T}$. We see that the proxy is tracking the EFR quite well. The connection between the EFR and the risk-free rate is not perfect and departures between the two are evident. Japan has the largest discrepancies. The observed differences between r and $\widetilde{r}$ 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 described in Equation (5).

As the relation between the conjectured and the estimated EFRs is not perfect, we investigate other variables. Our candidates are the ingredients of $\widetilde{r}$ 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

[9] 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.

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, 2022). 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.[10] 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 unobserved 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), capital shocks to common arbitrageurs, coordinated monetary policy, or regulatory reforms. 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 of 0.94 and 0.80, respectively, and R^2 around 0.8. The weakest variables are the U.S.-only ones: EFFR, CD-Treasury spread with coefficients of about 0.2 and R^2 below 0.07, and CDS, which is insignificant on its own. Our initial proxy for the EFR $\widetilde{r}$ is close to the most important variables with a coefficient of 0.7 and R^2 of 0.7.

[10] We implement regressions in changes due to the strong persistence of both left- and right-hand side variables, which increases the risk of spurious results.

Moving to multivariate regressions, we find that, at the 3-month horizon, convenience-adjusted Treasury rates and LIBOR are the two variables that remain significant. 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. Significance of LIBOR is consistent with the representation of EFR as a weighted average of the collateralized and uncollaterlized rates.

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. Thus our proxy for the risk-free rate $\widetilde{r}$ is selected as a significant variable.

One may try mapping the 1-year results to Equation (5) and infer the implicit costs of non-risk-based constraints. Since the regression coefficients on LIBOR and $\widetilde{r}$ are 0.38 and 0.50, Equation (5) would imply a cost $m = 0.38$.[11] The two coefficients do not add up to 1, which could be due to month and currency fixed effects, potentially omitted variables, and because $\widetilde{r}$ is a noisy proxy for r^c. Thus, the link to Equation (5), although attractive, should be considered with caution.

It is interesting that OIS is not significant in multivariate regressions. Some might view this as surprising in the context of the 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

[11]The implicit value of m matches the baseline calibration of Gârleanu and Pedersen (2011), who use $m = 0.4$. Margins vary quite a bit between low values for derivatives and high values for funded assets. As Gârleanu and Pedersen (2011) point out, margins are opaque and their estimates also depend on the specific combination of brokers and clients and the time period. Given that we interpret m as a cost associated with a host of non-risk-based constraints, not just margins only, the value 0.38 may be considered plausible.

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.

5.2 Is the EFR different from LIBOR?

As mentioned earlier, one concern could be that LIBOR is a good proxy for the EFR and, thus, all the explanatory power in the model is driven by the extra cash flows in the form of the dividend yield η. First, Figure 5 explicitly compares our EFR with LIBOR. It is evident that i is substantively different from r during the post-crisis period (they are similar before the crisis as part of our identification strategy, up to a noise term). The EFR is lower than LIBOR, consistent with the theory.[12]

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 SDF-based 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 22 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 for a fairer comparison across periods. That specification suggests a 28 bps XCCY rate in absolute value for a 100 bps

[12]The 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 with the corresponding Central Banks adopting negative interest rate policies (NIRP) in 2012, 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.

spread between LIBOR and r. This is economically significant, as 28 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.

5.3 IC wedges

While our SDF-implied XCCY rates match the evidence reasonably well, there are differences from observed rates. These differences could reflect model misspecification or IC wedges due to differences in margins and positioning across the IRS and XCCY markets, as highlighted in Section 2.2.

We use the Gârleanu and Pedersen (2011) framework outlined in Section 2.2 to distinguish between the two interpretations. Equation (3) suggests that departures from the risk-based valuation should be related to the shadow cost of capital, which equals the difference between collateralized and uncollateralized interest rates. Since we do not observe the former, we exploit its relation to the EFR. Equations (2) and (5) imply that the difference between LIBOR and EFR, $r^u - r$, is also proportional to the shadow cost of capital ψ.

Armed with this observation we evaluate whether the differences between SDF-implied and observed XCCY rates, $XCCY^e$, are related to $r^u - r$, which we label ψ for simplicity. We regress changes in $XCCY^e$ on changes in ψ and control variables. We measure the shadow cost of capital in two different ways, using U.S. interest rates or using the first principal component of $r^u - r$ across countries in our sample. We focus on the three-month maturity for better comparability to other variables, but we find qualitatively similar results using the one-month maturity.

We consider three broad groups of control variables, starting with the intermediary factors advocated in the literature. Specifically, we measure IC using 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). We also considered the leverage of security broker-dealers (Adrian, Etula, and Muir, 2014, AEM), but do not report it because it is measured quarterly, and, therefore, not comparable to other variables in multivariate regressions.

Second, we consider measures of uncertainty: the Jurado, Ludvigson, and Ng (2015) real, macroeconomic, and financial uncertainty measures; the Bekaert and Hoerova (2014) uncertainty and risk aversion measures; and the CBOE VIX index. We only report results for variables that remain significant in multivariate regressions.

In the third group we use indicators of the opportunity cost of money like the certificate of deposit rate over Treasury yield spread (Nagel, 2016), or of distress in the banking sector like the U.S. LIBOR-OIS spread. The latter is insignificant. See the Online Appendix for details.

The results in Table 6 indicate that ψ is significant regardless of the specific proxy. In columns (1) and (4), we control for currency fixed effects to absorb cross-currency differences in XCCY changes. The estimated coefficient also changes little in sign when we move from the univariate regressions in columns (1) and (4) to multivariate regressions with controls in columns (2) to (3) and (5) to (6). The negative sign of the coefficient suggests negative net positions in XCCY by the least risk averse dealer. Although this interpretation might be taking the margin-based framework too literally, it is not inconsistent with recent work by Du, Hebert, and Li (2022), who suggest that dealers reduce CIP arbitrage activity and increase Treasury swap spread positions in the post-GFC period.

Overall, the evidence suggests that the component of XCCY premiums that is not explained by the IRS-implied SDF can be interpreted as IC wedges, i.e., premiums reflecting IC. Revisiting Table 3, 40% of XCCY premium is thus attributed to IC.

The literature has explored many IC proxies. We check, in Table 7, whether our measure of shadow cost of capital brings novel information for research considerations. We, therefore, regress changes in ψ on changes in similar proxies for "intermediary variables" from the literature discussed in the context of Table 6. We report only those coefficients that remain statistically significant in univariate regressions.

While some of the variables in Table 7 are statistically significant, the amount of variation in ψ they explain remains low, according to the R^2 of the regressions. This suggests that our theoretically motivated shadow cost of capital could represent a new tool for the study of the impact of IC on asset valuations.

Last, but not least, we evaluate if intermediary risk aversion affects the cross-section of IC wedges in the spirit of Haddad and Muir (2021). In this context, we interpret IC wedges as excess returns after hedging out the risk-based pricing components.

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 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, we conduct similar cross-sectional tests for the IC wedges of both 5-year XCCY rates and 6-month forward premiums.

Specifically, we regress changes in IC wedges 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 6 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 of IC wedges associated with XCCY to intermediary risk aversion, while that relation is much noisier for forward premiums.

6 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 XCCY that have traded at non-zero prices. These violations have prominently been linked to various types of IC.

We examine the dynamics of the term structure of CIP violations across G11 currencies in a unifying framework. Specifically, we quantify the impact of non-risk-based IC on CIP violations and differentiate that from the quantitative impact of risk-based IC and other conventional sources of risk.

For our analysis, we back out a stochastic discount factor from plain vanilla interest rate swap contracts and use this discount factor to price forward exchange rates and XCCY across all eleven currencies. Using that discount factor, we find no evidence of short-term CIP violations and explain about 60% of long-term CIP violations.

We rely on IC theory to interpret the wedge between model-implied and observed long-term CIP violations. Consistent with the theory, we find that the wedge correlates with the shadow cost of intermediary capital, and that IC wedges line up with the degree of intermediation of the different forward and XCCY contracts. We also provide evidence that the effective funding rate is a weighted average of collateralized and uncollateralized interest rates.

References

Acharya, Viral V., Lars A. Lochstoer, and Tarun Ramadorai, 2013, Limits to arbitrage and hedging: Evidence from commodity markets, *Journal of Financial Economics* 109, 441–465.

Adrian, Tobias, Nina Boyarchenko, and Or Shachar, 2017, Dealer balance sheets and bond liquidity provision, *Journal of Monetary Economics* 89, 92–109.

Adrian, Tobias, Erkko Etula, and Tyler Muir, 2014, Financial Intermediaries and the Cross-Section of Asset Returns, *Journal of Finance* 69, 2557–2596.

Andersen, Leif, Darrell Duffie, and Yang Song, 2019, Funding Value Adjustments, *Journal of Finance* 74, 145–192.

Anderson, Alyssa, Wenxin Du, and Bernd Schlusche, 2019, Arbitrage Capital of Global Banks, *Working Paper University of Chicago*.

Augustin, Patrick, Mikhail Chernov, Lukas Schmid, and Dongho Song, 2021, Benchmark interest rates when the government is risky, *Journal of Financial Economics* pp. 74–100.

Avdjiev, Stefan, Wenxin Du, Cathérine Koch, and Hyun Song Shin, 2019, The Dollar, Bank Leverage, and Deviations from Covered Interest Parity, *American Economic Review: Insights* pp. 193–208.

Baba, Naohiko, and Frank Packer, 2009, Interpreting deviations from covered interest parity during the financial market turmoil of 2007-08, *Journal of Banking and Finance* pp. 1953–1962.

Baba, Naohiko, Frank Packer, and Teppei Nagano, 2008, The spillover of money market turbulence to FX swap and cross-currency swap markets, *BIS Quarterly Review*.

Bahaj, Saleem, and Ricardo Reis, 2018, Central Bank Swap Lines, *CEPR Discussion Paper No. DP13003*.

Bech, Morten L., and Elizabeth Klee, 2011, The mechanics of a graceful exit: Interest on reserves andsegmentation in the federal funds market, *Journal of Monetary Economics* 58, 415–431.

Bekaert, Geert, and Marie Hoerova, 2014, The VIX, the Variance Premium and Stock Market Volatility, *Journal of Econometrics* 183, 181–192.

Binsbergen, Jules H. van, William Diamond, and Marco Grotteria, 2022, Risk free interest rates, *Journal of Financial Economics* 143, 1–29.

Borio, Claudio, Robert McCauley, Patrick McGuire, and Vladyslav Sushko, 2016, Covered interest parity lost: understanding the cross-currency basis, *BIS Quarterly Review*.

Bottazzi, Jean-Marc, Jaime Luque, Mario Pascoa, and Suresh M. Sundaresan, 2012, Dollar Shortage, Central Bank Actions, and the Cross Currency Basis, *SSRN Working Paper 2167716*.

Boyarchenko, Nina, Thomas M. Eisenbach, Pooja Gupta, Or Shachar, and Peter Van Tassel, 2020, Bank-Intermediated Arbitrage, *Federal Reserve Bank of New York Staff Reports* 854.

Cameron, Matt, 2013, Goldman and the OIS Gold Rush, *Risk Magazine* June, 15–19.

Cenedese, Gino, Pasquale Della Corte, and Tianyu Wang, 2021, Currency Mispricing and Dealer Balance Sheets, *Journal of Finance* 76, 2763–2803.

Chernov, Mikhail, Lukas Schmid, and Andres Schneider, 2020, A macrofinance view of US Sovereign CDS premiums, *Journal of Finance* 75, 2809–2844.

Coffey, Niall, Warren Hrung, and Asani Sarkar, 2009, Capital Constraints, Counterparty Risk, and Deviations from Covered Interest Rate Parity, *Federal Reserve Bank of New York Staff Reports 393*.

Collin-Dufresne, Pierre, Robert S. Goldstein, and Christopher S. Jones, 2008, Identification of Maximal Affine Term Structure Models, *Journal of Finance* 63, 743–795.

Csavas, Csaba, 2016, Covered interest parity with default risk, *European Journal of Finance* pp. 1130–1144.

Dai, Qiang, and Kenneth J. Singleton, 2000, Specification Analysis of Affine Term Structure Models, *Journal of Finance* 55, 1943–1978.

Dooley, M. P., and P. Isard, 1980, Capital Controls, Political Risk, and Deviations from Interest-Rate Parity, *Journal of Political Economy* pp. 370–384.

Du, Wenxin, Benjamin Hebert, and Amy Huber, 2022, Are Intermediary Constraints Priced?, *Review of Financial Studies* Forthcoming.

Du, Wenxin, Benjamin Hebert, and Wenhao Li, 2022, Intermediary Balance Sheets and the Treasury Yield Curve, *Working Paper Stanford Graduate School of Business.*

Du, Wenxin, Joanne Im, and Jesse Schreger, 2018, The U.S. Treasury Premium, *Journal of International Economics* 112, 167–181.

Du, Wenxin, and Jesse Schreger, 2021, CIP Deviations, The Dollar, and Frictions in International Capital Markets, *NBER Working Paper Series 28777.*

Du, Wenxin, Alexander Tepper, and Adrien Verdelhan, 2018, Deviations from Covered Interest Rate Parity, *Journal of Finance* 73, 915–957.

Duffie, Darrell, and Kenneth J. Singleton, 1997, An Econometric Model of the Term Structure of Interest-Rate Swap Yields, *Journal of Finance* 52, 1287–1321.

Duffie, Darrel, and Kenneth J. Singleton, 1999, Modeling Term Structures of Defaultable Bonds, *Review of Financial Studies* 12, 687–720.

Fang, Xiang, and Yang Liu, 2021, Volatility, Intermediaries, and Exchange Rates, Journal of Financial Economics, forthcoming.

Fleckenstein, Matthias, and Francis Longstaff, 2020, Renting Balance Sheet Space: Intermediary Balance Sheet Rental Costs and the Valuation of Derivatives, *Review of Financial Studies* 33, 5051–5091.

Fletcher, Donna J., and Larry W. Taylor, 1994, A Nonparametric Analysis of Covered Interest Parity in Long-Date Capital Markets, *Journal of International Money and Finance* pp. 459–475.

Fletcher, Donna J., and Larry W. Taylor, 1996, "Swap" Covered Interest Parity in Long-Date Capital Markets, *Review of Economics and Statistics* pp. 530–538.

Fong, Wai-Ming, Giorgio Valente, and Joseph K.W. Fung, 2010, Covered interest arbitrage profits: The role of liquidity and credit risk, *Journal of Banking and Finance*.

Frenkel, Jacob A., and Richard M. Levich, 1975, Covered Interest Arbitrage: Unexploited Profits?, *Journal of Political Economy* pp. 325–338.

Gârleanu, Nicolae, and Lasse Heje Pedersen, 2011, Margin-based Asset Pricing and Deviations from the Law of One Price, *Review of Financial Studies* 24, 1980–2022.

Georgievska, Ljubica, 2020, The persistent element of the no-arbitrage framework: discount rates, bases, and opportunity cost of collateral, *Working Paper University of California Los Angeles*.

Greenwood, Robin, Samuel Hanson, Jeremy Stein, and Adi Sundaram, 2020, A Quantity-Driven Theory of Term Premiums and Exchange Rates, *Working Paper*.

Griffolli, Tommaso Mancini, and Angelo Ranaldo, 2011, Limits to Arbitrage during the Crisis: Funding Liquidity Constraints and Covered Interest Parity, *Swiss National Bank Working Paper*.

Haddad, Valentin, and Tyler Muir, 2021, Do Intermediaries Matter for Aggregate Asset Prices?, *Journal of Finance, forthcoming*.

He, Zhiguo, Bryan Kelly, and Asaf Manela, 2017, Intermediary asset pricing: New evidence from many asset classes, *Journal of Financial Economics* 126, 1–35.

Hull, John, and Alan White, 2013, LIBOR vs. OIS: The Derivatives Discounting Dilemma, *Journal of Investment Management* 11, 14–27.

Ivashina, Victoria, David S. Scharfstein, and Jeremy C. Stein, 2015, Dollar Funding and the Lending Behavior of Global Banks, *Quarterly Journal of Economics* 130, 1241–1281.

Jiang, Zhengyang, Arvind Krishnamurthy, and Hanno Lustig, 2018, Foreign Safe Asset Demand for US Treasurys and the Dollar, *American Economic Review P&P* 108, 537–541.

Jiang, Zhengyang, Arvind Krishnamurthy, and Hanno Lustig, 2019, Foreign Safe Asset Demand and the Dollar Exchange Rate, *Working Paper Stanford University*.

Johannes, Michael, and Suresh M. Sundaresan, 2007, The Impact of Collateralization on Swap Rates, *Journal of Finance* 62, 383–410.

Joslin, Scott, Anh Le, and Kenneth J. Singleton, 2013, Why Gaussian macro-finance term structure models are (nearly) unconstrained factor-VARs, *Journal of Financial Economics* 109, 604 – 622.

Joslin, Scott, Wenhao Li, and Yang Song, 2019, The Term Structure of Liquidity Premium, *working paper*.

Jurado, Kyle, Sydney C. Ludvigson, and Serena Ng, 2015, Measuring Uncertainty, *American Economic Review* 105, 1177–1216.

Keynes, John Maynard, 1923, *A tract on monetary reform.* (MacMillan and Co., Ltd., London).

Kohler, Daniel, and Benjamin Müller, 2018, Covered interest rate parity, relative funding liquidity and cross-currency repos, *Working Paper Swiss National Bank*.

Krishnamurthy, Arvind, and Annette Vissing-Jorgensen, 2012, The Aggregate Demand for Treasury Debt, *Journal of Political Economy* 120, 233–267.

Levich, Richard, 2012, FX Counterparty Risk and Trading Activity in Currency Forward and Future Markets, *Review of Financial Economics* 21, 102–110.

Liao, G., 2020, Credit migration and covered interest parity, *Journal of Financial Economics* 138, 504–525.

Litzenberger, Robert, 1992, Swaps: Plain and fanciful, *Journal of Finance* 47, 831–850.

Liu, Jianan, 2019, Comovement in Arbitrage Limits, *SSRN Working Paper 3242862*.

Longstaff, Francis A., 2004, The Flight-to-Liquidity Premium in U.S. Treasury Bond Prices, *Journal of Business* 77, 511–526.

McGuire, Patrick, and Goetz von Peter, 2012, The Dollar Shortage in Global Banking and the International Policy Response, *International Finance* 15, 155–178.

Nagel, Stefan, 2016, The Liquidity Premium of Near-Money Assets, *Quarterly Journal of Economics* 131, 1927–1971.

Pasquariello, Paolo, 2014, Financial Market Dislocations, *Review of Financial Studies* 27, 1868–1914.

Popper, Helen, 1993, Long-term covered interest parity: evidence from currency swaps, *Journal of International Money and Finance* pp. 439–448.

Ramaswamy, Krishna, and Suresh M. Sundaresan, 1986, The valuation of floating-rate instruments: Theory and evidence, *Journal of Financial Economics* 17, 251 – 272.

Rime, Dagfinn, Andreas Schrimpf, and Olav Syrstad, 2019, Covered Interest Parity Arbitrage, *Working Paper*.

Skinner, Frank S., and Andrew Mason, 2011, Covered interest rate parity in emerging markets, *International Review of Financial Analysis*.

Spears, Taylor, 2019, Discounting collateral: quants, derivatives and the reconstruction of the 'risk-free rate' after the financial crisis, *Economy and Society, forthcoming*.

Taylor, M., 1987, Covered Interest Rate Parity: A high-frequency, high-quality data study, *Econometrica* pp. 429–438.

Taylor, M. P., 1989, Covered Interest Arbitrage and Market Turbulence, *Economic Journal* pp. 376–391.

Tuckman, Bruce, and Pedro Porfirio, 2003, Interest Rate Parity, Money Market Basis Swaps, and Cross-Currency Basis Swaps, *Lehman Brothers: Fixed Income Liquid Markets Research*.

Wong, Alfred, David Leung, and Calvin Ng, 2016, Risk-Adjusted Covered Interest Parity: Theory and Evidence, *SSRN Working Paper 2834798*.

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. We highlight several regulatory capital and macroprudential regulations that were enacted during our sample period: SEC Net Capital Rule; Basel II; Basel II Trading Book; Dodd-Frank Act; Basel III; Fair Value Adjustments; Supplementary Leverage Ratio Regulation; Enhanced Supplementary Leverage Regulation; Overnight Reverse Repurchase Facility; Liquidity Coverage Ratio; Money Market Fund Reform. Source: Adrian, Boyarchenko, and Shachar (2017), Andersen, Duffie, and Song (2019), Fleckenstein and Longstaff (2020).

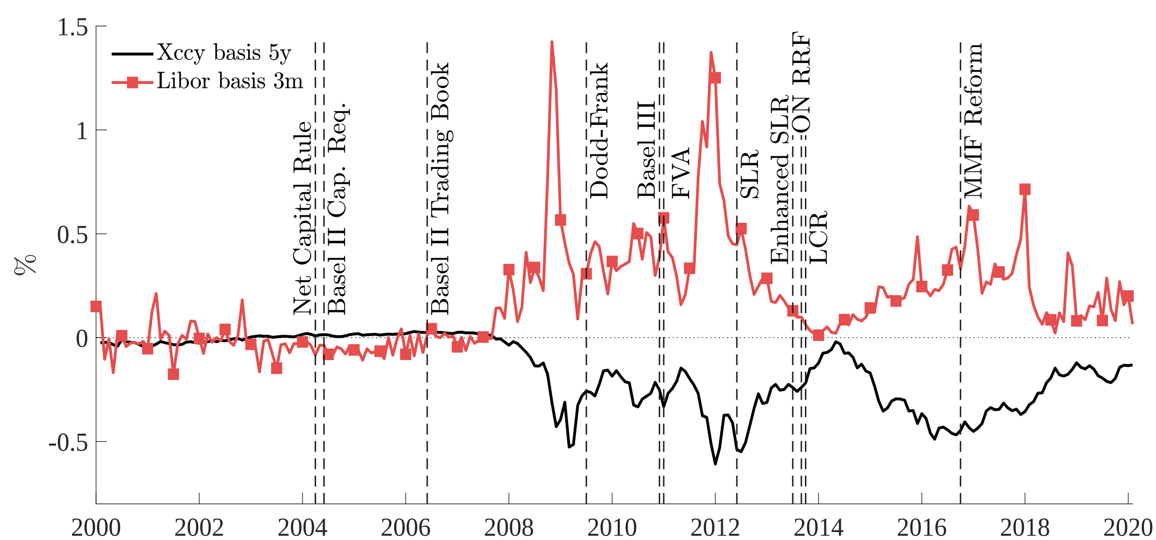

Figure 2: **Time-series of forward basis and XCCY for NZD, EUR, and JPY.** In these figures, we report the time series of the forward basis (3 and 6 months, based on LIBOR in the data and on EFR in the model) or XCCY rates (5 and 20 years) implied from the model and compare it with the data. The sample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

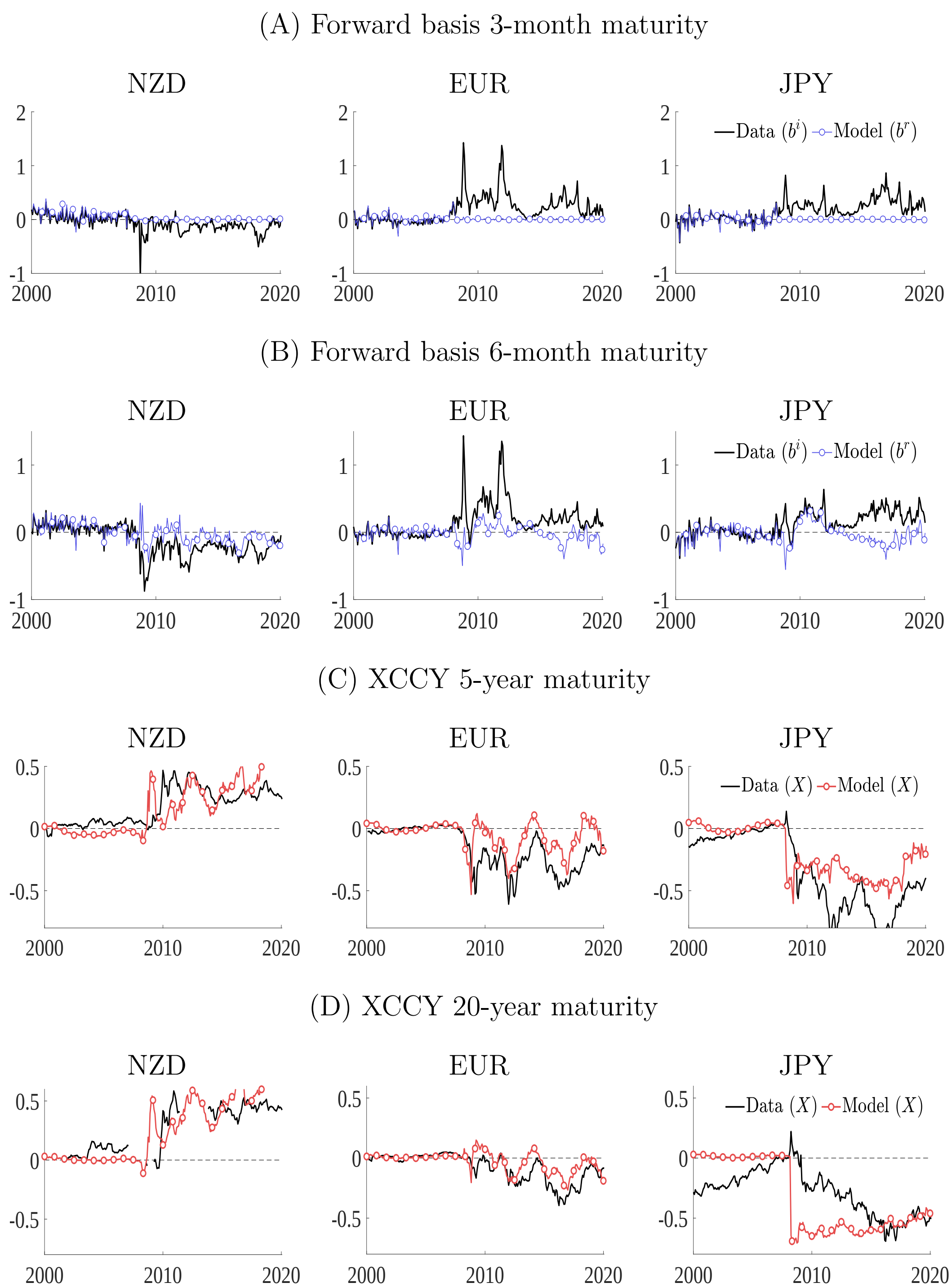

Figure 3: **Forward basis and XCCY rate.** We report the mean of the forward basis (based on LIBOR in the data and on EFR in the model) and the cross-currency basis swap rate (in bps). We also report the cross-sectional average of absolute rates, AVG. All exchange rates are expressed as the USD price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

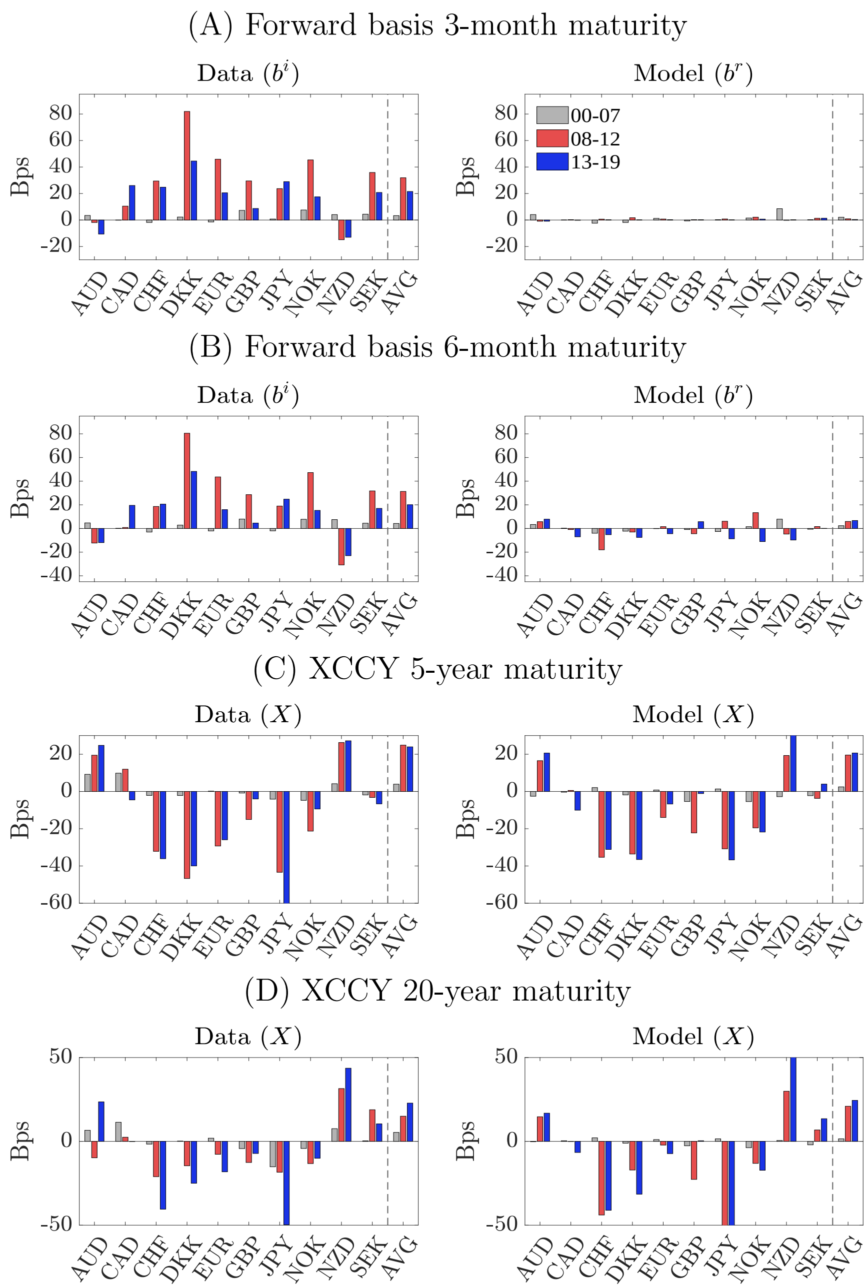

Figure 4: **Comparison of 1Y Interest Rate Proxies.** Each figure compares the model-implied 1-year interest rate to the predicted one and given by $\Delta r = -0.01 + 0.38 \cdot \Delta\text{LIBOR} + 0.50 \cdot \Delta(\text{Treasury} + \lambda) - 0.50 \cdot \Delta\text{CDS}$, where LIBOR corresponds 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.

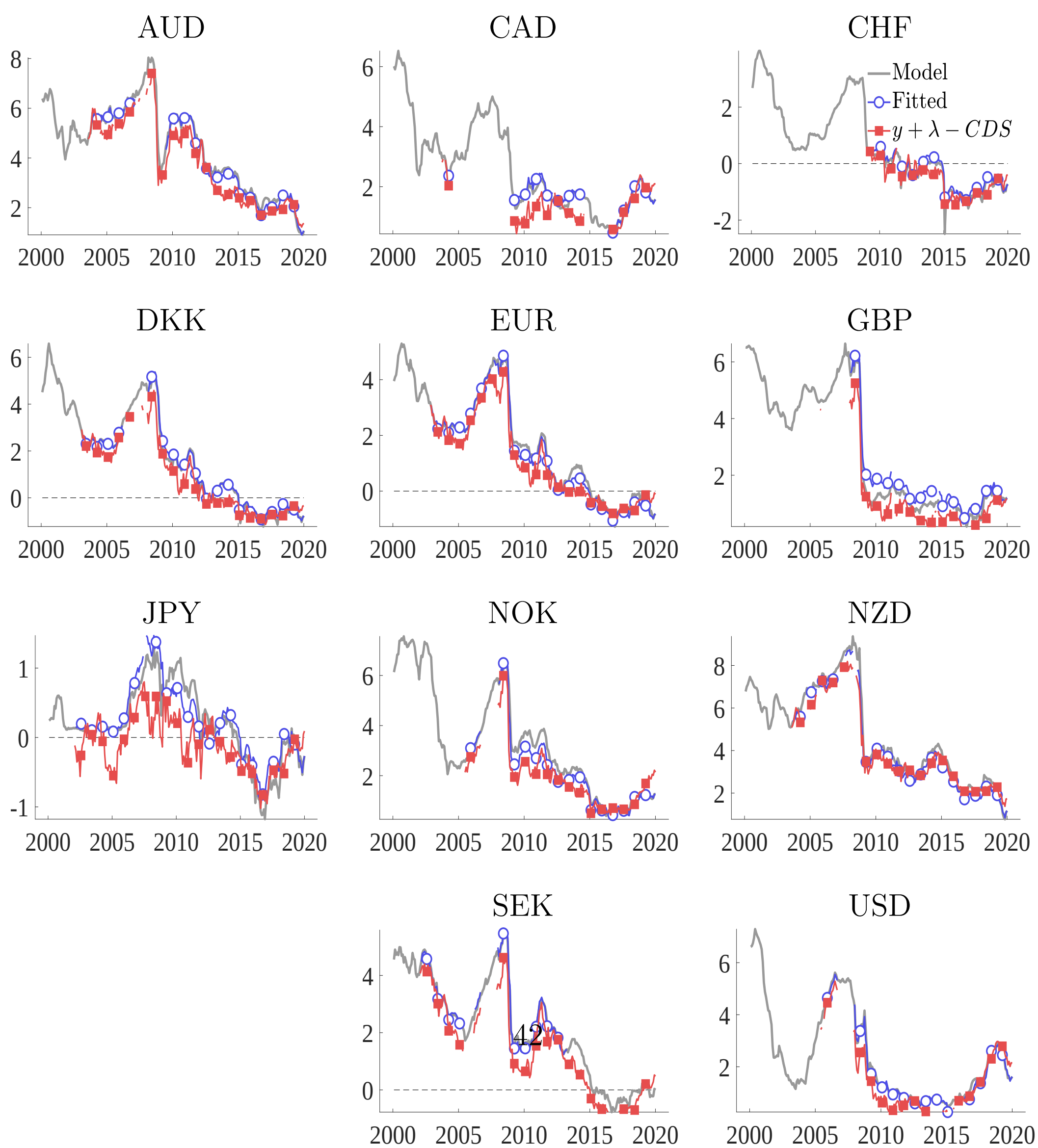

Figure 5: **Comparison of 3M Interest Rate Proxies.** Each figure compares the model-implied 3-month interest rate to the predicted one and given by $\Delta r = -0.00 + 0.49 \cdot \Delta\text{LIBOR}+0.51 \cdot \Delta(\text{Treasury}+\lambda)$, where LIBOR corresponds to the country-specific Libor/interbank 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.

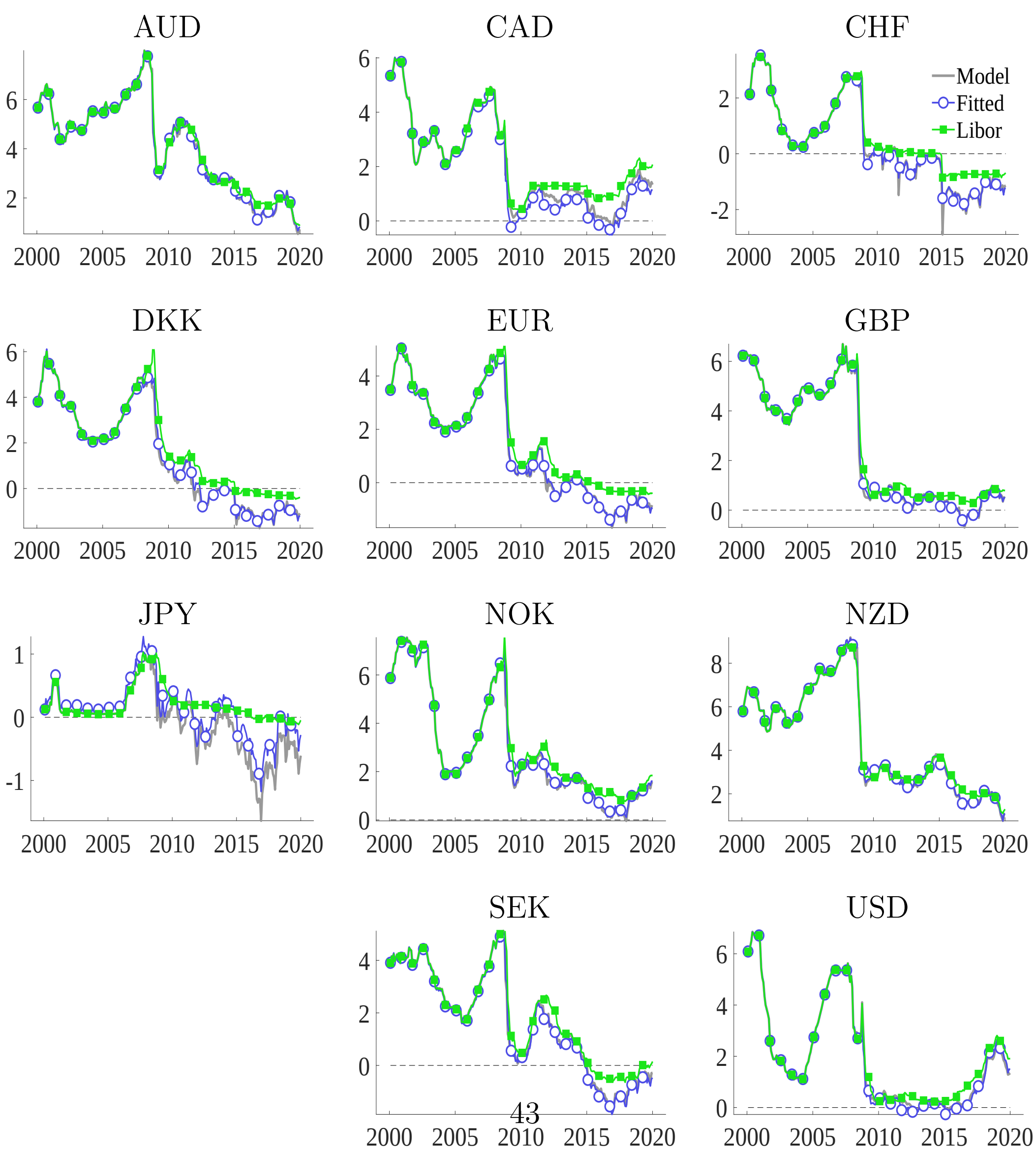

Figure 6: **Factor Exposure of XCCY Basis Swap Spread and Forward Premium Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, NOK, we regress changes in the spread between the observed and model-implied (i) 5-year XCCY basis swap rate and (ii) the 6-month forward premium on the Haddad and 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.

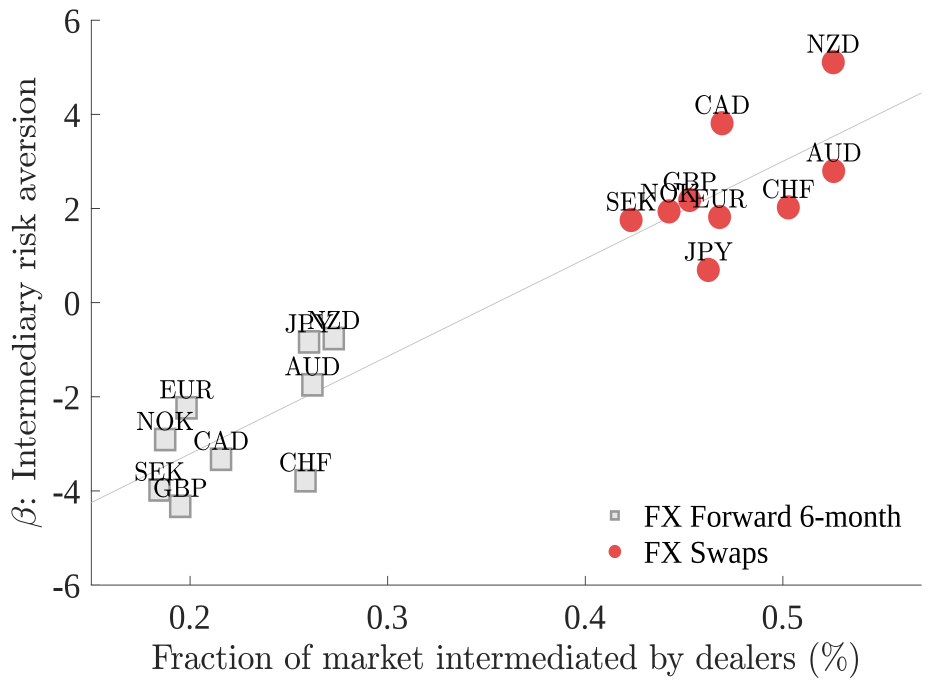

Table 1: Cash flows from a plain vanilla and fixed-for-fixed cross-currency basis swap

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 $i_t^* + X$ 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 $i_t^* + X$ 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 CMS^* respectively.

		$S = \$1/€1$	Cash flows at time		
		XC Basis Swap	0	t	T
Panel A	XC Swap	EUR Leg	$+ €1$	$- €(i_{t-1}^* + X)$	$- €(i_{T-1}^* + X) - €1$
		USD Leg	$- \$S_0$	$+ \$S_0 i_{t-1}$	$+ \$S_0 i_{T-1} + \$S_0$
Panel B	€ IRS	Fix Leg		$- €CMS^*$	$- €CMS^*$
		Float Leg		$+ €i_{t-1}^*$	$+ €i_{T-1}^*$
	\$ 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 cross-currency and interest swap rates - 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 ($XCCY^m$) 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
% Explained by Model	59.37	53.90	62.18	46.93	63.52	68.14	76.17	65.34	35.67	73.24	48.59
% Explained by Error	40.63	46.10	37.82	53.07	36.48	31.86	23.83	34.66	64.33	26.76	51.41

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. Reported standard errors are robust and adjusted for heteroscedasticity, but correcting standard errors for serial correlation or different types of clustering yields similar results. 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	0.58***												
	(0.04)												
OIS		0.75***											
		(0.17)											
IBOR			0.94***									0.49***	0.38***
			(0.05)									(0.05)	(0.09)
λ				0.31***									
				(0.05)									
Treasury+λ					0.80***							0.51***	0.50***
					(0.03)							(0.04)	(0.07)
EFFR						0.21***							
						(0.03)							
CD-Treasury							0.17***						
							(0.05)						
CDS								-0.06					-0.50***
								(0.11)					(0.09)
Treasury-CDS									0.52***				
									(0.05)				
Treasury+λ-CDS										0.70***			
										(0.06)			
BOX											0.37***		
											(0.06)		
Constant	-0.01***	-0.01	-0.00*	-0.02***	-0.01***	-0.02***	-0.02***	-0.02***	-0.01***	-0.01***	-0.01***	-0.00**	-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 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	0.317	0.323	0.618	0.086	0.661	0.072	0.043	-0.000	0.249	0.436	0.123	0.738	0.533
adj.R^2	0.643	0.641	0.801	0.523	0.823	0.071	0.042	0.474	0.605	0.703	0.123	0.863	0.737

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 ($\left|XCCY5y^D\right|$) 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 $+\lambda$), 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) $\left\|XCCY5y^D\right\|$	(2) $\left\|XCCY5y^D\right\|$	(3) $\left\|XCCY5y^D\right\|$	(4) $\left\|XCCY5y^D\right\|$	(5) $\left\|XCCY5y^D\right\|$
3M$(i-r)$	21.98***	28.34***	26.15***	25.96***	27.77***
	(1.41)	(1.99)	(2.09)	(2.54)	(3.28)
Treasury$+\lambda$			-0.84***	-1.09***	1.96**
			(0.27)	(0.34)	(0.77)
CDS				12.78*	7.03
				(6.50)	(4.49)
Constant	15.74***	13.18***	15.05***	15.09***	11.43***
	(0.62)	(0.80)	(1.00)	(1.17)	(1.56)
OBSERVATIONS	1,405	1,405	1,405	1,269	1,269
CCY FE	NO	NO	NO	NO	YES
MONTH FE	NO	YES	YES	YES	YES
MAT	3M	3M	3M	3M	3M
adj.R^2	0.175	0.182	0.185	0.195	0.766

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 XCCY basis swap rates ($\Delta XCCY^e$) on changes in the shadow cost of constraints measured as the difference between an uncollateralized (i.e., Libor) and collateralized (i.e., EFR) interest rate for the USD (ψ^{us3m}) or based on the first principal component of the shadow cost of constraints across all currencies (ψ^{pc3m}). Since the sign of a principal component is undetermined, we switch the sign on the coefficients of ψ^{pc3m} to align with that of ψ^{us3m}. We also control for other proxy candidates for shadow cost of constraints: the He, Kelly, and Manela (2017) intermediary capital ratio factor (HKM-ICR); the trade-weighted U.S. dollar index (USD Factor); the Bekaert-Horeova uncertainty (BH-UC); the VIX index (VIX); the certificate of deposit rate over Treasury yield spread (CD-Treasury); the Jurado, Ludvigson, and Ng (2015) financial uncertainty (JNL-FU12). All tenors for pricing errors are based on 5 year contracts. 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 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.

	(1) $\Delta XCCY^e$	(2) $\Delta XCCY^e$	(3) $\Delta XCCY^e$	(4) $\Delta XCCY^e$	(5) $\Delta XCCY^e$	(6) $\Delta XCCY^e$
ψ^{us3m}	-44.84***	-39.41***	-32.54***			
	(7.49)	(6.48)	(6.83)			
ψ^{pc3m}				-1.34***	-1.38***	-1.73***
				(0.26)	(0.25)	(0.25)
HKM-ICR		138.33***	336.90***		230.46***	517.87***
		(39.01)	(81.88)		(39.34)	(71.65)
BH-RA		-0.10***	-0.10**		-0.10**	-0.06
		(0.04)	(0.04)		(0.04)	(0.04)
BH-UC		-0.13***	-0.12**		-0.14***	-0.11**
		(0.04)	(0.05)		(0.05)	(0.05)
VIX		0.49***	0.57***		0.44***	0.46***
		(0.16)	(0.18)		(0.16)	(0.16)
CD-Treasury		1.71*	1.91*		2.65***	2.46**
		(0.91)	(1.06)		(0.97)	(1.03)
JNL-FU12		0.88	-0.65		-4.99	-8.51*
		(3.06)	(5.11)		(3.19)	(4.95)
USD FACTOR			-0.35***			-0.72***
			(0.12)			(0.12)
Constant	0.14	-0.77	0.72	-0.01	4.90	8.29*
	(0.11)	(2.94)	(4.87)	(0.12)	(3.07)	(4.71)
OBSERVATIONS	2,354	2,228	1,509	2,354	2,228	1,509
CCY FE	YES	YES	YES	YES	YES	YES
MONTH FE	NO	NO	NO	NO	NO	NO
MAT	5Y	5Y	5Y	5Y	5Y	5Y
w.R^2	0.078	0.105	0.126	0.053	0.101	0.167
adj.R^2	0.075	0.101	0.121	0.050	0.097	0.162

Table 7: Determinants of Shadow Cost of Capital

In this table, we report results from the regression of changes in the 3-month shadow cost of capital $\Delta\psi$ (Libor minus model-implied effective funding rate EFR) on changes of proxy candidates for shadow cost of capital. The shadow cost of capital is measured either using the USD 3-month LIBOR-EFR (ψ^{usd3m}) or using the first principal component of the shadow cost of constraints across all currencies (ψ^{pc3m}). We report only those coefficients that remain statistically significant in univariate regressions: the Ted spread (TED); the He, Kelly, and Manela (2017) intermediary capital ratio factor (HKM-ICR); the certificate of deposit rate over Treasury yield spread (CD-Treasury); the trade-weighted U.S. dollar index (USD Factor); the Bekaert-Horeova uncertainty (BH-UC); the Jurado, Ludvigson, and Ng (2015) macroeconomic uncertainty (JNL-MU12). Standard errors are robust. We report the adjusted R^2 values from the panel regressions. The sample period is January 2000 to December 2019.

	(1) $\Delta\psi^{usd3m}$	(2) $\Delta\psi^{usd3m}$	(3) $\Delta\psi^{usd3m}$	(4) $\Delta\psi^{usd3m}$	(5) $\Delta\psi^{usd3m}$	(6) $\Delta\psi^{pc3m}$	(7) $\Delta\psi^{pc3m}$	(8) $\Delta\psi^{pc3m}$	(9) $\Delta\psi^{pc3m}$	(10) $\Delta\psi^{pc3m}$
TED	-0.03**									
	(0.01)									
HKM-ICR		-2.24***			-2.20***					
		(0.70)			(0.63)					
CD-Treasury			-0.04**		-0.04**	0.61*				
			(0.02)		(0.02)	(0.35)				
USD FACTOR				0.00*			0.20*			0.15*
				(0.00)			(0.11)			(0.09)
BH-UC								0.02*		
								(0.01)		
JNL-MU12									40.75*	38.54*
									(20.92)	(19.72)
Constant	0.00	0.00	0.00	0.00	0.00	0.03	0.01	0.03	0.03	0.03
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.07)	(0.09)	(0.07)	(0.07)	(0.09)
OBSERVATIONS	240	227	240	167	227	240	167	239	240	167
MAT	3M	3M	3M	3M	3M	3M	3M	3M	3M	3M
adj.R^2	0.027	0.052	0.062	0.020	0.117	0.021	0.076	0.067	0.070	0.119