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ABSTRACT

We revisit the recent literature on persistent deviations from covered interest parity (CIP) by showing theoretically that CIP violations imply arbitrage opportunities only if uncollateralized interbank lending rates are riskless. In the absence of observable riskless discount rates, we extract them empirically using a simple no-arbitrage framework. They deliver novel quantitative benchmarks for foreign exchange contracts that match observed forward currency premiums and cross-currency basis swap rates well. The no-arbitrage benchmarks account for about two thirds of the alleged CIP deviations, while the residual pricing errors line up with measures of intermediary constraints and the expensiveness of the U.S. dollar.

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1 Introduction

With an average daily turnover of \$6.6 trillion (BIS, 2019), the foreign exchange market is among the most liquid trading venues in the world. Fluctuations in exchange rates are intimately related to fluctuations in sovereign default risk (Du and Schreger, 2016; Na, Schmitt-Grohé, Uribe, and Yue, 2018), and foreign exchange reserves represent a primary safety cushion for governments around the world (Bianchi, Hatchondo, and Martinez, 2018). Thus, it is no surprise that scores of researchers have studied the determination of exchange rates and their implications (e.g., Engel, 2016; Maggiori, 2017; Kremens and Martin, 2019; Lustig and Verdelhan, 2019; Lustig, Stathopoulos, and Verdelhan, 2019; Gopinath, Boz, Casas, Diez, Gourinchas, and Plagborg-Moeller, 2020).

One of the pillars of how we think about the international currency system is that of covered interest parity (e.g., Keynes, 1923). According to covered interest parity (henceforth CIP), the cross-currency basis, that is, the difference between the forward premium relative to the spot exchange rate and the cross-country interest rate differential, should be equal to zero. At longer horizons above one year, CIP is characterized through zero rates of cross-currency basis swaps (xccy). These are foreign exchange derivative instruments that allow for the exchange of two variable rate loans denominated in different currencies.

A growing and important literature provides evidence of persistent CIP violations, at least since the 2008 Global Financial Crisis (GFC) (e.g., Ivashina, Scharftstein, and Stein, 2015). Figure 1 offers an example of the evidence of non-zero basis and xccy rates in the case of the Euro, using LIBOR as a proxy for the riskless rates, following standard practice in the literature. Departures from zero for both time series are taken as prima facie evidence that CIP is violated. Notably, CIP violations suggest the existence of arbitrage opportunities (e.g., Du, Tepper, and Verdelhan, 2018), which in turn violates the backbone of our understanding of financial markets. Prominent explanations include a combination of leverage and funding constraints associated with the expensiveness of the U.S. dollar (Ivashina, Scharftstein, and Stein, 2015; Avdjiev, Du, Koch, and Shin, 2019) and balance sheet constraints of financial intermediaries (Du, Tepper, and Verdelhan, 2018; Andersen, Duffie, and Song, 2019; Fleckenstein and Longstaff, 2020).

We argue that the key to understanding the behavior of CIP is the nature of effective funding rates (EFRs) faced by market participants. In stark contrast to conventional wisdom, we show, in a model-free setting, that, under the null hypothesis of no arbitrage, both the basis and xccy rates cannot be simultaneously zero unless the EFR is equivalent to LIBOR. Intuitively, because both forward and xccy transactions are fully collateralized, the EFR should be free of counterparty risk. Thus, given the possibility of distress in uncollateralized interbank markets, proxying for the EFR by LIBOR is empirically questionable. This observation prompts a question about the appropriate quantitative benchmark for testing the CIP relation.

Tests of the short-term CIP validity depend on the availability of EFRs or riskless interest rates. Standard practice in the literature is to proxy for these rates by Treasury yields, or, bank funding rates, such as LIBOR, because financial institutions are the ones most likely to enforce no-arbitrage relations by trading on price discrepancies.[1] However, there is significant evidence that Treasury yields may be too low, because they provide safety and convenience (e.g., Bansal and Coleman, 1996; Krishnamurthy and Vissing-Jorgensen, 2012; Nagel, 2016), giving rise to a non-zero Treasury-based forward exchange rate basis (Du, Im, and Schreger, 2018; Jiang, Krishnamurthy, and Lustig, 2018). In addition, there is significant evidence that bank funding rates may be too high because they can embed default risk, which could deliver a non-zero LIBOR-based forward basis (Du, Im, and Schreger, 2018; Du, Tepper, and Verdelhan, 2018). An additional complication for obtaining representative riskless rates is that counterparties face differential funding costs depending on their credit risk and bargaining power. Thus, we take the view that true risk-free rates are unobservable.

We propose a novel no-arbitrage benchmark by applying a standard stochastic discount factor approach to characterize present values of cash flows associated with foreign exchange contracts. We respect the fact that both short-term forward and long-term xccy contracts are free of counterparty risk because of collateralization. This implication is central to our perspective on the CIP anomaly, because the basis constructed using the corresponding risk-free rates is zero, while the xccy rate is not.

We operationalize our framework to back out the cost of safe and riskless assets. We posit that their dynamics are part of an affine no-arbitrage model, and use data for G11 countries from January 2000 to December 2019. In a first step, we extract the latent risk-free rates in each country by imposing the 3-month CIP in terms of the model-based risk-free rates and by matching the term structure of IRS rates in the respective countries. In a second step, we relate the estimated latent risk-free rates to various observable candidates. We find that only three variables are significant. The risk-free rates are closely associated with Treasury yields adjusted for convenience, and sovereign credit default swap (CDS) rates. In addition, there is a modest contribution from interbank rates, i.e., LIBOR. Thus, the estimated risk-free rate is a linear combination of LIBOR, a Treasury rate adjusted for convenience, and credit risk.

[1] We refer to interbank rates as LIBOR, regardless of the currency of denomination.

We next use the riskless rates implied from our model to revisit the evidence on long-term CIP deviations characterized by non-zero cross-currency swap rates. We compute model-implied xccy rates, which were not used in the estimation. While the basis is zero at horizons of one year and less, we find that xccy rates are non-zero, just like in the data. The average 5-year xccy rate across countries during the post-crisis period is 24 basis points (bps) vs. 21 bps in the model. Our no-arbitrage model would furthermore imply zero Treasury basis and positive swap spreads. That is because in our model, there is no difference between risk-free and Treasury rates.

Some variation in xccy rates remains unexplained by our model. We show that these xccy pricing errors reflect market frictions that correlate with financial intermediary health. We find that the leverage of bank holding companies (He, Kelly, and Manela, 2017) and the trade-weighted U.S. dollar index (Avdjiev, Du, Koch, and Shin, 2019; Jiang, Krishnamurthy, and Lustig, 2018) are two highly significant variables co-varying with changes in the xccy pricing errors. Cross-sectional regressions show that expected changes in the cross-currency pricing errors line up with their beta sensitivities to both factors. Lastly, we implement the Haddad and Muir (2018) cross-sectional analysis of intermediation costs and assets' exposure to a measure of intermediary risk aversion. We find that swaps are intermediated more than forward contracts, and that their risk aversion sensitivities line up accordingly. The overall evidence is consistent with the role of intermediaries in xccy premiums.

We evaluate the relative contribution to the variation in xccy rates by our model and the significant intermediary fractions in the spirit of Campbell and Ammer (1993). Our variance-covariance decomposition on the 5-year xccy rates shows that our no-arbitrage framework explains, on average, about 68% of the variation in their levels. The remaining 32% of their variation that is left unexplained is accounted for by financial intermediary leverage and the expensiveness of the USD.

Taken at face value, our findings lead to a novel and important conclusion. A simple no-arbitrage framework with the most minimal set of assumptions is helpful for understanding, to a first order, the dynamics of the forward-spot exchange rate basis and xccy rates. Our analysis suggests that non-zero and large xccy rates do not necessarily imply arbitrage opportunities. In fact, our no-arbitrage benchmarks provide a good quantitative account of the levels and dynamics of these rates. Nonetheless, we associate our pricing errors with factors that are likely correlated with balance sheet costs (Du, Tepper, and Verdelhan, 2018; Andersen, Duffie, and Song, 2019; Fleckenstein and Longstaff, 2020). That result lends credence to the role of intermediary based asset pricing for quantitatively realistic models (He, Krishnamurthy, and Milbradt, 2019).

Related literature

Our work relates to the growing literature on international asset pricing anomalies in developed economies documented since the GFC. One prominent stream of research focuses on the failure of covered interest parity (CIP), a no-arbitrage condition that equates the interest rate differential in two countries with the corresponding forward-spot exchange rate premium. Since the GFC, there is increasing evidence that the CIP condition, implied by interbank interest rates such as LIBOR, has been violated. The evidence suggests that borrowing costs in USD cash markets are lower than those implied by comparable synthetic USD loans constructed using foreign currency loans, spot and forward exchange rates. Taken at face value, this suggests the presence of persistent and systematic arbitrage opportunities (Du, Tepper, and Verdelhan, 2018).

In terms of possible explanations of CIP deviations, early work points towards frictions in global intermediation of USD funding. See, among others, Baba, Packer, and Nagano (2008), Coffey, Hrung, and Sarkar (2009), Griffolli and Ranaldo (2011), McGuire and von Peter (2012), and Bottazzi, Luque, Pascoa, and Sundaresan (2012). Ivashina, Scharftstein, and Stein (2015) provide a model that rationalizes these effects through foreign banks that cut USD lending more than U.S. banks in crisis times. Coffey, Hrung, and Sarkar (2009) and Bahaj and Reis (2018) support such views by documenting that U.S. swap lines to foreign central banks are effective in reducing CIP deviations. Other explanations relate to an increase in bank counterparty risk associated with the crisis (Tuckman and Porfirio, 2003; Baba and Packer, 2009; Skinner and Mason, 2011; Levich, 2012; Csavas, 2016; Wong, Leung, and Ng, 2016; Alfred Wong, 2018), while Coffey, Hrung, and Sarkar (2009) and Fong, Valente, and Fung (2010) provide evidence in support of both credit and liquidity risk as drivers of CIP deviations. Garleanu and Pedersen (2011) point towards binding margin constraints. Additional contributions are made by Akram, Rime, and Sarno (2008), Goldberg, Kennedy, and Miu (2010), and Iida, Kimura, and Sudo (2016).

CIP deviations have persisted even after funding conditions became less strained in the post-crisis period, which is often associated with a reduced ability to conduct arbitrage in response to post-crisis regulation. Du, Tepper, and Verdelhan (2018) document a co-movement in CIP deviations with bank balance sheet constraints (see also Andersen, Duffie, and Song (2019) and Fleckenstein and Longstaff (2020) for related arguments.) Similarly, Cenedese, Corte, and Wang (2020) point towards post-crisis leverage constraints as a source of CIP deviations. Borio, McCauley, McGuire, and Sushko (2016) argue that an increase in hedging demand coupled with reduced balance sheet capacity is responsible for persistent deviations. See also Liao (2016) for arguments of hedging demand. Anderson, Du, and Schlusche (2019) indicate that banks' arbitrage positions reduced in response to a reduction in wholesale funding supply following a regulatory change in the money market mutual fund industry in 2016. Avdjiev, Du, Koch, and Shin (2019) cite banks' leverage constraints as a source of friction, driven by the strength of the dollar. In that same spirit, Du, Hebert, and Wang (2019) argue that CIP deviations may reflect financial intermediaries' shadow costs of capital constraints.[2]

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. They conclude, on the basis of these costs, that CIP arbitrage is possible for only a subset of highly capitalized banks. The potential arbitrage profits are, however, much smaller than what LIBOR-based measures would indicate. Andersen, Duffie, and Song (2019) question benefits of CIP arbitrage to bank shareholders in the light of required funding value adjustments.

One common feature across previous work is that CIP deviations are measured using observable interest rates. Our key distinction is that we consider true discount rates to be unobservable when we analyze CIP deviations from the no-arbitrage perspective. Moreover, we provide new testable predictions for long-term CIP deviations suggested by non-zero prices of cross-currency basis swaps.

Another stream of research focuses on CIP deviations measured using Treasury yields, giving rise to a non-zero Treasury basis. Du, Im, and Schreger (2018) document the Treasury basis across G11 and emerging countries and relate it to relative differences in convenience yields. Jiang, Krishnamurthy, and Lustig (2018, 2019) formalize the greater convenience yield for USD assets based on foreign investors' demand for safe assets. The two studies differ in their measurement of the Treasury basis, with the former relying on fitted par yield curves and the latter using secondary market bond prices. By studying the Treasury basis, we also relate broadly to the literature that examines the convenience yield embedded in Treasury bonds (Bansal and Coleman, 1996; Grinblatt, 2001; Krishnamurthy, 2002; Longstaff, 2004; Gurkaynak, Sack, and Wright, 2007; Goyenko, Subrahmanyam, and Ukhov, 2011; Krishnamurthy and Vissing-Jorgensen, 2012; Nagel, 2016).

A third pricing anomaly that appeared since the GFC is the emergence of negative swap spreads, i.e., interbank interest swap rates have fallen below maturity-matched Treasury yields. The extant evidence almost exclusively focuses on the U.S., and cites reasons related to hedging demand (Klingler and Sundaresan, 2019), dealer funding costs (Lou, 2009), increases in regulatory leverage ratios (Boyarchenko, Gupta, Steele,

[2]Discussions and test of CIP deviations go back to Keynes (1923). For early work, see also Frenkel and Levich (1975); Dooley and Isard (1980); Taylor (1987, 1989); Popper (1993); Fletcher and Taylor (1994, 1996).

and Yen, 2018; Jermann, 2020), a declining convenience yield, (Klingler and Sundaresan, 2018), or U.S. default risk (Augustin, Chernov, Schmid, and Song, 2020). Du, Im, and Schreger (2018) relate CIP deviations to cross-country swap spread differentials. We provide formal cross-country evidence on the dynamics and cross-sectional differences in negative swap spreads.

Another major difference from prior work is that we study arbitrage anomalies across multiple asset classes in a relative valuation sense. From that perspective, we also relate closely to Pasquariello (2014), who studies commonalities across market dislocations in stock, foreign exchange, and money markets. Hazelkorn, Moskowitz, and Vasudevan (2020) suggest that demand for leverage explains arbitrage anomalies across foreign exchange and equity markets.

2 A no-arbitrage perspective

We start by presenting our assumptions. We then successively present the no-arbitrage valuation approach to forward and xccy rates. We end by discussing our empirical strategy.

2.1 Assumptions

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

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

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

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

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

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

If international financial markets are integrated, these investment strategies denominated in Euros are available to a U.S. investor as well, but she would have to convert U.S. dollars to Euros first, and then back. Let S_t denote the USD value of EUR 1 at time t. In a setting without arbitrage opportunities, this implies, as emphasized by Backus, Foresi, and Telmer (2001), that

$$E_0\left[M_{0,T}(S_T/S_0)\widehat{R}_{0,T}\right] = 1,$$

so that

$$E_0\left[\widehat{M}_{0,T}\widehat{R}_{0,T}\right] = E_0\left[M_{0,T}(S_T/S_0)\widehat{R}_{0,T}\right].$$

This equality establishes a tight connection between movements in exchange rates, and the stochastic discount factors $\widehat{M}_{0,T}$, and $M_{0,T}$.

2.2 Forward rates

A currency contract that is struck at time 0 to sell €1 forward at time T for the price $\$F_{0,T}$ has a net USD cash flow of $F_{0,T} - S_T$. A textbook forward valuation strategy would involve two portfolios. One portfolio is long USD: sell €1 forward and invest $F_{0,T}\exp(-Tr_{0,T})$ at the annualized domestic $T-$period "risk-free rate" $r_{0,T}$. The second portfolio is short EUR: borrow € $\exp(-T\widehat{r}_{0,T})$ at the annualized foreign $T-$period "risk-free rate" $\widehat{r}_{0,T}$. The two portfolios have identical payouts at maturity, thereby making their time-0 values expressed in USD the same:

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

implying that

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

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

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

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

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

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

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

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

The forward price in logs is given by

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

2.3 Cross-currency basis swap rates

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

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

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

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

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

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

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

As a next step we link the xccy cash flows to those of interest rate swaps (IRS). Specifically, we swap both the USD and the EUR interest rates into fixed rates using an IRS in each currency, at prices CMS and $\widehat{CMS}$, respectively (CMS stands for "constant maturity swaps"). We illustrate these cash flows in Table 1B.

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

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

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

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

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

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

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

2.3.1 Cross-sectional variation in xccy rates

In order to understand the drivers behind the xccy rate X, divide the expression in (5) by $\widehat{CMS}_{0,T}$ and add 1:

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

Now let's introduce a hypothetical par bond (constant maturity Treasury, CMT) that corresponds to the discount factor M'. Its coupon rate is:

$$CMT'_{0,T} \equiv \frac{1 - E_0\left[M'_{0,T}\right]}{\sum_{t=1}^{T} E_0\left[M'_{0,t}\right]}.$$

A foreign hypothetical bond has a similar notation. Then,

$$\begin{aligned} 1 + \frac{X_{0,T}}{\widehat{CMS}_{0,T}} &= \frac{CMS_{0,T}}{CMT'_{0,T}} \cdot \left(\frac{\widehat{CMS}_{0,T}}{\widehat{CMT}'_{0,T}}\right)^{-1} \cdot w_0 + \left(\frac{\widehat{CMS}_{0,T}}{\widehat{CMT}'_{0,T}}\right)^{-1} \cdot (1 - w_0) \quad (6) \\ &\approx \frac{CMS_{0,T}}{CMT'_{0,T}} \cdot \left(\frac{\widehat{CMS}_{0,T}}{\widehat{CMT}'_{0,T}}\right)^{-1}, \quad (7) \end{aligned}$$

with weight $w_0 \equiv (1 - E_0[M'_{0,T}]) \cdot (1 - E_0[\widehat{M}'_{0,T}])^{-1}$. The approximation is based on the assumption $w_0 \approx 1$. We use this approximation for expositional purposes and for cross-sectional analysis.

The first relation (equality) implies that $\widehat{CMS}_{0,T} = \widehat{CMT}'_{0,T}$ and $CMS_{0,T} = CMT'_{0,T}$, then $X_{0,T} = 0$. Otherwise, we should expect to see time-series and cross-sectional variation in xccy rates. The second relation (approximation) implies that cross-sectional variation in the appropriately normalized xccy rate X is driven by the differences in the domestic and foreign interest rates. Specifically, the sign of X is driven by the cross-country differences in swap spreads, defined here as ratios of of swap rates and (hypothetical) par bond rates.

As a result, we link cross-sectional variation in X to another post-GFC phenomenon referred to as negative swap spreads. The evidence is that post-GFC in the U.S., $CMS_{0,T} - CMT_{0,T} < 0$, or, equivalently, $CMS_{0,T}/CMT_{0,T} < 1$ for large T, with the same results for its foreign counterparts. Equation (7) links xccy to swap spreads relative to the hypothetical bonds. Thus, we can evaluate whether this distinction is important and whether the explanation of the swap spread is connected to the appropriate discount factor.

2.3.2 Discounting at LIBOR

There is one prominent case that leads to a zero xccy rate. This is the case of discounting at LIBOR, which was a prominent paradigm prior to the GFC (e.g., Duffie and Singleton, 1997). To see this, note that in that case $E_0\left[M'_{0,T}\right] = 1/(1+i_{0,T})^T$. As a result,

$$CMS_{0,T} = \frac{1 - E_0\left[M'_{0,T}\right]}{\sum\limits_{t=1}^{T} E_0\left[M'_{0,t}\right]}. \tag{8}$$

The foreign swap rate $\widehat{CMS}$ obtains a similar expression. That is, $CMS = CMT'$ and $\widehat{CMS} = \widehat{CMT}'$ in Equation (6).

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, and, by the end of 2008, the whole industry had switched to OIS (e.g., Hull and White, 2013, Cameron, 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 implied by (5) is quantitatively similar to the observed one.

2.3.3 Back to forward rates

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

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

The last expression can be re-written in terms of LIBOR basis (in levels):

$$\frac{F_{0,T}}{S_0}\frac{(1+\widehat{i}_{0,T})^T}{(1+i_{0,T})^T} = 1 - \frac{X_{0,T}}{F_{0,T}}(1+i_{0,T})^{-T}.$$

If $E_0[M_{0,T}] = (1+i_{0,T})^{-T}$, then the LIBOR basis is equal to 1, and $X_{0,T} = 0$, as in the previous subsection where we explored discounting at LIBOR. Otherwise, the log

LIBOR basis is equal to:

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

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

2.4 Empirical strategy

In order to evaluate the quantitative success of our view of forward premiums and xccy rates, we need to obtain estimates of collateral-adjusted SDFs M' and $\widehat{M}'$. Our strategy is to infer these objects from prices of domestic and foreign IRS. We do so, on a country-by-country basis, by estimating an affine term structure model designed to fit the interbank rates and the corresponding IRS curve of a given country. Having obtained a collection of M' and $\widehat{M}'$, we can evaluate expressions in Equations (3) and (5) and compare them to the observed forward premiums and xccy rates, respectively. Importantly, we are not using the data on $X_{0,T}$ for the estimation. We exploit the identity $b^r_{0,T} = 0$ to identify foreign rates $\widehat{r}$ via the observed forward premiums.

2.4.1 A model

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

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

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

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

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

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

While we are using the interest rate r_t in the same way as a risk-free rate appears in a classical framework, we have to be careful with its interpretation. We cannot estimate the true risk-free rate using the data on OTC interest rate derivatives alone. What we estimate is the effective funding rate in these markets. It is reasonable to think of this rate as risk-free because the specific instruments that we are using are fully collateralized. A different effective funding rate could be associated with different markets. See, for example, Binsbergen, Diamond, and Grotteria (2019) for equity options, and Fleckenstein and Longstaff (2018) for Treasury note futures. We refer to r_t as the risk-free rate and emphasize its interpretation when appropriate.

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

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

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

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

$$CMS_{0,T} \;=\; \frac{\sum\limits_{t=1}^{T} E_0\left[M'_{0,t}\left(e^{i_{t-1,t}}-1\right)\right]}{\sum\limits_{t=1}^{T} E_0\left[M'_{0,t}\right]}. \tag{9}$$

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.

2.4.2 Identification

As specified, the model is under-identified. We adopt the canonical form used by Joslin, Le, and Singleton (2013) and choose the latent state z_t so that the matrix $\Phi - \nu$

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

governing the dynamics under the risk-adjusted distribution is diagonal. Further, because both loadings δ_r and covariance matrix Σ control the scale of r_t, we set the former to unity. All other parameters are free.

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

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

We rely on the second observation and set $\widehat{r}'_{0,T} = r'_{0,T} - \rho_{0,T}$ for $T = 3$ months. In combination with the previous identifying assumption, this helps in identifying the connection between the domestic and foreign interest rates. It also helps with separating r and r', or, equivalently, η in each country. Finally, the third observation, in combination with the other two assumptions, identifies r and η separately.

3 Evidence

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

3.1 Data

We use a panel data set on interest and exchange rates for G11 countries from January 2000 to December 2019. G11 currencies include the USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, NOK.[4] Specifically, we obtain information on spot

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

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

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

The black lines in Figure 4 show spreads between IRS and constant maturity Treasuries (CMT) of matching maturities at the 20-year horizon (NOK and NZD are not available, AUD and SEK have only partial data, and, thus, not reported). The USD pattern is prominently discussed in the literature. We observe interesting cross-country variation in these spreads.

3.2 Results

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

3.2.1 Forward bases

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

3.2.2 Xccy rates

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

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

3.2.3 Swap spreads

Figure 4 shows that swap spreads are no longer negative if Treasuries, CMT, are replaced by the hypothetical ones computed from our advocated discount rates, CMT'.[5]

[5]Although there little to no data for NOK, NZD, AUD, and SEK, we can compute the theoretical spreads relative to CMT'. All of them are positive. The DKK observed and theoretical spreads look

Augustin, Chernov, Schmid, and Song (2020) also use a no-arbitrage model to explain negative U.S. swap spreads. Their explanation is based on credit risk of the U.S. Treasury. Our explanation is silent about this specific channel because we are sidestepping the modeling of government bonds. Nevertheless, we can explore the role of credit risk by connecting yields on the hypothetical bonds to the ones on actual bonds.

Figure 5 explores the cross-sectional impact of swap spreads on xccy rates via Equation (7). The two top panels plot the relation between the xccy rate and the spread in swap spreads when CMT is the benchmark for 5 and 20-year maturities. We see that the relation is weak to negative. However, when we replace CMT with CMT', as our framework requires, we get a strong positive relation between the two. Thus, the differences in xccy rates are indeed determined by the "spread in spreads", i.e., cross-country differences in how swap rates differ from their riskless counterparts.

4 Interpretation of the evidence

This section has two main objectives. First, we relate the estimated risk-free rate to observable variables. Second, we evaluate economic sources of xccy pricing errors.

The first objective has a dual purpose: investigating the source of the empirical success of our model and developing economic intuition for the estimated risk-free rate. One might worry that our results are driven by the latent cost of collateral η, which mechanically adjusts LIBOR, as a proxy for r, so that $r' = r - \eta$ prices assets correctly. That we set η to a constant should alleviate this concern as r rather than r' is doing all the work in our model. Nevertheless, we can explicitly verify if the estimated r is closely associated with LIBOR. More broadly, it is useful to understand the structure of effective funding rates in the fixed-income swap markets. Thus, one would want to evaluate which variables r is related to. For that purpose, we use panel regressions that are informative about which variables correlate contemporaneously most strongly with r.

The second objective parallels our discussion of the Treasury basis and negative swap spreads. While our model does not speak to these concepts directly, it makes it easier to understand the economic sources behind these phenomena. Likewise, we interpret xccy pricing errors as a manifestation of economic phenomena not explicitly accounted for by our model.

very similar to the EUR ones.

4.1 Risk-free rate

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

Yields on Treasury bonds, $y_{0,T}$, continue to serve as a natural starting point to think about 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; Li and Song, 2019).[6] Having obtained the U.S. convenience premium λ, we obtain the foreign $\widehat{\lambda}$ from the Treasury basis via

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

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

Because reliable CDS information is available only at maturities starting at 1 year, the shortest interest rate that we can evaluate is for $T = 1$ year. Figure 6 plots $r_{0,T}$ and its proxy $\widetilde{r}_{0,T}$. We see that the proxy is tracking the risk-free rate quite well, but there are also evident departures. Japan is the strongest example of large discrepancies, and Sweden is one of the better fitting ones. While there is a reasonably close association between $\widetilde{r}$ and the model-implied interest rate, differences between them are not surprising. Even if there is no noise associated with the ingredients of

[6] 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.

$\widetilde{r}$, it does not account for risk associated with the interbank market, and so it may not be capturing the effective funding rates of dealers.

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

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

When evaluating the univariate regressions, we focus on the magnitude of the estimated coefficient (the closer to 1 the better) and the within R^2. The leading variables here are LIBOR and the convenience-adjusted Treasury with coefficients around 0.6, and R^2 around 0.5. The weakest variables are the U.S.-only ones: EFFR, CD-Treasury spread, and the option box spread with coefficients below 0.2 and R^2 below 0.06. Our initial proxy for the risk free rate $\widetilde{r}$ occupies an intermediate position with the coefficient of 0.48 and R^2 of 0.33.

Moving to multivariate regressions, we find that, at the 3-month horizon, LIBOR and convenience-adjusted Treasury rates are the two variables that remain significant. This finding is supportive of our prior that interbank funding costs should also be related to effective funding rates in swap markets. Initially, we have allowed y and λ

to appear separately, but the estimated coefficients were nearly identical, so we have combined them into one intuitive variable with no loss in R^2. We had also included the other candidate variables in the multivariate regression, but we subsequently removed them because they turned out to be statistically insignificant.

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

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

Figures 6 and 7 compare the estimated r with the best prediction according to the multivariate regressions presented in columns (12) and (13) in Table 3. The predictions are for the changes, so we obtain predictions for levels by cumulating the changes. At the 1-year horizon, the predicted r is more accurate than $\widetilde{r}$ and, in fact, is very close to r. At the 3-month horizon, the prediction tracks r almost perfectly.

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

4.2 Is the risk-free rate different from LIBOR?

As mentioned earlier, one concern could be that LIBOR is in fact a good proxy for the effective funding rate and, thus, all the explanatory power in the model is driven by the latent cost of collateral. First, we can see from column (3) of Table 3 that the regression coefficient of Δr on Δi is significantly different from one. Next, Figure 7 compares explicitly our risk-free rate with LIBOR. It is evident that i is substantively different from r during the post-crisis period (we set them to be similar before the crisis as part of our identification strategy). Thus, the difference in forward basis and xccy valuation comes from a different funding rate.

As a further characterization of the difference between the two rates, we consider theoretical connection between this difference and swap rates X. The no-arbitrage framework suggests that xccy rates are zero only under the strong assumption that risk-free rates are 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 effective funding rates.

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 4.

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

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

4.3 Xccy pricing errors

We have proposed a descriptive model that does not account for any frictions. While our fit of xccy rates is reasonable, there are pricing errors. We interpret the xccy pricing errors as potential components of risk premiums associated with intermediary constraints or other frictions that could be reflected in asset prices. This motivates us to investigate whether there is a covariation between xccy pricing errors and various measures of financial frictions that have been developed in the literature.

We consider three broad groups of variables. First, we measure intermediary constraints using leverage of security broker-dealers (Adrian, Etula, and Muir, 2014, AEM) and of bank holding companies (He, Kelly, and Manela, 2017, HKM), the trade-weighted U.S. dollar index, which proxies for the limited willingness of intermediaries to provide USD funding and demand for USD associated with the convenience of USD assets (Avdjiev, Du, Koch, and Shin, 2019; Jiang, Krishnamurthy, and Lustig, 2018), and total dealers' cash balances with the Federal Reserve. We do not report results for the latter variable because it turns out that their coefficients are statistically insignificant. Second, we use a number of measures of uncertainty: the Jurado, Ludvigson, and Ng (2015) real uncertainty, macroeconomic uncertainty, and financial uncertainty measures; the Bekaert and Hoerova (2014) uncertainty and risk aversion measures; and the CBOE VIX index. We also examine country-specific measures of uncertainty, including the Baker, Bloom, and Davis (2016) economic policy uncertainty indices, implied volatility from 5x10-year swaptions, and xccy bid-ask spreads. As none of these country-specific measures are significant, we do not report their results. Third, we use indicators of distress in the banking sector via the U.S. Treasury over Eurodollar (TED) spread, and the LIBOR-OIS spread. The latter is available for all countries. Details on all data sources are available in the online appendix.

Table 5 summarizes the relation between changes in xccy pricing errors and changes in candidate variables by regressing the former on the latter at the monthly frequency. The AEM is an exception as it is available at the quarterly frequency only. We run panel regressions and examine the connection of changes in xccy pricing errors to the individual variables, and to all of them together in a multivariate setting (with the exception of AEM). We add currency fixed effects to focus on the within variation at the exchange rate level.

Many of the variables are significant individually. Only three variables remain significant in the multivariate regression: the HKM intermediary leverage ratio, the U.S. dollar index, and LIBOR-OIS spread. We report our final specification of the multivariate regression where we exclude the insignificant variables. While many of the uncertainty measures are individually significant in univariate regressions, their significance is soaked up by the measure of intermediary leverage and the strength of the U.S. dollar. We check if they indeed capture the risk premiums by implementing cross-sectional regressions of expected changes in the xccy pricing errors on the beta exposures to the candidate risk factors. Specifically, we first estimate for each currency the exposure (i.e., betas) of changes in xccy pricing errors to changes in the intermediary leverage ratio and the U.S. dollar index. We then relate the average xccy pricing errors to these betas and estimate, with some abuse of language, the "price of risk" associated with the pricing errors.

We plot the average realized changes in pricing errors against their predicted counterparts based on factor exposures and risk prices in Figure 8. The predicted pricing errors line up with xccy "returns" quite well. The estimated risk prices are statistically significant. Our cross-section is small, so the evidence is merely suggestive. The evidence is consistent with the view that our model omits important sources of risk premiums in these markets.

Haddad and Muir (2018) caution that a cross-sectional relation between excess returns and factor exposures to intermediary leverage may simply reflect high excess returns in times when dealers also happen to be constrained. They suggest to overcome this interpretation by focusing on a cross-section of asset classes. Evidence in favor of intermediary-based asset pricing is tied to a positive cross-sectional relation between the cost of intermediation for a given asset class and its exposure to intermediary risk aversion. Thus, in a last step, we conduct similar cross-sectional tests for the pricing errors of both 5-year xccy swap rates and 6-month forward premiums.

We regress changes in pricing errors on the Haddad and Muir (2018) 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 9 convey two messages. First, for all currencies, FX swaps are on average more intermediated through dealers than FX forwards. Second, there appears to be a positive link between the amount of dealer activity and exposure to intermediary risk aversion for xccy swaps, while that relation is much noisier for forward premiums. Remember that our fit of forward premiums was tight, and that we matched the broad pattern of xccy swap rates. However, our model is not able to eliminate all pricing errors for xccy swaps. Overall, this evidence further supports that we omit intermediation risk premiums in our framework.

4.4 Variance decomposition

We use the results from Table 5 to better understand the relative contribution of our model and intermediary risks to the variation in xccy rates. The pricing errors $\Delta\widehat{xccy}$ are defined as the difference between the observed ($xccy^d$) and model-implied ($xccy^m$)

xccy rates. According to column (12) in Table 5, we have that changes in the xccy pricing errors are approximately a linear combination of changes in the intermediary capital ratio (ICR) and changes in the level of the USD factor (USD), with loadings of 180.49 and -0.44, respectively.

We use these loadings to decompose the variance of xccy rates into contributions from $xccy^m$, ICR, and USD. We follow Campbell and Ammer (1993), and split pairwise covariances between the respective factors equally. We report in the first row of Table 6 the fraction of the variance of the observed xccy rate levels explained by the total variance derived from all components. On average the three factors capture about 95.21% of the observed xccy rate variance, with some variation across currencies. The last three rows in Table 5 provide a decomposition of the total variance into the respective components. The contribution to the total variance from the model-implied xccy rate is 68.33%, on average, and represents, therefore, the bulk of the variation. The contribution of the ICR and USD represent, on average, 11.40% and 20.28% of the total variance.

Our decomposition is consistent with our earlier evidence that the model captures the magnitudes of xccy rates reasonably well (about two thirds, to be specific), but that there are departures from the data which can be connected to factors associated with financial intermediary health (Adrian, Etula, and Muir, 2014; He, Kelly, and Manela, 2017; Haddad and Muir, 2018) and balance sheet costs (Du, Tepper, and Verdelhan, 2018; Andersen, Duffie, and Song, 2019; Fleckenstein and Longstaff, 2018).

5 Conclusion

In the era following the global financial crisis, prices in fixed income and exchange rate markets have exhibited patterns that are unusual from the perspective of classical textbook theories, and are, therefore, considered to be anomalies. Covered interest parity has been violated, cross-currency basis swaps have traded at non-zero prices, and swap spreads have turned negative. We examine the dynamics of all three asset classes across G11 currencies in a unifying way using a no-arbitrage framework.

First, we assume that true risk-free rates are unobserved and latent. Second, we assume the existence of a latent pricing kernel, implying that traded prices are consistent with no-arbitrage. Third, we assume that OTC derivatives transactions are fully collateralized and that collateral is costly.

Under these assumptions, we proceed and back out the true unobserved discount rates from plain vanilla interest rate swap contracts. We show that the collateral-adjusted implied discount rates consistently price forward and exchange rates and cross-currency basis swaps across all ten currencies. Thus, true discount rates consistent with no-arbitrage can jointly reconcile CIP deviations, non-zero cross-currency basis swap prices, and negative swap spreads.

There remain non-trivial pricing errors in xccy rates. We relate the cross-section of these pricing errors to measures of intermediaries' leverage and their willingness to provide USD funding. The evidence suggests that the law of one price still holds to a first order, but that there are departures due to the constraints faced by market participants.

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

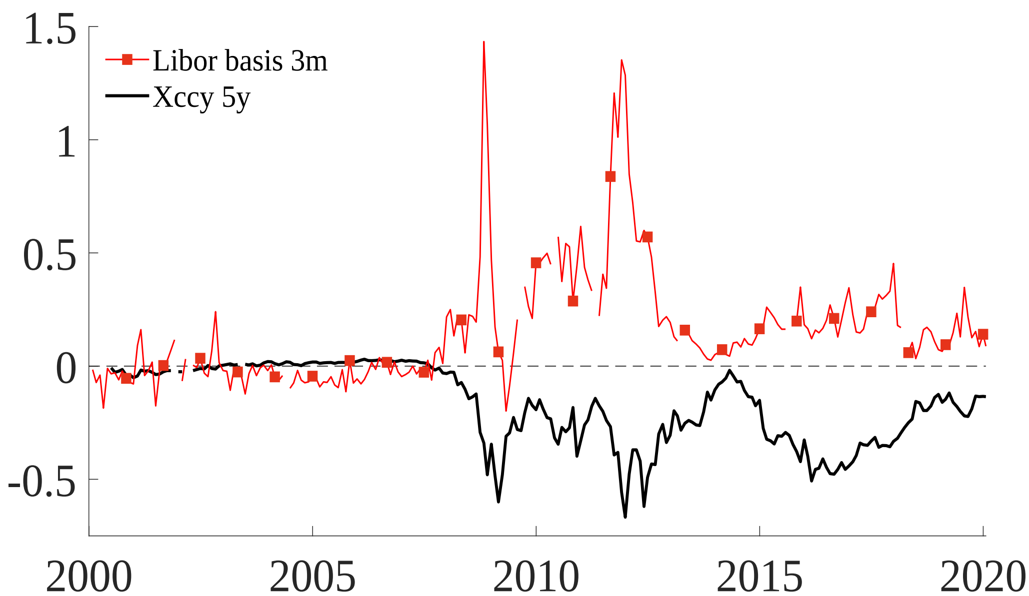
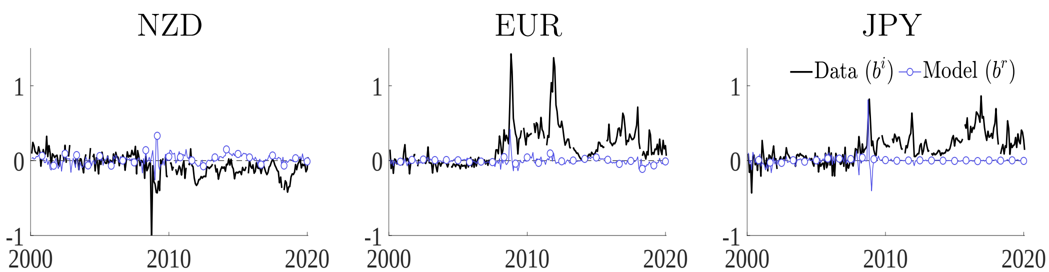
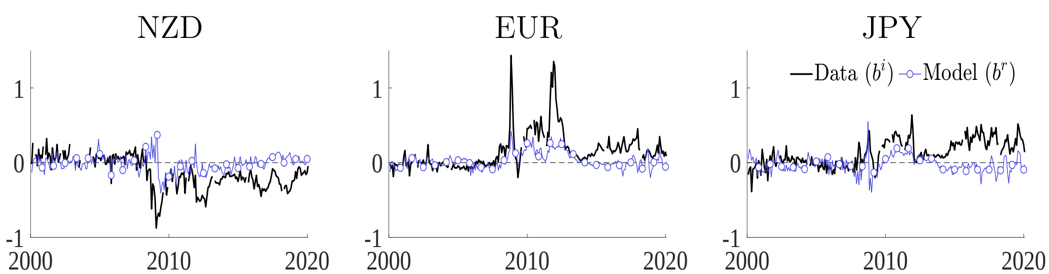


Figure 2: **Time-series of forward basis and xccy for NZD, EUR, and JPY.** In these figures, we report the time series of the forward basis (3 and 6 months, based on LIBOR in the data and on risk-free rate in the model) or xccy rates (5 and 20 years) implied from the model and compare it with the data.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

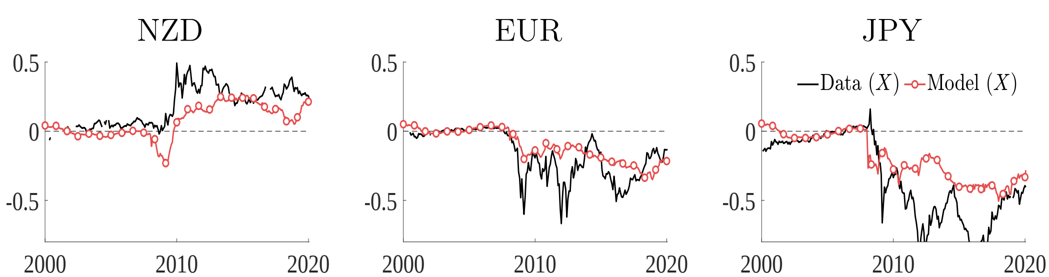
(A) Forward basis 3-month maturity



(B) Forward basis 6-month maturity



(C) xccy 5-year maturity



(D) xccy 20-year maturity

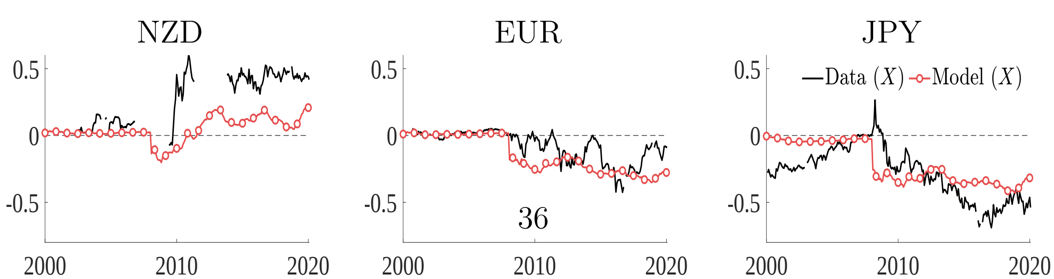


Figure 3: **Forward basis and xccy rate.** We report the mean of the forward basis (based on LIBOR in the data and on risk-free rate in the model) and the cross-currency basis swap rate (in bps). We also report the cross-sectional average of absolute rates, AVG. All exchange rates are expressed as the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

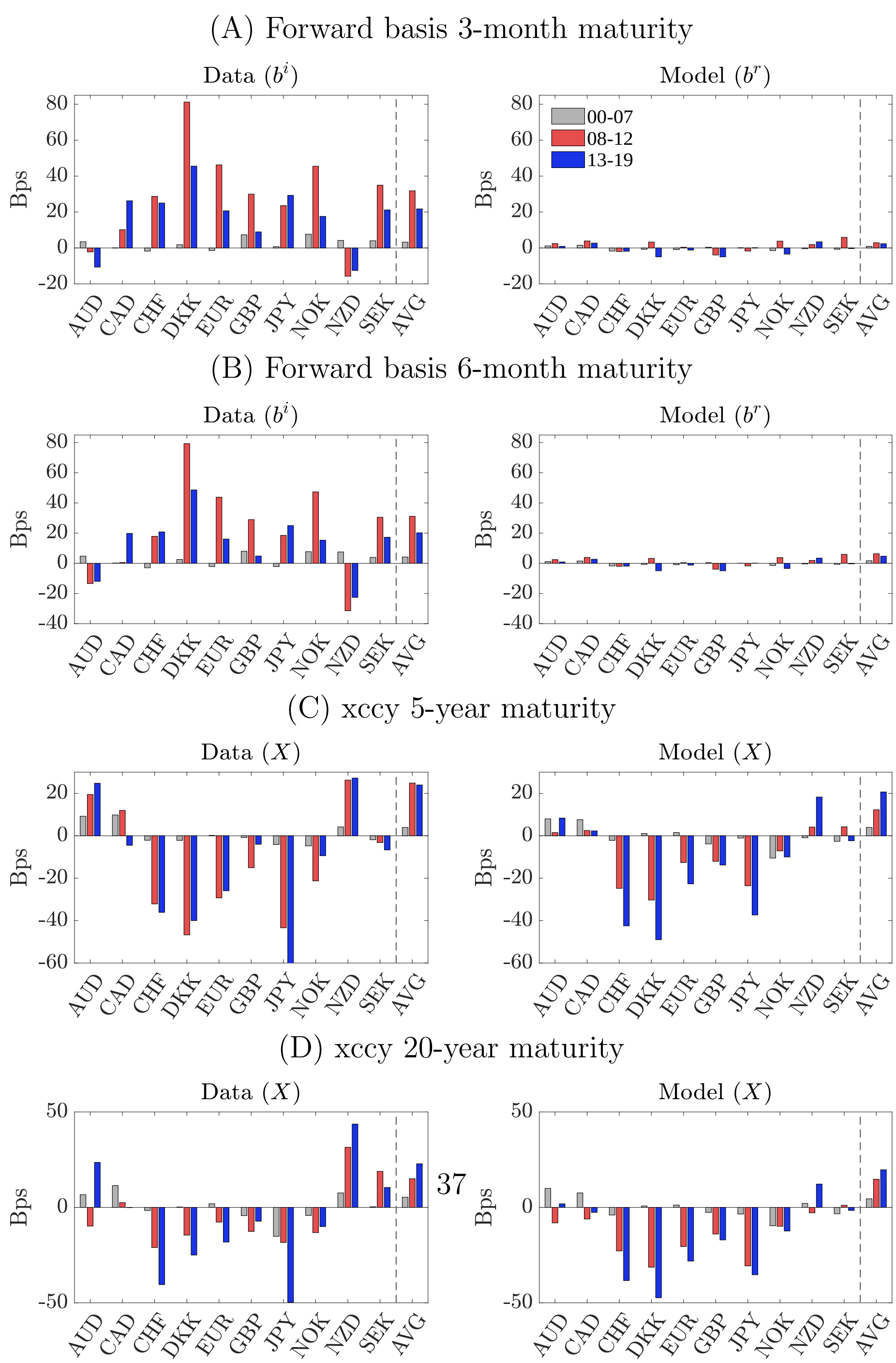


Figure 4: **Swap spread 20-year maturity.** In these figures, we report the time series of the swap spread from the model and compare it with the data. The sample is for selected G11 currencies:Canadian dollar (CAD), Swiss franc (CHF), Euro (EUR), British pound (GBP), Japanese yen (JPY), abd U.S. dollar (USD). Because of availability of Treasury data, Norwegian krone (NOK), and New Zealand dollar (NZD) have no swap spread data. For the same reason, Australian dollar (AUD) and Swedish krona (SEK) have very limited data. Danish krone (DKK) is very similar to EUR and, thus, not displayed. The sample period is January 2000 to December 2019. Source: Bloomberg.

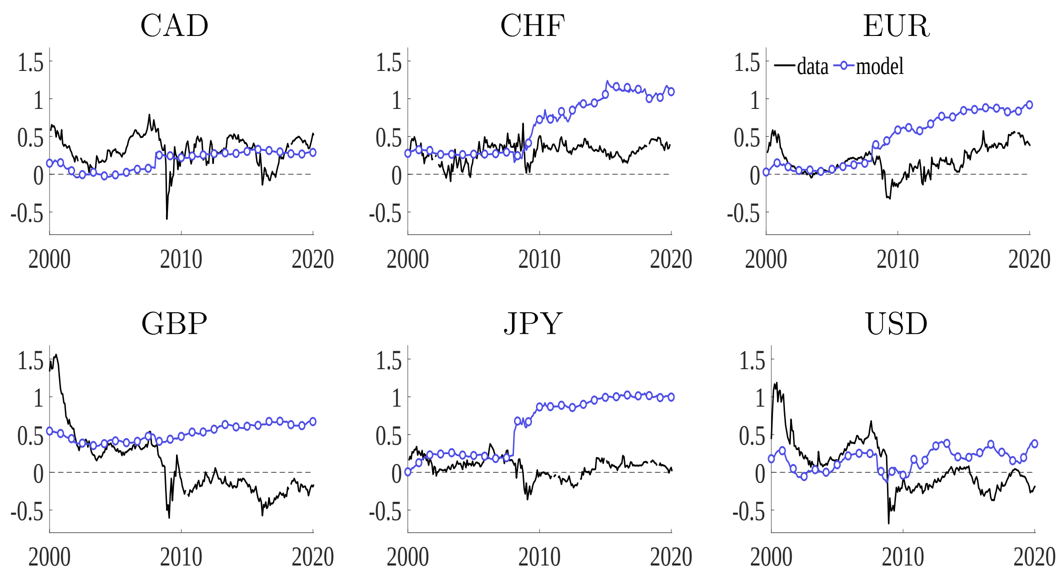
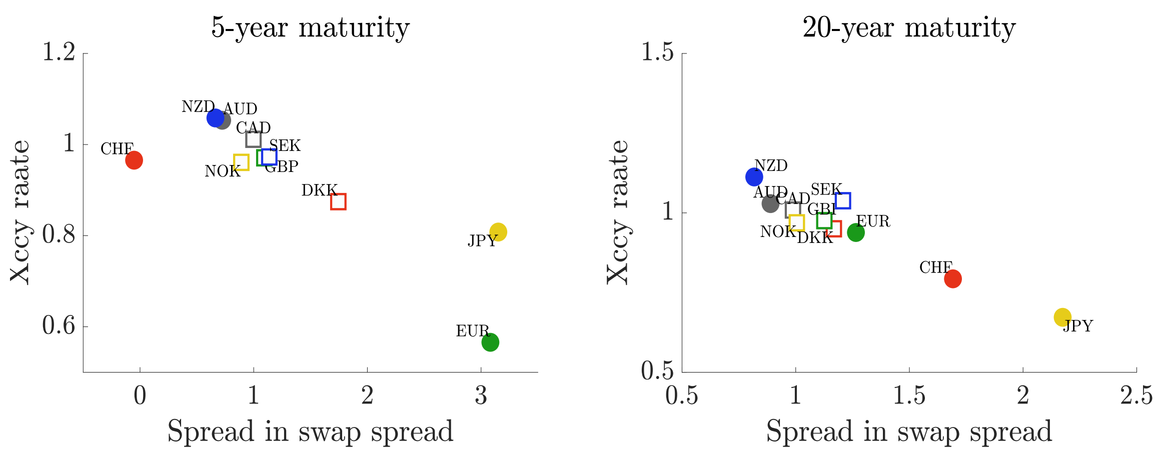


Figure 5: **Cross-sectional variation in Xccy.** In these figures, we report the scatter plot corresponding to Equation (7). For each country, we remove the top and bottom 1% for both the left-side and the right-side of the Equation and take their respective time-series averages. "xccy rate" refers to $1 + X_{0,T}/\widehat{CMS}_{0,T}$. In Panel (A), "spread in swap spread" is computed by $CMS_{0,T}/CMT_{0,T} \cdot \left(\widehat{CMS}_{0,T}/\widehat{CMT}_{0,T}\right)^{-1}$ whereas it is $CMS_{0,T}/CMT'_{0,T} \cdot \left(\widehat{CMS}_{0,T}/\widehat{CMT}'_{0,T}\right)^{-1}$ in Panel (B). The sample is for the G11 currencies: 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.

(A) Data



(B) Model

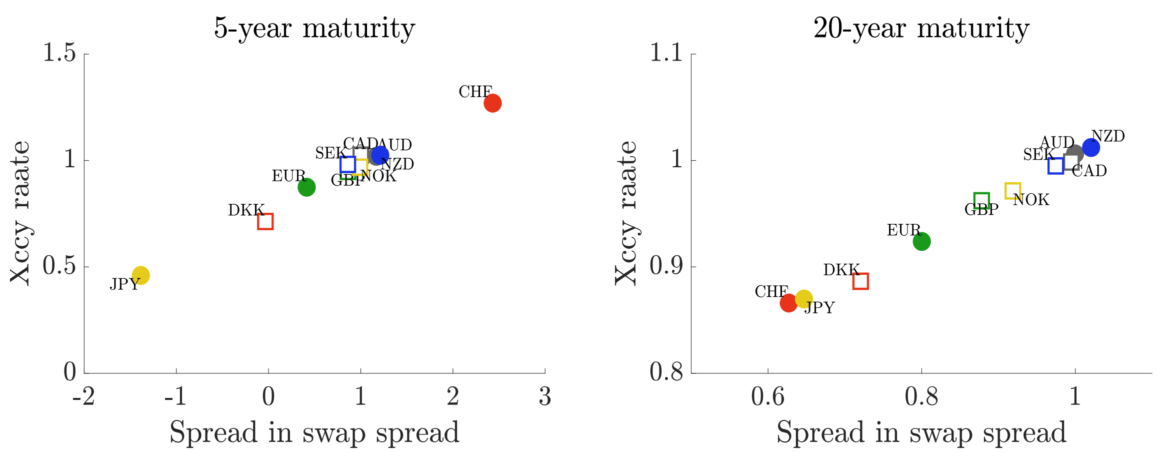


Figure 6: **Comparison of 1Y Interest Rate Proxies.** Each figure compares the model-implied 1-year interest rate to the predicted one and given by $\Delta r = -0.01 + 0.17 \cdot \Delta\text{LIBOR} + 0.40 \cdot \Delta(\text{Treasury} + \lambda) - 0.24 \cdot \Delta\text{CDS}$, where LIBOR corresponds to the country-specific Libor/interbank rate, Treasury corresponds to the country-specific Treasury rate, and λ refers to the country-specific convenience yield, computed as the Treasury basis plus the U.S. Refcorp-Treasury spread, and CDS corresponds to the country-specific 1-year local currency denominated CDS premium (we use the USD denomination if the local currency CDS is not available). We use G11 currencies, i.e., USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. We use Libor rates for USD, JPY, GBP, CHF, Cdor rates for Canada, Euribor rates for EUR, BBSW rates for AUD, BKBM rates for NZD, Stibor rates for SEK, Cibor rates for DKK, Nibor rates for NOK. The sample period is January 2000 to December 2019. Source: Bloomberg.

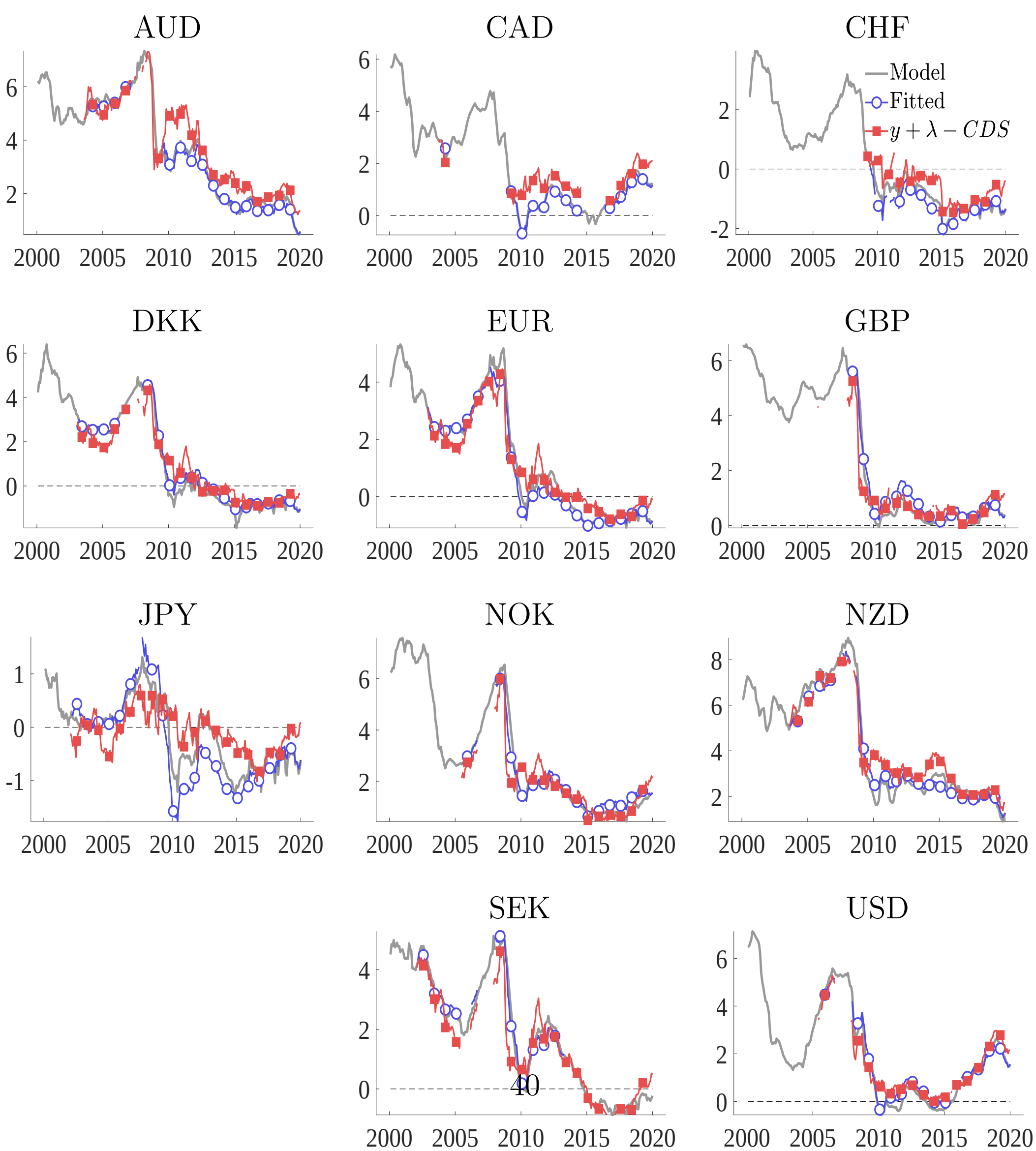


Figure 7: **Comparison of 3M Interest Rate Proxies.** Each figure compares the model-implied 3-month interest rate to the predicted one and given by $\Delta r = -0.01 + 0.31 \cdot \Delta\text{LIBOR} + 0.38 \cdot \Delta(\text{Treasury} + \lambda)$, where LIBOR corresponds to the country-specific Libor/interbank rate, Treasury corresponds to the country-specific Treasury rate, and λ refers to the country-specific convenience yield, computed as the Treasury basis plus the U.S. Refcorp-Treasury spread. We use G11 currencies, i.e., USD, JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, and NOK. We use Libor rates for USD, JPY, GBP, CHF, Cdor rates for Canada, Euribor rates for EUR, BBSW rates for AUD, BKBM rates for NZD, Stibor rates for SEK, Cibor rates for DKK, Nibor rates for NOK. The sample period is January 2000 to December 2019. Source: Bloomberg.

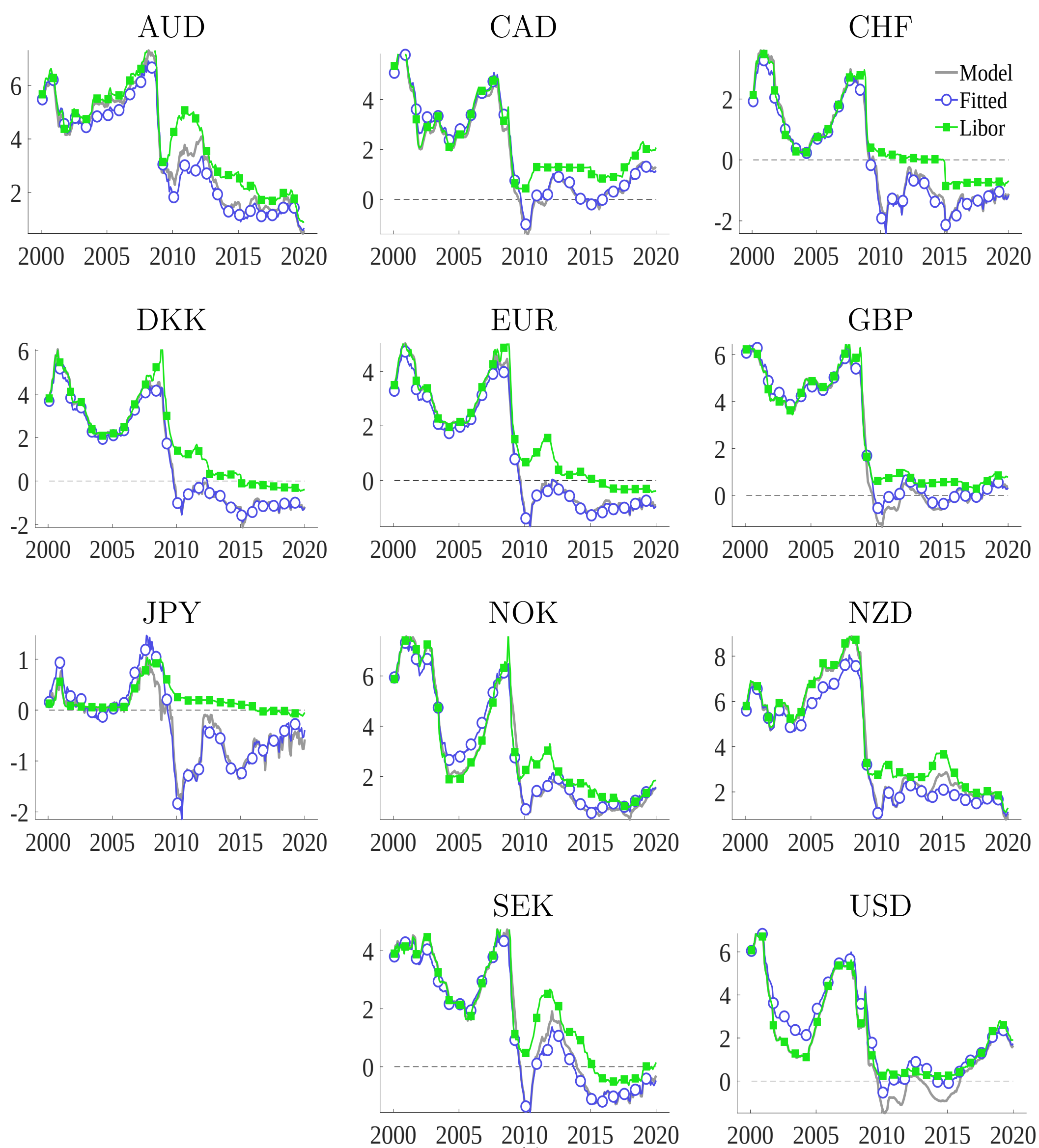
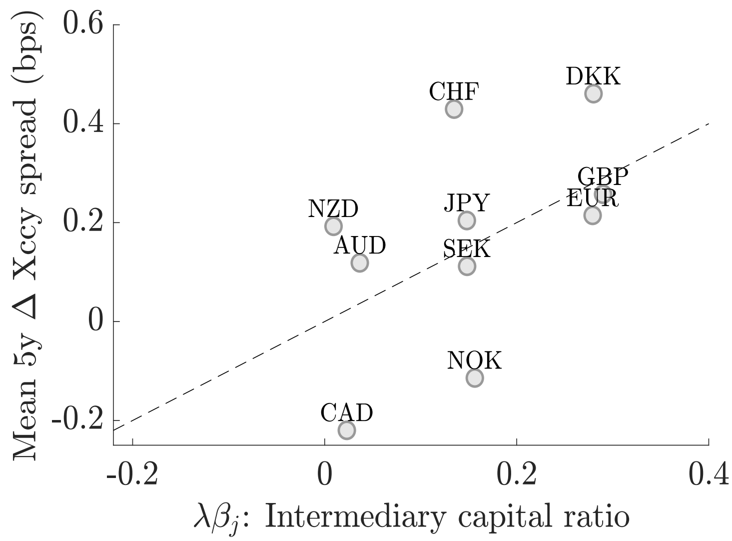


Figure 8: **Factor Exposure of Xccy Basis Swap Spread Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied 5-year xccy basis swap rate on a risk factor, i.e., $\Delta xccy\_spread_t = \alpha + \beta \cdot RF_t + \varepsilon_t$. We then project the average level of the xccy spread on the estimated betas $\widehat{\beta}$. We use two risk factors: changes in the He, Kelly, and Manela (2017) intermediary capital ratio (Δ HKM-ICR); changes in the trade-weighted U.S. dollar index (Δ USD Factor). The sample period is January 2000 to December 2019.



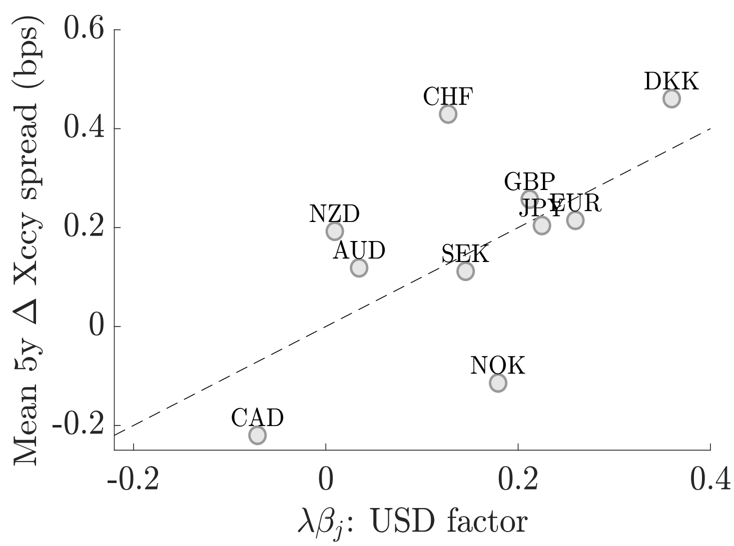


Figure 9: **Factor Exposure of Xccy Basis Swap Spread and Forward Premium Deviations.** For each of JPY, GBP, CAD, EUR, AUD, CHF, NZD, SEK, DKK, we regress changes in the spread between the observed and model-implied (i) 5-year xccy basis swap rate and (ii) the 6-month forward premium on the Haddad and Muir (2018) intermediary risk aversion factor, i.e., $\Delta \text{xccy spread}_{t+1} = \alpha + \beta RF_t + \varepsilon_t$. 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.

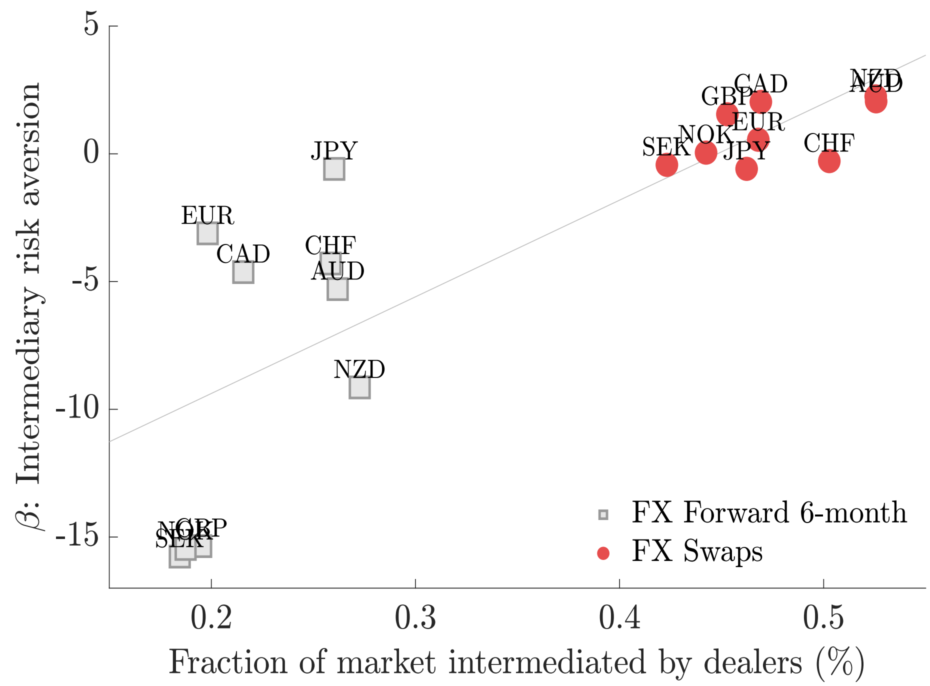


Table 1: Cash flows from a plain vanilla and fixed-for-fixed cross-currency basis swap

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

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

Table 2: Factor Structure in International Anomalies - By Currency

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

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

Table 3: Model-implied Interest Rates and Candidate Proxies

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

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

Table 4: Xccy Rates and Spreads between LIBOR and Model-Implied Interest Rates

In this table, we report results from the panel regressions where we project the absolute values of the observed 5-year xccy rates ($|xccy5y^D|$) on the spread between LIBOR and model-implied interest rates at the 3-month maturity (3M(i-3)). 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.

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

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

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

Table 6: Model-implied Variance Decomposition

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

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