The Term Structure of Currency Carry Trade Risk Premia

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Abstract

Fixing the investment horizon, the returns to currency carry trades decrease as the maturity of the foreign bonds increases. Across developed countries, the local currency term premia, which increase with the maturity of the bonds, offset the currency risk premia. Similarly, in the time-series, the predictability of foreign bond returns in dollars declines with the bonds' maturities. Leading no-arbitrage models in international finance do not match the downward term structure of currency carry trade risk premia. We derive a simple preference-free condition that no-arbitrage models need to reproduce in the absence of carry trade risk premia on long-term bonds.

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Carry trades correspond to simple investment strategies that are funded by borrowing in low interest rate currencies and invest in high interest rate currencies. What are their expected returns? According to the uncovered interest rate parity (U.I.P.) condition, which assumes that investors are risk-neutral, expected carry trade returns should be zero. Yet, empirically, borrowing in low interest rate currencies and investing in Treasury bills of high interest rate currencies delivers large excess returns. This is the U.I.P puzzle, and it gave rise to a large literature that studies the role of systematic risk in exchange rates and expectational errors. Our paper revisits the empirical evidence on carry trades and deepens the puzzle.

Our paper explores the properties of the same carry trade investment strategy implemented with long-term bonds, and compares it to the standard strategy that uses Treasury bills. We focus on the same set of G10 countries and consider the strongest predictors of bond and currency returns: the level and slope of the yield curve. The first strategy we consider goes long the bonds of high interest rate currencies and short the bonds of low interest rate currencies, whereas the second strategy goes long the bonds of flat yield curve currencies and short the bonds of steep yield curve currencies. Most importantly, the investment horizon is one month, as in the classic tests of the U.I.P. condition, not ten years, as in the tests of the U.I.P condition over long horizons. We find that, as the maturity of the bonds increases, the average excess return declines to zero. In other words, whereas the carry trade implemented with Treasury bills is profitable, the carry trade implemented with long-term bonds is not. Similar results emerge in individual country time-series predictability tests: as the maturity of the bonds increases, the predictability of the cross-country differences in dollar bond returns disappears.

The downward-sloping term structure of carry trade risk premia that we uncover represents a challenge for the leading models in international finance. To illustrate this point, we simulate the multi-country model of Lustig, Roussanov, and Verdelhan (2011). This reduced-form model, derived from the term structure literature, offers a flexible and transparent account of the U.I.P puzzle. Yet, we show that it implies a counterfactual flat term structure of currency carry trade risk premia. Our paper explains why other recent no-arbitrage models that replicate the U.I.P. puzzle fail to replicate the absence of carry trade risk premia at the long end of the yield curve.

To guide future work in international finance, we derive a simple preference-free characterization of carry trade risk premia on infinite-maturity bonds when financial markets are complete. We rely on Alvarez and

Jermann (2005)'s decomposition of the stand-investor's marginal utility or pricing kernel into a permanent and a transitory component. To be precise, we show that the difference between domestic and foreign long-term bond risk premia, expressed in domestic currency, is determined by the difference in the volatilities of the permanent components of the stochastic discount factors (SDFs, e.g., the growth rate of the stand-in investor's marginal utility). Our preference-free result is the bond equivalent of the usual expression for the carry trade risk premium in no-arbitrage models: when borrowing and investing in Treasury bills, the carry trade risk premium is equal to the differences in volatilities of the SDFs (see Bekaert, 1996, Bansal, 1997, and Backus, Foresi, and Telmer, 2001). This condition is the basis of most explanations of the U.I.P. puzzles. Our novel characterization similarly imposes additional restrictions on foreign and domestic SDFs.

To link our theoretical results to our empirical findings, two assumptions are required: (i) long-term bond returns are good proxy for infinite-maturity bond returns, and (ii) the level and the slope of the yield curve summarize all the relevant predictors of carry trade excess returns. Under these two assumptions, the significant downward-sloping term structure of carry trade risk premia implies that differences in SDFs' volatilities must be significantly larger than differences in their permanent components' volatilities. To obtain stricter implications, let us consider a benchmark case: the absence of carry trade risk premia on long-term bonds. In this case, the volatilities of the permanent components need to be equalized across currencies. This is a natural benchmark that is not rejected by the data, unlike its short-term bond equivalent. But it is only a benchmark: the large standard errors around average excess returns on long-term bonds in the data still leave room for cross-country differences in the volatilities of the SDF's permanent components. Those differences, however, are clearly bounded.

Armed with our preference-free results, we revisit a large class of dynamic asset pricing models that have been used to study U.I.P. violations, ranging from the reduced-form term structure model of Vasicek (1977) and Cox, Ingersoll, and Ross (1985) to their more recent multi-factor versions, to the Campbell and Cochrane (1999) external habit model, the Bansal and Yaron (2004) long run risks model, the disaster risk model, and the reduced-form model of Lustig, Roussanov, and Verdelhan (2011). We focus on models of the real SDF, given that there is no evidence that inflation risk can account for U.I.P. deviations (if anything, U.I.P. works better in high inflation environments, see Bansal and Dahlquist, 2000) or for cross-country variation in local currency term premia. None of the models we consider can replicate our empirical findings in their standard calibrations. But, when feasible, we show how to modify and calibrate these models to match the absence of carry trade risk premia on long-term bonds.

Our results are related to, but different from, the long-run U.I.P condition. The long-run U.I.P condition compares foreign and domestic long-term interest rates to long-term changes in exchange rates. Meredith and Chinn (2005) find that long-run U.I.P is a potentially valid description of the data. However, empirical tests lack power in finite samples: intuitively, there are few non-overlapping observations of 10-year windows available so far. From no-arbitrage conditions, we show that long-run U.I.P. always holds for temporary shocks and thus long-run U.I.P deviations have to come from permanent shocks to exchange rates. For long-run U.I.P to hold at all times, exchange rates must not have any permanent shocks and thus be stationary in levels (up to a deterministic time trend). Yet, exchange rate stationarity in levels is sufficient but not necessary to satisfy our preference-free condition on the volatilities of the permanent SDF components. As a result, the carry trade risk premia on long-term bonds could be zero without implying that long-run U.I.P always holds. Under some additional regularity conditions, in that case, long-run U.I.P would hold on average in no-arbitrage models.

Recent work has documented a downward sloping term structure of risk premia in equity markets (van Binsbergen, Brandt, and Koijen, 2012), real estate markets (Giglio, Maggiori, and Stroebel, 2015), and volatility markets (Dew-Becker, Giglio, Le, and Rodriguez, 2017). Backus, Boyarchenko, and Chernov (2016) provide a general analysis of the term structure of asset returns. Our work confirms the same pattern in currency markets, and offers a preference-free interpretation. Our theoretical result, although straightforward, has not been derived or used in the literature. On the one hand, at the short end of the maturity curve, currency risk premia are high when there is less overall risk in foreign countries' pricing kernels than at home (Bekaert, 1996, Bansal, 1997, and Backus, Foresi, and Telmer, 2001). On the other hand, at the long end of the maturity curve, local bond term premia compensate investors mostly for the risk associated with transitory innovations to the pricing kernel (Bansal and Lehmann, 1997; Hansen and Scheinkman, 2009; Alvarez and Jermann, 2005; Hansen, 2012; Hansen, Heaton, and Li, 2008; Backus, Chernov, and Zin, 2014; Borovička, Hansen, and Scheinkman, 2016; Backus, Boyarchenko, and Chernov, 2016). In this paper, we combine those two insights to derive general theoretical results under the assumption of complete financial markets. Foreign bond returns allow us to compare the permanent components of the SDFs, which as Alvarez and Jermann (2005) show, are the main drivers of the SDFs. Our preference-free condition does not apply to exchange rate models that allow for market segmentation (see, e.g., Gabaix and Maggiori, 2015, and Bacchetta and van Wincoop, 2005, for leading examples). It remains to be determined whether these models can fit our facts, so we leave this as an open question for future research.

The rest of the paper is organized as follows. Section 1 focuses on the time-series and cross-section of foreign bond risk premia. Section 2 compares recent no-arbitrage models to the empirical term structure of currency carry trade risk premia. In Section 3, we derive the no-arbitrage, preference-free theoretical restriction imposed on bond returns and SDFs. Section 4 links long-term U.I.P. to the properties of exchange rates. Section 5 concludes. An Online Appendix contains supplementary material and all proofs not presented in the main body of the paper.

1 Foreign Bond Returns in the Time-Series and Cross-Section

We first describe the data and the notation, and then turn to our empirical results on the time-series and cross-sectional properties of foreign government bond returns.

1.1 Data

Our benchmark sample, to which we refer as the G-10 sample, consists of a small homogeneous panel of developed countries with liquid bond markets: Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, and the U.K. The domestic country is the United States. We calculate the returns of both coupon and zero-coupon bonds for these countries.

Our data on total return bond indices were obtained from Global Financial Data. The dataset includes a 10-year government bond total return index, as well as a Treasury bill total return index, in U.S. dollars and in local currency. The data are monthly, starting in 1/1951 and ending in 12/2015. We use the 10-year bond returns as a proxy for long maturity bond returns. While Global Financial Data offers, to the best of our knowledge, the longest time-series of government bond returns available, the series have two key limitations. First, they pertain to coupon bonds, while the theory presented in this paper pertains to zero-coupon bonds. Second, they only offer 10-year bond returns, not the entire term structure of bond returns. To address these issues, we also use zero-coupon bond prices. Our zero-coupon bond dataset covers the same benchmark sample of G10 countries, but from at most 1/1975 to 12/2015, with different countries entering the sample at different dates. The details are in the Data Appendix.

Finally, we collect data on inflation rates and sovereign credit ratings. Inflation rates are calculated using monthly Consumer Price Index (CPI) data from Global Financial Data, whereas sovereign credit ratings are from Standard & Poor's, available over the 7/1989 to 12/2015 period. To construct averages of credit ratings, we assign each rating to a number, with a smaller number corresponding to a higher rating. The details are in the Data Appendix.

1.2 Notation

We now introduce our notation for bond prices, exchange rates, and bond and currency returns. In all cases, foreign variables are denoted as the starred version of their U.S. counterpart.

Bonds $P_t^{(k)}$ denotes the price at date t of a zero-coupon bond of maturity k, while $y_t^{(k)}$ denotes its continuously compounded yield: $\log P_t^{(k)} = -ky_t^{(k)}$. The one-period holding return on the zero-coupon bond is $R_{t+1}^{(k)} = P_{t+1}^{(k-1)}/P_t^{(k)}$. The log excess return on the domestic zero-coupon bond, denoted $rx_{t+1}^{(k)}$, is equal to

$$rx_{t+1}^{(k)} = \log\left[R_{t+1}^{(k)}/R_t^f\right], \tag{1}$$

where the risk-free rate is $R_t^f = R_{t+1}^{(1)} = 1/P_t^{(1)}$. Finally, r_t^f denotes the log risk-free rate: $r_t^f = \log R_t^f = y_t^{(1)}$.

Exchange Rates The nominal spot exchange rate in foreign currency per U.S. dollar is denoted S_t. Thus, an increase in S_t implies an appreciation of the U.S. dollar relative to the foreign currency. The log currency excess return, given by

$$rx_{t+1}^{FX} = \log\left[\frac{S_t}{S_{t+1}}\frac{R_t^{f,*}}{R_t^f}\right] = r_t^{f,*} - r_t^f - \Delta s_{t+1}, \tag{2}$$

is the log excess return of a strategy in which the investor borrows at the domestic risk-free rate, R_t^f, invests at the foreign risk-free rate, $R_t^{f,*}$, and converts the proceeds into U.S. dollars at the end of the period.

Bond Risk Premia The log return on a foreign bond position (expressed in U.S. dollars) in excess of the domestic (i.e., U.S.) risk-free rate is denoted $rx_{t+1}^{(k),\$}$. It can be expressed as the sum of the bond log excess return in local currency plus the log excess return on a long position in foreign currency:

$$rx_{t+1}^{(k),\$} = \log \left[\frac{R_{t+1}^{(k),*}}{R_t^f} \frac{S_t}{S_{t+1}} \right] = \log \left[\frac{R_{t+1}^{(k),*}}{R_t^{f,*}} \frac{R_t^{f,*}}{R_t^f} \frac{S_t}{S_{t+1}} \right] = \log \left[\frac{R_{t+1}^{(k),*}}{R_t^{f,*}} \right] + \log \left[\frac{R_t^{f,*}}{R_t^f} \frac{S_t}{S_{t+1}} \right] = rx_{t+1}^{(k),*} + rx_{t+1}^{FX}. \quad (3)$$

Taking conditional expectations, the total term premium in dollars consists of a foreign bond risk premium, $E_t[rx_{t+1}^{(k),*}]$, plus a currency risk premium, $E[rx_{t+1}^{FX}] = r_t^{f,*} - r_t^f - E_t[\Delta s_{t+1}]$.

We are not the first to study the relation between domestic and foreign bond returns. Prior work, from Campbell and Shiller (1991) to Bekaert and Hodrick (2001) and Bekaert, Wei, and Xing (2007), show that investors earn higher returns on foreign bonds from a country in which the slope of the yield curve is currently higher than average for that country. Ang and Chen (2010) and Berge, Jordà, and Taylor (2011) show that yield curve variables can also be used to forecast currency excess returns. These authors, however, do not examine the returns on foreign bond portfolios expressed in domestic currency. The following papers consider foreign bond returns in U.S. dollars: Dahlquist and Hasseltoft (2013) study international bond risk premia in an affine asset pricing model and find evidence for local and global risk factors, while Jotikasthira, Le, and Lundblad (2015) study the co-movement of foreign bond yields through the lenses of an affine term structure model. Our paper revisits the empirical evidence on bond returns without committing to a specific term structure model.

1.3 Time-Series Predictability of Foreign Bond Returns

To study the properties of the cross-country differences in expected bond excess returns, we first run individual currency predictability regressions on variables that can be used to predict bond and currency returns. We focus on the level and the slope of the term structure, the two predictors that have been shown to forecast both bond and currency returns.

We regress the 10-year dollar bond log excess return differential ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$) on the short-term interest rate differential ($r_t^{f,*} - r_t^f$, Panel A of Table 1) and on the yield curve slope differential ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B of Table 1), focusing on the post-Bretton Woods sample period (1/1975 – 12/2015). Given that the 10-year dollar bond log excess return differential (left columns) can be decomposed into the sum of currency log excess returns (rx_{t+1}^{FX}) and local currency bond log excess return differentials ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$) as noted in Equation (3), we also regress each of those two components on the same predictors (middle and right columns, respectively). By construction, the sum of the slope coefficients in the middle and right columns equals the slope coefficients in the corresponding left column.

For each individual country regression, we report Newey and West (1987) standard errors, setting the value of the lag truncation parameter (kernel bandwidth) to $S = 29$, following the recommendation of Lazarus, Lewis, Stock, and Watson (2018) to use $S = 1.3T^{1/2}$, where T is the number of observations. Given the well-known potential issues with using asymptotic distributions for statistical inference in finite samples, we use the Kiefer and Vogelsang (2005) non-standard fixed-b distributions for inference. In particular, we report fixed-b p-values for t-statistics and Wald tests using the methodology discussed in Vogelsang (2012). The panel regressions include country fixed effects, and standard errors are calculated using the Driscoll and Kraay (1998) methodology, which corrects for heteroskedasticity, serial correlation, and cross-equation correlation. As discussed in Vogelsang (2012), the fixed-b distributions also apply to test statistics that use the Driscoll and Kraay (1998) standard errors, so we also report fixed-b p-values for our panel regression coefficients. Our results are not driven either by the choice of kernel bandwidth or by the use of fixed-b distributions for inference: Table A1 in the Online Appendix refers to the same regressions as Table 1, but reports Newey and West (1987) and Driscoll and Kraay (1998) standard errors calculated with a kernel bandwidth of $S = 6$.

When using interest rate differentials as predictors, there is no consistent evidence in support of predictability of 10-year bond return differentials in dollars: indeed, out of nine countries, there is evidence of return predictability only for Japan. The slopes and constants are insignificant for all the other countries in the sample. In a panel regression, the slope coefficient is small and not statistically different from zero. In addition, we cannot reject the null that all constants in the invidual country regressions are zero and, similarly, that all slope coefficients are zero (p-values of 0.82 and 0.19, respectively). This evidence supports the view that there is no difference between expected dollar returns on long bond returns in these countries.

To better understand the lack of predictability, we decompose the dollar excess return differential into currency log excess returns and local currency bond log excess return differentials. As seen in the table, currency log excess returns are strongly forecastable by interest rate differentials (Hansen and Hodrick, 1980; Fama, 1984): as documented in the existing literature, higher than usual interest rate differentials in a given country pair predict higher than usual currency log excess returns. In a joint test of all slope coefficients, we can reject the null that interest rates do not predict currency excess returns. But, while Treasury bill return differentials in U.S. dollars are forecastable, long-term bond return differentials in U.S. dollars are not. The deterioration of return predictability for long-maturity bonds, compared to Treasury bills, appears to be due to the offsetting effect of local currency bond returns: higher than usual interest rate differences in a given country predict lower local currency bond return differences. Again, we can reject the null that interest rate differences do not predict local currency bond return differences at the 1% confidence level. In the panel regression, the local bond return slope coefficient is -1.34, largely offsetting the 1.98 slope coefficient in the currency excess return regression. The net effect on dollar bond returns is only 65 basis points, the slope coefficient is not statistically significant, and the panel regression adjusted R^2 is -0.05%.[1] Thus, from the perspective of a U.S. investor, the time variation in the currency excess return is largely offset by the variation in the local term premium.

When using yield curve slope differentials as predictors, a similar finding emerges. On the one hand, currency log excess returns are forecastable by yield curve slope differentials: a steeper than usual slope in a given country predicts lower than usual currency log excess returns. On the other hand, a slope steeper than usual in a given country also predicts higher local currency bond excess returns. In the panel regression, the local bond excess return slope coefficient is 3.96, more than offsetting the -2.02 slope coefficient in the currency excess return regression. The local currency bond return predictability merely confirm the results for U.S. bond excess returns documented by Fama and Bliss (1987), Campbell and Shiller (1991), and Cochrane and Piazzesi (2005). The net effect on dollar bond excess return differences is 194 basis points in a surprising direction: a steeper slope seems to weakly forecast *higher* dollar returns for foreign bonds, rather than lower dollar returns as for foreign Treasury Bills. The slope coefficient in the panel regression is statistically significant, albeit marginally (the p-value is 0.06). From the perspective of a U.S. investor, the time variation in the currency excess return is more than offset by the variation in the local term premium. This reverses the usual carry trade logic: investors want to short the currencies with lower than average slopes to harvest the local bond term premium. We turn now to the economic significance of these results.

To do so, we explore the risk-return characteristics of a simple investment strategy that goes long the foreign

[1] Note that the local currency log bond excess return differential contains the interest rate differential with a negative sign: $rx_{t+1}^{(10),*} - rx_{t+1}^{(10)} = r_{t+1}^{(10),*} - r_{t+1}^{(10)} - (r_t^{f,*} - r_t^f)$, where $r_{t+1}^{(10),*}$ and $r_{t+1}^{(10)}$ are the foreign and domestic holding-period bond returns. Thus, the local currency log bond excess return differential is highly predictable by the interest rate differential (right column of Panel A in Table 1), simply because the interest rate spread in effect predicts itself, as it is a component (with a negative sign) of the dependent variable. When we regress the local currency log return differential (instead of the excess returns) on the interest rate differential, there is no evidence of predictability (see Online Appendix). Measurement error in the short rate could also give rise to a negative relation, even in the absence of true predictability. The slope of the yield curves predicts the local currency log return differential.

Table 1: Time-Series Predictability Regressions

	Bond dollar return diff. $rx^{(10),\$} - rx^{(10)}$							Currency excess return rx^{FX}							Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$							Slope Diff.	Obs.
	α	s.e.	p-val	β	s.e.	p-val	R^2(%)	α	s.e.	p-val	β	s.e.	p-val	R^2(%)	α	s.e.	p-val	β	s.e.	p-val	R^2(%)	p-val	
	Panel A: Short-Term Interest Rates																						
Australia	0.01	[0.02]	0.69	-0.15	[0.91]	0.87	-0.20	-0.02	[0.02]	0.29	1.29	[0.55]	0.03	0.56	0.03	[0.01]	0.04	-1.44	[0.52]	0.01	1.51	0.20	492
Canada	0.02	[0.02]	0.27	-1.10	[0.69]	0.13	0.11	-0.01	[0.02]	0.49	1.22	[0.58]	0.05	0.46	0.03	[0.01]	0.01	-2.32	[0.52]	0.00	3.64	0.02	492
Germany	0.01	[0.02]	0.60	1.52	[1.18]	0.22	0.37	0.02	[0.02]	0.37	2.49	[1.05]	0.03	1.71	-0.01	[0.01]	0.59	-0.97	[0.40]	0.03	0.48	0.55	492
Japan	0.06	[0.03]	0.11	2.37	[0.71]	0.00	1.13	0.07	[0.03]	0.04	3.11	[0.70]	0.00	3.48	-0.01	[0.02]	0.43	-0.74	[0.41]	0.09	0.13	0.47	492
New Zealand	-0.03	[0.05]	0.50	0.69	[1.06]	0.53	-0.03	-0.07	[0.03]	0.05	2.23	[0.44]	0.00	3.14	0.04	[0.03]	0.25	-1.54	[0.88]	0.10	1.62	0.20	492
Norway	-0.02	[0.02]	0.45	0.72	[0.57]	0.23	0.08	-0.02	[0.02]	0.22	1.74	[0.55]	0.00	2.26	0.01	[0.01]	0.60	-1.02	[0.34]	0.01	0.97	0.22	492
Sweden	-0.00	[0.02]	0.94	-0.64	[0.86]	0.47	-0.02	-0.02	[0.02]	0.49	0.89	[0.88]	0.34	0.25	0.01	[0.01]	0.26	-1.53	[0.52]	0.01	2.02	0.23	492
Switzerland	0.02	[0.02]	0.37	1.16	[0.90]	0.23	0.33	0.05	[0.02]	0.03	2.45	[0.79]	0.01	2.43	-0.03	[0.01]	0.02	-1.29	[0.44]	0.01	1.69	0.30	492
U.K.	-0.02	[0.03]	0.52	1.02	[1.03]	0.34	0.04	-0.05	[0.03]	0.13	2.69	[1.24]	0.04	2.44	0.03	[0.02]	0.08	-1.67	[0.49]	0.00	1.39	0.32	492
Panel	–	–	–	0.65	[0.50]	0.23	-0.05	–	–	–	1.98	[0.49]	0.00	1.82	–	–	–	-1.34	[0.33]	0.00	1.37	0.00	4428
Joint zero p-val			0.82			0.19				0.16			0.00				0.08			0.00		0.32	
	Panel B: Yield Curve Slopes																						
Australia	0.06	[0.03]	0.04	3.84	[1.69]	0.04	1.54	0.00	[0.02]	0.90	-1.00	[1.16]	0.41	-0.02	0.05	[0.02]	0.00	4.84	[0.96]	0.00	7.65	0.03	492
Canada	0.04	[0.02]	0.04	4.04	[1.23]	0.00	2.25	-0.00	[0.01]	0.98	-0.72	[0.79]	0.39	-0.07	0.04	[0.01]	0.00	4.76	[0.81]	0.00	9.09	0.00	492
Germany	0.00	[0.02]	0.93	0.50	[1.57]	0.76	-0.18	-0.01	[0.02]	0.78	-3.05	[1.37]	0.04	1.15	0.01	[0.01]	0.45	3.55	[0.82]	0.00	4.07	0.11	492
Japan	0.00	[0.02]	0.90	-0.32	[1.12]	0.78	-0.19	-0.01	[0.02]	0.62	-4.18	[0.94]	0.00	2.91	0.01	[0.01]	0.24	3.85	[0.81]	0.00	3.96	0.02	492
New Zealand	0.08	[0.05]	0.17	2.94	[2.35]	0.24	1.26	-0.01	[0.04]	0.71	-1.60	[1.28]	0.24	0.62	0.09	[0.04]	0.02	4.55	[1.41]	0.00	7.41	0.11	492
Norway	-0.00	[0.02]	0.88	0.59	[0.98]	0.56	-0.12	-0.01	[0.02]	0.52	-2.03	[0.97]	0.05	1.33	0.01	[0.01]	0.46	2.62	[0.52]	0.00	3.35	0.07	492
Sweden	0.02	[0.02]	0.51	3.12	[1.21]	0.02	2.12	-0.00	[0.02]	0.98	-0.13	[1.02]	0.90	-0.20	0.02	[0.01]	0.19	3.25	[0.82]	0.00	5.29	0.06	492
Switzerland	0.00	[0.03]	0.95	0.97	[1.05]	0.38	-0.06	-0.02	[0.03]	0.41	-3.59	[1.27]	0.01	1.97	0.02	[0.01]	0.09	4.55	[1.00]	0.00	8.82	0.01	492
U.K.	0.02	[0.02]	0.47	1.59	[1.28]	0.24	0.17	-0.02	[0.03]	0.48	-3.17	[1.62]	0.07	2.11	0.04	[0.01]	0.01	4.75	[0.85]	0.00	7.95	0.03	492
Panel	–	–	–	1.94	[0.96]	0.06	0.42	–	–	–	-2.02	[0.82]	0.02	0.83	–	–	–	3.96	[0.66]	0.00	6.08	0.00	4428
Joint zero p-val			0.48			0.08				1.00			0.01				0.01			0.00		0.00	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^{f}$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 29. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 29. All p-values are fixed-b p-values, calculated using the approximation of the corresponding fixed-b asymptotic distribution in Vogelsang (2012).

bond and shorts the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate (or the foreign yield curve slope is lower than the U.S. yield curve slope), and goes long the U.S. bond and shorts the foreign bond when the U.S. interest rate is higher than the foreign country's interest rate (or the U.S. yield curve slope is lower than the foreign yield curve slope). To assess the risk-return trade-off, investors commonly look at the corresponding Sharpe ratios of an investment strategy, defined as the expected return less the risk-free rate divided by its standard deviation. In the absence of arbitrage opportunities, there is a one-to-one mapping from the R^2s in predictability regressions to the unconditional Sharpe ratios on investment strategies that exploit the predictability (see Cochrane, 1999, on p.75-76). Table 2 shows this mapping. The top panel reports the results of the interest rate-based strategy, whereas the bottom panel focuses on the slope-based strategy. The very low R^2s reported in Table 1 lead to low returns and Sharpe ratios in Table 2.

In our sample, none of the individual country dollar returns on the interest rate level strategy are statistically significant. The equally-weighted dollar return on the interest rate strategy is 0.70% per annum, and this return is not statistically significant. The equally-weighted annualized Sharpe ratio is 0.11, not significantly different from zero. Furthermore, with the exception of Sweden, none of the individual country dollar bond returns on the slope strategy are statistically significant either, even though the currency excess returns and the local currency bond returns typically are. The equally-weighted return on the slope strategy is -1.03% per annum, and this dollar return is not statistically significant. The equally-weighted annualized Sharpe ratio is -0.13, not significantly different from zero. In short, there is no evidence of economically significant dollar return predictability. This is not to say that no significant return predictability exists in our sample. To the contrary, there are large currency excess returns (with a Sharpe ratio of 0.47 for an equally-weighted portfolio in the interest-rate strategy) — these are the well-known deviations from U.I.P., i.e. the carry trade returns — and there are large local currency long-term bond returns (with a Sharpe ratio of 0.51 for an equally-weighted portfolio in the interest-rate strategy) — these are related to the well-known deviations from the expectation hypothesis, i.e. the local term premia. The two corresponding Sharpe ratios are larger than the one on the U.S. aggregate stock market. However, the currency risk premia and the local term premia cancel out, so the Sharpe ratio on foreign bond returns in dollars is not significant. From the vantage point of a U.S. investor, there is no evidence of economically significant return predictability in long foreign bonds.

Table 2: Dynamic Long-Short Foreign and U.S. Bond Portfolios

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Short-Term Interest Rates														
Australia	1.28	[2.23]	14.27	0.09	[0.16]	3.55	[1.75]	11.40	0.31	[0.16]	-2.27	[1.34]	8.46	-0.27	[0.15]
Canada	-0.46	[1.42]	9.10	-0.05	[0.16]	0.86	[1.13]	6.97	0.12	[0.16]	-1.31	[0.85]	5.50	-0.24	[0.15]
Germany	2.19	[1.93]	12.45	0.18	[0.16]	3.86	[1.72]	11.13	0.35	[0.16]	-1.67	[1.16]	7.30	-0.23	[0.16]
Japan	0.93	[2.19]	14.31	0.07	[0.16]	1.54	[1.73]	11.31	0.14	[0.16]	-0.61	[1.43]	9.03	-0.07	[0.16]
New Zealand	0.65	[2.56]	16.91	0.04	[0.16]	3.84	[1.87]	12.25	0.31	[0.17]	-3.19	[1.76]	11.47	-0.28	[0.16]
Norway	0.69	[1.98]	13.00	0.05	[0.16]	3.17	[1.60]	10.61	0.30	[0.16]	-2.48	[1.42]	8.99	-0.28	[0.16]
Sweden	-0.40	[2.03]	12.85	-0.03	[0.15]	2.30	[1.73]	11.14	0.21	[0.16]	-2.71	[1.38]	8.68	-0.31	[0.16]
Switzerland	0.57	[2.05]	12.84	0.04	[0.16]	1.14	[1.97]	12.22	0.09	[0.15]	-0.57	[1.13]	7.62	-0.07	[0.16]
United Kingdom	0.89	[2.00]	12.76	0.07	[0.15]	3.09	[1.59]	10.26	0.30	[0.16]	-2.20	[1.25]	8.21	-0.27	[0.15]
Equally-weighted	0.70	[1.00]	6.36	0.11	[0.16]	2.59	[0.87]	5.55	0.47	[0.17]	-1.89	[0.60]	3.73	-0.51	[0.16]
	Panel B: Yield Curve Slopes														
Australia	-1.88	[2.22]	14.26	-0.13	[0.16]	2.89	[1.80]	11.41	0.25	[0.17]	-4.76	[1.34]	8.37	-0.57	[0.14]
Canada	-2.07	[1.41]	9.08	-0.23	[0.15]	1.23	[1.10]	6.97	0.18	[0.16]	-3.30	[0.83]	5.43	-0.61	[0.16]
Germany	1.98	[1.92]	12.46	0.16	[0.16]	5.17	[1.71]	11.09	0.47	[0.16]	-3.19	[1.14]	7.25	-0.44	[0.15]
Japan	-0.71	[2.21]	14.31	-0.05	[0.16]	4.60	[1.73]	11.24	0.41	[0.16]	-5.31	[1.42]	8.90	-0.60	[0.16]
New Zealand	-0.18	[2.57]	16.91	-0.01	[0.16]	3.49	[1.90]	12.26	0.28	[0.17]	-3.67	[1.77]	11.45	-0.32	[0.15]
Norway	-0.56	[2.05]	13.00	-0.04	[0.15]	2.84	[1.73]	10.61	0.27	[0.16]	-3.40	[1.39]	8.97	-0.38	[0.15]
Sweden	-3.62	[1.99]	12.81	-0.28	[0.16]	1.32	[1.75]	11.15	0.12	[0.17]	-4.94	[1.38]	8.60	-0.57	[0.16]
Switzerland	0.47	[2.00]	12.84	0.04	[0.15]	4.80	[1.91]	12.15	0.40	[0.15]	-4.33	[1.17]	7.51	-0.58	[0.16]
United Kingdom	-2.73	[1.95]	12.73	-0.21	[0.16]	2.06	[1.61]	10.29	0.20	[0.16]	-4.79	[1.30]	8.12	-0.59	[0.16]
Equally-weighted	-1.03	[1.24]	7.82	-0.13	[0.16]	3.16	[1.08]	6.68	0.47	[0.16]	-4.19	[0.80]	5.04	-0.83	[0.16]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate (or the foreign yield curve slope is lower than the U.S. yield curve slope), and go long the U.S. bond and short the foreign country bond when the U.S. interest rate is higher than the country's interest rate (or the U.S. yield curve slope is lower than the foreign yield curve slope). Results based on interest rate levels are reported in Panel A and results based on interest rate slopes are reported in Panel B. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). The holding period is one month. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 1/1975–12/2015.

1.4 Cross-Sectional Properties of Foreign Bond Returns

After focusing on time-series predictability, we turn now to cross-sectional evidence, as in Lustig and Verdelhan (2007). There is no mechanical link between the time-series and cross-sectional evidence. The time-series regressions test whether a predictor that is higher than its average implies higher returns, while the cross-sectional tests show whether a predictor that is higher in one country than in others implies higher returns in that country (see Hassan and Mano, 2014). Our cross-sectional evidence echoes some papers that study the cross-section of bond returns: Koijen, Moskowitz, Pedersen, and Vrugt (2016) and Wu (2012) examine the currency-hedged returns on 'carry' portfolios of international bonds, sorted by a proxy for the carry on long-term bonds. But these papers do not examine the interaction between currency and term risk premia, the topic of our paper.

We sort countries into three portfolios on the level of the short-term interest rates or the slope of their yield curves. Portfolios are rebalanced every month and those formed at date t only use information available at that date. Portfolio-level log excess returns are obtained by averaging country-level log excess returns across all countries in the portfolio. We first describe results obtained with the 10-year bond indices, reported in Table 3, and then turn to the zero-coupon bonds to study the whole term structure, presented in Figure 2.

1.4.1 Sorting by Interest Rates

We start with the currency portfolios sorted by short-term interest rates. In order to focus on the conditional carry trade, we use interest rates in deviation from their past 10-year rolling mean as the sorting variable.[2] Thus, the first (third) portfolio includes the conditionally low (high) interest rate currencies. Clearly, the classic uncovered interest rate parity condition fails in the cross-section: the currencies in the third portfolio only depreciate by 60 basis points per year on average, not enough to offset the 2.65% interest rate difference and thus delivering a 2.05% return. As a result, average currency excess returns increase from low- to high-interest-rate portfolios, ranging from -0.61% to 2.05% per year over the last 40 years. Thus, the long-short currency carry trade (invest in Portfolio 3, short Portfolio 1) implemented with Treasury bills delivers an average annual log excess return of $2.05\% - (-0.61\%) = 2.66\%$ and a Sharpe ratio of 0.36 (Panel B), higher than the Sharpe ratio on the U.S. S&P500 equity index over the same period.

[2] The unconditional carry trade goes long in currencies with an average high interest rates (Hassan and Mano, 2014). That is not the focus of our paper.

However, implementing the carry trade with long-term bonds does not yield a similar performance, as local currency bond premia decrease from low- to high-interest-rate portfolios, from 3.53% to −0.25%, implying that the 10-year bond carry trade entails a local currency bond premium of $-0.25\% - 3.53\% = -3.78\%$ (Panel C). Therefore, the long-short currency carry trade implemented with long-term government bonds, which is the sum of the two long-short returns above, delivers a negative average return of $-3.78\% + 2.66\% = -1.12\%$ (Panel D) that is not statistically significant. This contrasts with the equivalent trade using Treasury bills. The dollar bond risk premia (−1.12%) are statistically different from the carry trade risk premia (2.66%) because the local term premia (−3.78%) are statistically significant. Investors have no reason to favor the long-term bonds of a particular set of countries on the basis of average returns after converting the returns into the same currency (here, U.S. dollars).

Inflation risk is not a good candidate explanation for local currency bond excess returns, because inflation is higher and more volatile in countries that are in the last portfolio than in the first portfolio (Panel A of Table 3, left section). Similarly, sovereign default risk is not a plausible explanation, given that the countries in the first portfolio tend to have slightly better credit ratings than the countries in the last portfolio. The relatively high term premium of 3.53% in the first portfolio thus corresponds to relatively low inflation and default risk.

1.4.2 Sorting by Slopes

Similar results emerge when we sort countries into portfolios by the slope of their yield curves. There is substantial turnover in these portfolios, more so than in the usual interest rate-sorted portfolios. On average, the flat slope currencies (first portfolio) tend to be high interest rate currencies, while the steep slope currencies (third portfolio) tend to be low interest rate currencies. As expected, average currency log excess returns decline from 2.41% per annum on the first portfolio (low slope, high short-term interest rates) to −1.24% per annum on the third portfolio (high slope, low short-term interest rates) over the last 40 years (Panel B). Therefore, a long-short position of investing in flat-yield-curve currencies (Portfolio 1) and shorting steep-yield-curve currencies (Portfolio 3) delivers a currency excess return of 3.65% per annum and a Sharpe ratio of 0.49. Our findings confirm those of Ang and Chen (2010): the slope of the yield curve predicts currency excess returns at the short end of the maturity spectrum.

However, those currency premia are offset by term premia, as local currency bond excess returns and

Table 3: Cross-sectional Predictability: Bond Portfolios

Portfolio		Sorted by Short-Term Interest Rates 1	2	3	$3-1$	Sorted by Yield Curve Slopes 1	2	3	$1-3$
		Panel A: Portfolio Characteristics							
Inflation rate	Mean	2.90	3.45	4.81	1.91	4.89	3.41	2.87	2.02
	s.e.	[0.16]	[0.19]	[0.23]	[0.20]	[0.23]	[0.19]	[0.18]	[0.19]
	Std	1.03	1.23	1.48	1.30	1.39	1.16	1.20	1.26
Rating	Mean	1.45	1.25	1.49	0.04	1.54	1.38	1.28	0.25
	s.e.	[0.02]	[0.02]	[0.02]	[0.04]	[0.02]	[0.02]	[0.02]	[0.03]
Rating (adj. for outlook)	Mean	1.50	1.37	1.84	0.33	1.84	1.50	1.37	0.47
	s.e.	[0.03]	[0.02]	[0.02]	[0.04]	[0.02]	[0.02]	[0.02]	[0.03]
$y_t^{(10),*} - r_t^{*,f}$	Mean	1.52	0.92	-0.44	-1.96	-0.81	0.85	1.96	-2.76
		Panel B: Currency Excess Returns							
$-\Delta s_{t+1}$	Mean	-0.44	0.11	-0.60	-0.16	-0.95	0.38	-0.36	-0.58
$r_t^{f,*} - r_t^f$	Mean	-0.17	0.54	2.65	2.81	3.35	0.55	-0.88	4.23
rx_{t+1}^{FX}	Mean	-0.61	0.66	2.05	2.66	2.41	0.92	-1.24	3.65
	s.e.	[1.35]	[1.44]	[1.36]	[1.14]	[1.48]	[1.38]	[1.40]	[1.18]
	SR	-0.07	0.07	0.23	0.36	0.26	0.11	-0.14	0.49
		Panel C: Local Currency Bond Excess Returns							
$rx_{t+1}^{(10),*}$	Mean	3.53	2.60	-0.25	-3.78	-1.01	2.29	4.61	-5.61
	s.e.	[0.69]	[0.69]	[0.73]	[0.77]	[0.76]	[0.69]	[0.70]	[0.74]
	SR	0.80	0.58	-0.05	-0.77	-0.21	0.53	1.00	-1.18
		Panel D: Dollar Bond Excess Returns							
$rx_{t+1}^{(10),\$}$	Mean	2.92	3.26	1.80	-1.12	1.40	3.21	3.36	-1.96
	s.e.	[1.56]	[1.58]	[1.57]	[1.33]	[1.64]	[1.57]	[1.62]	[1.38]
	SR	0.29	0.32	0.18	-0.13	0.14	0.33	0.32	-0.22
$rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$	Mean	0.14	0.48	-0.98	-1.12	-1.38	0.43	0.59	-1.96
	s.e.	[1.64]	[1.64]	[1.73]	[1.33]	[1.81]	[1.63]	[1.75]	[1.38]

Notes: The countries are sorted by the level of their short term interest rates in deviation from the 10-year mean into three portfolios (left section) or the slope of their yield curves (right section). The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The table reports the average inflation rate, the standard deviation of the inflation rate, the average credit rating, the average credit rating adjusted for outlook, the average slope of the yield curve ($y^{(10),*} - r^{*,f}$), the average change in exchange rates (Δs), the average interest rate difference ($r^{f,*} - r^f$), the average currency excess return (rx^{FX}), the average foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*}$) and in U.S. dollars ($rx^{(10),\$}$), as well as the difference between the average foreign bond excess return in U.S. dollars and the average U.S. bond excess return ($rx^{(10),\$} - rx^{(10)}$). For the excess returns, the table also reports their Sharpe ratios (denoted SR). The holding period is one month. The log returns are annualized. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The data are monthly and the sample is 1/1975–12/2015.

currency excess returns move in opposite directions across portfolios. In particular, the first portfolio produces negative bond average excess returns of −1.01% per year, compared to 4.61% on the third portfolio (Panel C), so the slope carry trade generates an average local currency bond excess return of −5.61% per year. This result is not mechanical: the spread in the slopes is about half of the spread in local currency excess returns. The corresponding average dollar bond excess returns range from 1.40% to 3.36%, so the slope carry trade implemented with 10-year bonds delivers an average annual dollar excess return of −1.96% (Panel D), which is not statistically significant.

Importantly, this strategy involves long positions in bonds issued by countries with slightly lower average inflation, less inflation risk and slightly better average sovereign credit rating, despite the higher term premium. Therefore, the offsetting effect of local currency term premia is unlikely to be due to inflation or credit risk.

1.4.3 Looking Across Subsamples

To summarize the portfolio results and test their robustness across subsamples, Figure 1 plots the cumulative returns on interest rate and slope carry strategies over the entire sample. The full line is the cumulative local currency bond excess return (in logs), while the dashed line is the cumulative currency excess return (in logs). The cumulative dollar excess return is the sum of these two (not shown). The top panels sort by interest rates on the left and interest rate deviations on the right. The bottom panel sort by slope and slope deviations. Overall, the same patterns reappear in each these plots, but there are three noticeable differences.

First, when we sort on interest rates, the off-setting effect of local currency bond excess returns is slightly weaker, and the currency excess returns are larger, than when we sort on interest rate deviations. The latter sort only captures the conditional carry premium, while the former captures the entire carry premium. The carry trade risk premium is somewhat larger when sorting on interest rates, as documented by Lustig, Roussanov, and Verdelhan (2011) and Hassan and Mano (2014), as these sorts capture both the conditional carry trade premium (long in currencies with currently high interest rates) and the unconditional carry trade premium (long in currencies with high average interest rates). Since our paper is about the conditional currency carry trade, we focus mainly on the demeaned sorts. Second, when we use interest rate (deviations) sorts, the negative contribution of local currency bond premia weakens starting in the mid 80s (90s), compared to the earlier part of the sample. This is not surprising given that interest rates started to converge across G10 countries in the second part of the sample (Wright, 2011). In other words, when G10 countries have similar low interest rates, these rates do not predict large differences in term premia across countries. Third, the offsetting effect of local currency bond premia is larger when we sort on the slope of the yield curve, and it does not weaken in the second part of the sample.

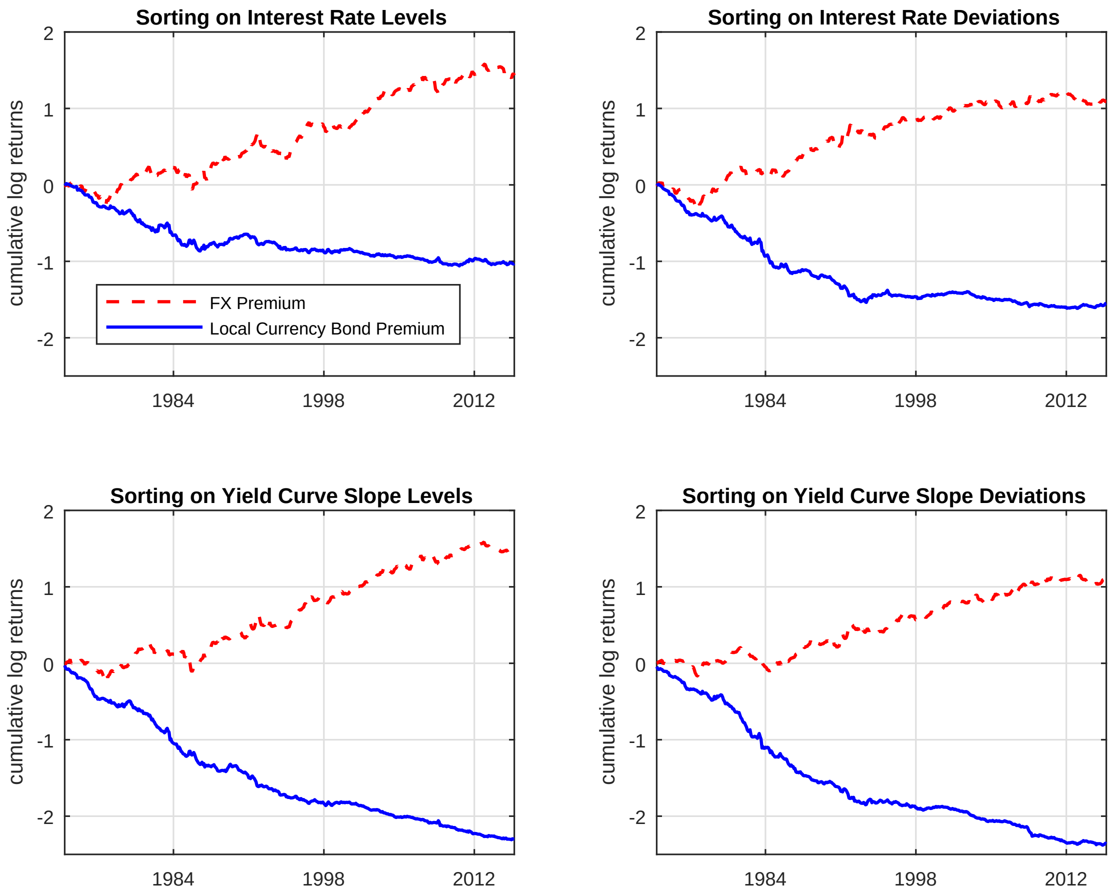

Figure 1: FX Premia and Local Currency Bond Premia— The figure shows the cumulative log excess returns on interest rate and slope carry investment strategies that go long in high interest rate (flat slope) currencies and short low interest rate (steep slope) currencies. The full line is the cumulative log currency excess return. The dashed line is the cumulative log local currency bond excess return. The cumulative log dollar excess return is represented by the sum of these 2 lines. At each date t, the countries are sorted by the interest rate (slope of the yield curves) into three portfolios. The slope of the yield curve is measured by the difference between the 10-year yield and the 3-month interest rate at date t. The holding period is one month. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The data are monthly and the sample is 1/1975–12/2015.

1.4.4 Looking Across Maturities

The results we discussed previously focus on the 10-year maturity. We now turn to the full maturity spectrum, using the zero-coupon bond dataset. The panel is unbalanced, and because of data limitations, we can only examine three-month holding period returns over the 4/1985–12/2015 sample, which mitigates predictability (4/1985 is the first month for which we have data for at least three foreign bonds). As a result, the standard errors are larger than in the benchmark sample. Building on the previous results, countries as sorted by the slope of their term structures because the slope appears as the strongest predictor of both currency and term premia over this sample. We sort only on the slope of the yield curve, not on the slope deviation from its 10-year mean, because of the shorter sample.

Our findings are presented graphically in Figure 2, which shows the dollar log excess returns as a function of the bond maturity, using the same set of funding and investment currencies. Investing in the short-maturity bills of countries with flat yield curves (mostly high short-term interest rate countries), while borrowing at the same horizon in countries with steep yield curves (mostly low short-term interest rate countries) leads to an average dollar excess return of 2.67% per year (with a standard error of 1.50%). This is the slope version of the standard currency carry trade. However, when we implement the same strategy using longer maturity bonds instead of short-term bills, the dollar excess return decreases monotonically as the maturity of the bonds increases. The zero-coupon findings confirm our previous results and seem to rule out measurement error in the 10-year coupon bond indices as an explanation. At the long end (15-year maturity), the bond term premium more than offsets the currency premium, so the slope carry trade yields a (non-significant) average annual dollar return of −2.18% (with a standard error of 2.28%). The average excess returns at the long-end of the yield curve are statistically different from those at the short-end: the difference between those returns correspond to the local term premia, which are equal to −4.85% with a standard error of 1.82%. Therefore, carry trade strategies that yield positive average excess returns when implemented with short-maturity bonds yield lower (or even negative) excess returns when implemented using long-maturity bonds.[3]

1.5 Robustness Checks

We consider many robustness checks, both regarding our time-series results and our cross-sectional results. The time-series predictability robustness checks are reported in Section A of the Online Appendix, whereas the

[3]When we use interest rate sorts, the term structure is flat: The carry premium is 3.71% per annum (with a standard error of 1.80%), while the local 10-year bond premium is only -0.21% per annum (with a standard error of 1.76%), so the dollar bond premium at the 15-year maturity is 3.50% (with a standard error of 2.32%). As noted in the previous subsection and apparent in the top panels of Figure 1, interest rates in levels do not predict bond excess returns in the cross-section over the 4/1985–12/2015 sample.

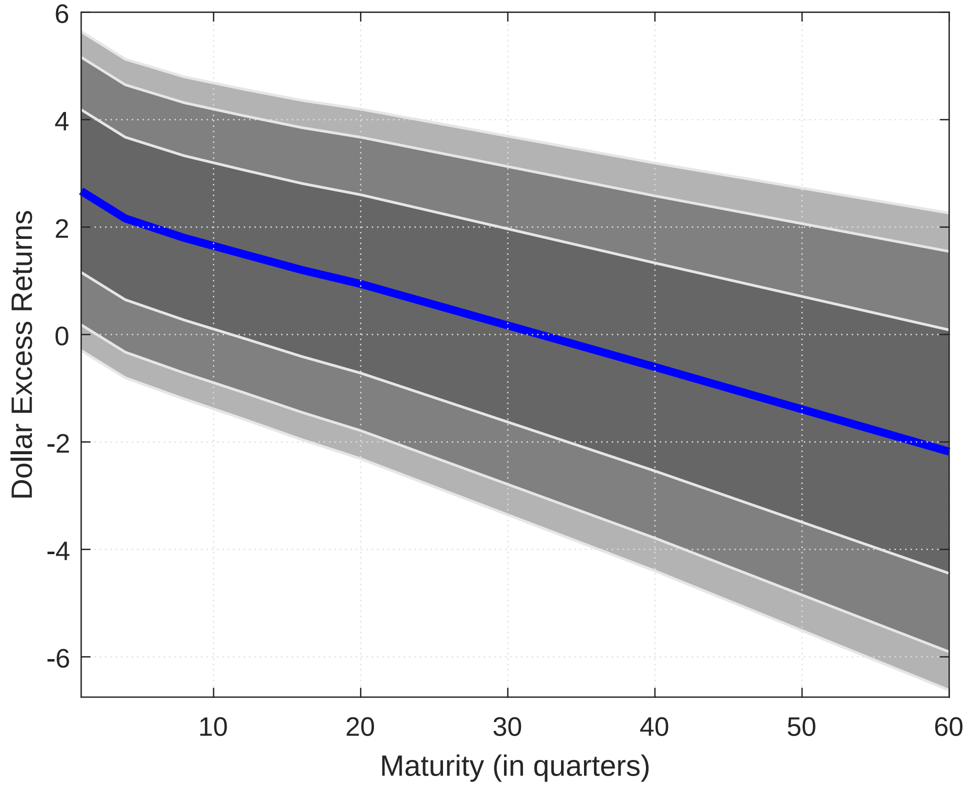

Figure 2: Long-Minus-Short Foreign Bond Risk Premia in U.S. Dollars— The figure shows the dollar log excess returns as a function of the bond maturities. Dollar excess returns correspond to the holding period returns expressed in U.S. dollars of investment strategies that go long and short foreign bonds of different countries. The unbalanced panel of countries consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. At each date t, the countries are sorted by the slope of their yield curves into three portfolios. The first portfolio contains countries with flat yield curves while the last portfolio contains countries with steep yield curves. The slope of the yield curve is measured by the difference between the 10-year yield and the three-month interest rate at date t. The holding period is one quarter. The returns are annualized. The dark shaded area corresponds to one-standard-error bands around the point estimates. The gray and light gray shaded areas correspond to the 90% and 95% confidence intervals. Standard deviations are obtained by bootstrapping 10,000 samples of non-overlapping returns. Zero-coupon data are monthly, and the sample window is 4/1985–12/2015.

cross-sectional portfolio robustness checks can be found in Section B of the Online Appendix.

As regards time-series predictability, we consider predictability regressions using inflation and sovereign credit as additional controls, we consider an alternative decomposition of dollar bond returns into an exchange rate component and a local currency bond return difference, we include predictability results with GBP as the base currency and we report predictability results using different time-windows (10/1983–12/2007, 1/1975–12/2007, 10/1983–12/2015) and investment horizons (three months). We find that our main results are robust to those alternative specifications.

As regards cross-sectional currency portfolios, we consider different lengths of the bond holding period (three and twelve months), different time windows, different samples of countries, sorts by (non-demeaned) interest rate levels, and other potential explanations of excess returns. Our results appear robust to the choice of the bond holding period and across time windows. Furthermore, our results appear robust across several samples of countries. Introducing more countries adds power to the experiment, but forces us to consider less liquid and more default-prone bond markets. In what may be of particular interest, we show that inflation risk or credit risk are unlikely explanations for differences in term premia even in larger sets of countries and different time windows. For both our benchmark sample (1/1975 – 12/2015) and a longer sample (1/1951 – 12/2015), term premia are higher in low inflation countries. Thus, assuming that there is a positive association between average inflation rates and exposure to inflation risk, inflation risk does not account for our findings. This is true not only for our benchmark set of countries, but also for the extended sets of countries. Similarly, the cross-sectional patterns in term premia we observe empirically are not likely to be due to sovereign default risk. As seen in Table 3, countries with high average local currency bond premia have average credit ratings (both unadjusted and adjusted for outlook) that are lower than or similar to the ratings of countries with low average local currency bond premia. That finding is robust to considering different sample periods: it holds both in the long sample period (1/1951 – 12/2015) and in the 7/1989–12/2015 period, during which full ratings are available. Therefore, we find no empirical evidence in favor of an inflation- or credit-based explanation of our findings in this sample of G-10 currencies, and thus pursue a simple interest rate risk interpretation that seems the most relevant, especially over the last thirty years.

2 The Term Structure of Currency Carry Trade Risk Premia: A Challenge

In this section, we show that the downward sloping term structure of currency carry trade risk premia is a challenge even for a reduced-form model.

2.1 The Necessary Condition for Replicating the U.I.P Puzzle

We start with a review of a key necessary condition for replicating the U.I.P puzzle, established by Bekaert (1996) and Bansal (1997) and generalized by Backus, Foresi, and Telmer (2001). To do so, we first introduce some additional notation.

Pricing Kernels and Stochastic Discount Factors The nominal pricing kernel is denoted by $\Lambda_t(\varpi)$; it corresponds to the marginal value of a currency unit delivered at time t in the state of the world ϖ. The nominal SDF M is the growth rate of the pricing kernel: $M_{t+1} = \Lambda_{t+1}/\Lambda_t$. Therefore, the price of a zero-coupon bond that promises one currency unit k periods into the future is given by

$$P_t^{(k)} = E_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right). \tag{4}$$

SDF Entropy SDFs are volatile, but not necessarily normally distributed. In order to measure the time-variation in their volatility, it is convenient to use entropy, rather than variance (Backus, Chernov, and Zin, 2014). The conditional entropy L_t of any random variable X_{t+1} is defined as

$$L_t\left(X_{t+1}\right) = \log E_t\left(X_{t+1}\right) - E_t\left(\log X_{t+1}\right). \tag{5}$$

If X_{t+1} is conditionally lognormally distributed, then the conditional entropy is equal to one half of the conditional variance of the log of X_{t+1}: $L_t\left(X_{t+1}\right) = (1/2)var_t\left(\log X_{t+1}\right)$. If X_{t+1} is not conditionally lognormal, the entropy also depends on the higher moments: $L_t(X_{t+1}) = \kappa_{2t}/2! + \kappa_{3t}/3! + \kappa_{4t}/4! + \ldots$, where $\{\kappa_{it}\}_{i=2}^{\infty}$ are the cumulants of $\log X_{t+1}$.[4]

Exchange Rates When markets are complete, the change in the nominal exchange rate corresponds to the ratio of the domestic to foreign nominal SDFs:

$$\frac{S_{t+1}}{S_t} = \frac{\Lambda_{t+1}}{\Lambda_t}\frac{\Lambda_t^*}{\Lambda_{t+1}^*}. \tag{6}$$

The no-arbitrage definition of the exchange rate comes directly from the Euler equations of the domestic and foreign investors, for any asset return R^* expressed in foreign currency terms: $E_t[M_{t+1}R_{t+1}^* S_t/S_{t+1}] = 1$ and $E_t[M_{t+1}^* R_{t+1}^*] = 1$.

[4]The literature on disaster risk in currency markets shows that higher order moments are critical for understanding currency returns (see Brunnermeier, Nagel, and Pedersen, 2009; Gourio, Siemer, and Verdelhan, 2013; Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan, 2013; Chernov, Graveline, and Zviadadze, 2011).

Currency Risk Premia As Bekaert (1996) and Bansal (1997) show, in models with lognormally distributed SDFs the conditional log currency risk premium $E_t(rx^{FX})$ equals the half difference between the conditional variance of the log domestic and foreign SDFs. This result can be generalized to non-Gaussian economies. When higher moments matter and markets are complete, the currency risk premium is equal to the difference between the conditional entropy of two SDFs (Backus, Foresi, and Telmer, 2001):

$$E_t\left(rx_{t+1}^{FX}\right) = r_t^{f,*} - r_t^f - E_t(\Delta s_{t+1}) = L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+1}^*}{\Lambda_t^*}\right). \tag{7}$$

According to the U.I.P. condition, expected changes in exchange rates should be equal to the difference between the home and foreign interest rates and, thus, the currency risk premium should be zero. In the data, the currency risk premium is as large as the equity risk premium. Any complete market model that addresses the U.I.P. puzzle must thus satisfy a simple necessary condition: high interest rate countries must exhibit relatively less volatile SDFs. In the absence of differences in conditional volatility, complete market models are unable to generate a currency risk premium and the U.I.P. counterfactually holds in the model economy.

Why is the downward term structure of currency carry trade risk premia a challenge for arbitrage-free models? Intuitively, the models need to depart from risk neutrality in order to account for the U.I.P deviations at the short end of the yield curve and large carry trade risk premia. Yet, for the exact same investment horizon, the models need to deliver zero risk premia at the long end of the yield curve, thus behaving as if investors are risk-neutral. In the rest of the paper, we highlight this tension and describe necessary conditions for arbitrage-free models to replicate our empirical evidence.

2.2 An Example: A Reduced-Form Factor Model

We start by showing that even a flexible, N-country reduced-form model calibrated to match the currency carry trade risk premia does not replicate the evidence on long-term bonds. Several two-country models satisfy the condition described in Equation (7) and thus replicate the failure of the U.I.P. condition, but they cannot replicate the portfolio evidence on carry trade risk premia. The reason is simple: when those models are extended to multiple countries, investors in the models can diversify away the country-specific exchange rate risk and there are no cross-sectional differences in carry trade returns across portfolios. To the best of our knowledge, only two models can so far replicate the portfolio evidence on carry trades: the multi-country long-run risk model of Colacito, Croce, Gavazzoni, and Ready (2017) and the multi-country reduced-form factor model of Lustig, Roussanov, and Verdelhan (2011). We focus on the latter because of its flexibility and close forms, and revisit the long-run risk model in Section E of the Appendix, along with other explanations of the U.I.P. puzzle. Moreover, in the Online Appendix, we cover a wide range of term structure models, from the seminal Vasicek (1977) model to the classic Cox, Ingersoll, and Ross (1985) model and to the most recent, multi-factor dynamic term structure models. To save space, we focus here on their most recent international finance version, illustrated in Lustig, Roussanov, and Verdelhan (2014). To replicate the portfolio evidence on carry trades, Lustig, Roussanov, and Verdelhan (2011, 2014) show that no-arbitrage models need to incorporate global shocks to the SDFs along with country heterogeneity in the exposure to those shocks. Following Lustig, Roussanov, and Verdelhan (2014), we consider a world with N countries and currencies in a setup inspired by classic term structure models.

Using their benchmark calibration, we calculate the model-implied term structure of currency risk premia when implementing the slope carry trade strategy (invest in low yield slope currencies, short the high yield slope interest rate currencies). This is very similar to investing in high interest rate countries while borrowing in low interest rate countries. The simulation details are provided in the Appendix. Figure 3, obtained with simulated data, is the model counterpart to Figure 2, obtained with actual data. A clear message emerges: while this model produces U.I.P. deviations (and thus currency risk premia) at the short end of the yield curve, the model produces a flat term structure of currency carry trade risk premia. We turn now to a novel necessary condition that dynamic asset pricing models need to satisfy in order to generate a downward-sloping term structure.

3 Foreign Long-Term Bond Returns and the Properties of SDFs

In this section, we derive a novel, preference-free necessary condition that complete market models need to satisfy in order to reproduce the downward sloping term structure of currency carry trade risk premia. To do so, we first review a useful decomposition of the pricing kernel.

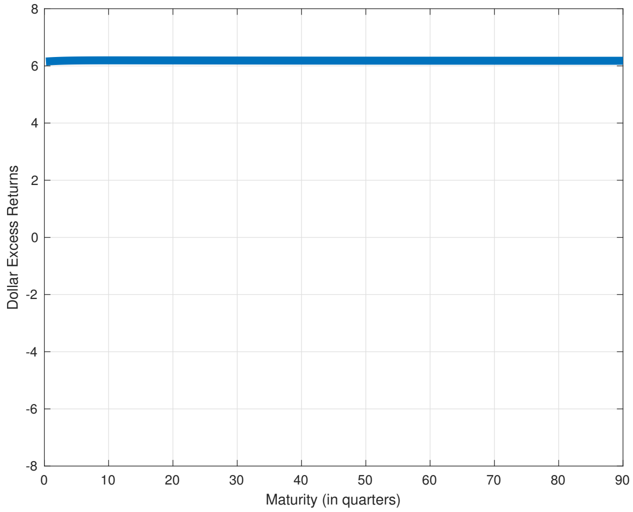

Figure 3: Simulated Long-Minus-Short Foreign Bond Risk Premia in U.S. Dollars— The figure shows the simulated average dollar log excess return of the slope carry trade strategy as a function of the bond maturities in the reduced-form model of Lustig, Roussanov, and Verdelhan (2014). At each date t, currencies are sorted into three portfolios by the slope of their yield curve (measured as the difference between the 10-year and the three-month yields). The first portfolio contains the currencies of countries with low yield slopes, while the third portfolio contains the currencies of countries with high yield slope. The slope carry trade strategy invests in the first portfolio and shorts the third portfolio. The model is simulated at the monthly frequency. The holding period is one month and returns are annualized.

3.1 Pricing Kernel Decomposition

Our results build on the Alvarez and Jermann (2005) decomposition of the pricing kernel Λ_t into a permanent component $\Lambda_t^{\mathbb{P}}$ and a transitory component $\Lambda_t^{\mathbb{T}}$ using the price of the long-term bond:

$$\Lambda_t = \Lambda_t^{\mathbb{P}} \Lambda_t^{\mathbb{T}}, \text{ where } \Lambda_t^{\mathbb{T}} = \lim_{k\to\infty} \frac{\delta^{t+k}}{P_t^{(k)}}, \tag{8}$$

where the constant δ is chosen to satisfy the following regularity condition: $0 < \lim_{k\to\infty} \frac{P_t^{(k)}}{\delta^k} < \infty$ for all t. Note that $\Lambda_t^{\mathbb{P}}$ is equal to:

$$\Lambda_t^{\mathbb{P}} = \lim_{k\to\infty} \frac{P_t^{(k)}}{\delta^{t+k}} \Lambda_t = \lim_{k\to\infty} \frac{E_t(\Lambda_{t+k})}{\delta^{t+k}}.$$

The second regularity condition ensures that the expression above is a martingale. Alvarez and Jermann (2005) assume that, for each $t+1$, there exists a random variable x_{t+1} with finite expected value $E_t(x_{t+1})$ such that almost surely $\frac{\Lambda_{t+1}}{\delta^{t+1}} \frac{P_{t+1}^{(k)}}{\delta^k} \leq x_{t+1}$ for all k. Under those regularity conditions, the infinite-maturity bond return is:

$$R_{t+1}^{(\infty)} = \lim_{k\to\infty} R_{t+1}^{(k)} = \lim_{k\to\infty} P_{t+1}^{(k-1)}/P_t^{(k)} = \frac{\Lambda_t^{\mathbb{T}}}{\Lambda_{t+1}^{\mathbb{T}}}. \tag{9}$$

The permanent component, $\Lambda_t^{\mathbb{P}}$, is a martingale and is an important part of the pricing kernel: Alvarez and Jermann (2005) derive a lower bound of its volatility and, given the size of the equity premium relative to the term premium, conclude that it accounts for most of the SDF volatility.[5] In other words, a lot of persistence in the pricing kernel is needed to jointly deliver a low term premium and a high equity premium. Throughout this paper we assume that stochastic discount factors $\frac{\Lambda_{t+1}}{\Lambda_t}$ and returns R_{t+1} are jointly stationary.

3.2 Main Preference-Free Result on Long-Term Bond Returns

We now use this pricing kernel decomposition to understand the properties of the dollar returns of long-term bonds. Recall that the dollar term premium on a foreign bond position, denoted by $E_t[rx_{t+1}^{(k),\$}]$, can be expressed as the sum of foreign term premium in local currency terms, $E_t[rx_{t+1}^{(k),*}]$, plus a currency risk premium, $E_t[rx_{t+1}^{FX}] = r_t^{f,*} - r_t^f - E_t[\Delta s_{t+1}]$. Here, we consider the dollar term premium of an infinite-maturity foreign bond, so we let $k \to \infty$.

Proposition 1. *If financial markets are complete, the foreign term premium on the long-term bond in dollars is equal to the domestic term premium plus the difference between the domestic and foreign entropies of the permanent components of the pricing kernels:*

$$E_t[rx_{t+1}^{(\infty),\$}] = E_t\left[rx_{t+1}^{(\infty)}\right] + L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) - L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}\right). \tag{10}$$

To intuitively link the long-run properties of pricing kernels to foreign bond returns and exchange rates, let us consider the simple benchmark of countries represented by stand-in agents with power utility and *i.i.d.*

[5]Proposition 2 in Alvarez and Jermann (2005) establishes that $L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) \geq E_t \log R_{t+1} - \log R_{t+1}^{\infty}$ for any return R_{t+1} and that $\frac{L\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right)}{L\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)} \geq Min\{1, \frac{E\log R_{t+1}/R_t^f - E\log R_{t+1}^{\infty}/R_t^f}{E\log R_{t+1}/R_t^f + L(1/R_t^f)}\}$ for any positive return R_{t+1} such that $E\log R_{t+1}/R_t^f + L(1/R_t^f) > 0$. Alvarez and Jermann (2005) take the latter expression to the data and report several lower bounds for the relative variance of the permanent component in their Table 2, page 1989. These lower bounds, obtained with either yields or holding-period returns on long-term bonds, range from 0.76 to 1.11. Thus, the variance of the permanent component is at least 76% of the total variance of the SDF.

consumption growth rates. In that case, all SDF shocks are permanent ($\Lambda_t = \Lambda_t^{\mathbb{P}}$ for all t). As we shall see, such model is counterfactual. In this model, the risk-free rate is constant, so bonds of different maturities offer the same returns. Foreign bond investments differ from domestic bond investments only because of the presence of exchange rate risk and, since consumption growth rates are $i.i.d$, exchange rates are stationary in changes but not in levels. Finally, carry trade excess returns are the same at the short end, (see Equation (7)) and at the long end (see Equation (10)) of the yield curve, so the term structure of currency carry trade risk premia is flat. A power utility model with only permanent shocks cannot match the facts.

Let us now turn to the opposite case: a model without permanent shocks in the SDF ($\Lambda_t = \Lambda_t^{\mathbb{T}}$ for all t). In case of an adverse temporary innovation to the foreign pricing kernel, the foreign currency appreciates, so a domestic position in the foreign bond experiences a capital gain. However, this capital gain is exactly offset by the capital loss suffered on the long-term bond as a result of the increase in foreign interest rates. Hence, interest rate exposure completely hedges the temporary component of the currency risk exposure. In this case, as Equation (10) shows, the long-term bond risk premium in dollars should be equal to the domestic term premium.

Beyond these two polar cases, Proposition 1 shows that in order to have differences across countries in bond risk premia, once converted in the same currency, no-arbitrage models need conditional entropy differences in the permanent component of their pricing kernels. If the domestic and foreign pricing kernels have identical conditional entropy, then high local currency term premia are always associated with low currency risk premia and vice-versa, so dollar term premia are identical across currencies.

Proposition 1 is thus the bond equivalent to the usual currency carry trade condition. We gather them below to emphasize their similarities:

$$\begin{aligned} E_t\left(rx_{t+1}^{FX}\right) &= L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+1}^*}{\Lambda_t^*}\right), \\ E_t[rx_{t+1}^{(\infty),\$}] - E_t\left[rx_{t+1}^{(\infty)}\right] &= L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) - L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}\right). \end{aligned}$$

To reproduce large currency carry trade risk premia, no-arbitrage models need large differences in the volatilities of their SDFs. To reproduce the absence of dollar bond risk premia, no-arbitrage models need to feature the same volatilities of the martingale components of their SDFs. As will shall see, this condition is a key tool to assess existing international finance models.

3.3 Additional Assumptions and Interpretation

Dynamic asset pricing models that generate small amounts of dollar return differential predictability may produce moments in small samples that fall within the confidence intervals, but it is useful to have a clear benchmark: Ours is no predictability in dollar bond return differentials for long bonds. This has been the null hypothesis in this literature. Our paper shows that this null cannot be rejected at longer maturities.

In order to use Proposition 1 to interpret our empirical findings, two additional assumptions are required.

Assumption 1: First, since very long-term bonds are rarely available or liquid, we assume that infinite-maturity bond returns can be approximated in practice by 10 and 15-year bond returns. The same assumption is also present in Alvarez and Jermann (2005), Hansen, Heaton, and Li (2008), Hansen and Scheinkman (2009), and Hansen (2012). It is supported by the simulation of the state-of-the-art Joslin, Singleton, and Zhu (2011) term structure model (see section H of the Online Appendix).

Assumption 2: Second, we assume that the level and slope of the yield curve summarize all the relevant information that investors use to forecast dollar bond excess returns. Proposition 1 pertains to *conditional* risk premia and is, thus, relevant for interpreting our empirical time series predictability results and the average excess returns of currency portfolios sorted by conditioning information (the level of the short-term interest rate or the slope of the yield curve). Building portfolios sorted by conditioning variables is a flexible, non-parametric approach to bringing in conditioning information. We cannot definitively rule out the possibility that there are other predictors. In all of the models that we consider, no other predictors exist (except in pathological, knife-edge cases), but this may not be true in other models. In particular, there is a lively empirical debate on whether there are unspanned macro variables that have incremental out-of-sample forecasting power for bond returns (see Bauer and Hamilton, 2017, for a thorough evaluation of the empirical evidence).

Under those two assumptions, a simple condition illustrates our empirical findings:

Condition 1. *In order for the conditional dollar term premia on infinite-maturity bonds to be identical across countries, when financial markets are complete, the conditional entropy of the permanent SDF component has to be identical across countries:* $L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = L_t\left(\frac{\Lambda^{\mathbb{P},*}_{t+1}}{\Lambda^{\mathbb{P},*}_t}\right)$*, for all* t.

If this condition fails, under Assumptions 1 and 2, portfolios sorted on conditioning variables produce non-zero currency carry trade risk premia at the long end of the term structure, as the conditional dollar term premia of long-maturity bonds differ across countries.[6] Condition 1 is satisfied when permanent shocks are common across countries ($\Lambda_{t+1}^{\mathbb{P}} = \Lambda_{t+1}^{\mathbb{P},*}$ for all t) and thus, in the absence of permanent shocks, when exchange rates are stationary in levels (up to a deterministic time trend). But note that the stationarity is sufficient but not necessary to satisfy Condition 1.

To develop some intuition for this condition, we rely on an example from Alvarez and Jermann (2005), who consider a model with conditionally log-normally distributed pricing kernels driven by both permanent and transitory shocks.

Example 1. *Consider the following pricing kernel (Alvarez and Jermann, 2005):*

$$\begin{aligned}
\log \Lambda_{t+1}^{\mathbb{P}} &= -\frac{1}{2}\sigma_P^2 + \log \Lambda_t^{\mathbb{P}} + \varepsilon_{t+1}^P, \\
\log \Lambda_{t+1}^{\mathbb{T}} &= \log \beta^{t+1} + \sum_{i=0}^{\infty} \alpha_i \varepsilon_{t+1-i}^T,
\end{aligned}$$

where α is a square summable sequence, and ε^P and ε^T are serially independent and normally distributed random variables with mean zero, variance σ_P^2 and σ_T^2, respectively, and covariance σ_{TP}. A similar decomposition applies to the foreign pricing kernel.

In this economy, Alvarez and Jermann (2005) show that the domestic term premium is given by the following expression: $E_t\left[rx_{t+1}^{(\infty)}\right] = \frac{1}{2}\alpha_0^2\sigma_T^2 + \alpha_0\sigma_{TP}$. Only transitory risk is priced in the market for long-maturity bonds: when marginal utility is transitorily high, interest rates increase because the representative agent wants to borrow, so long-term bonds suffer a capital loss. Permanent shocks to marginal utility do not affect the prices of long-term bonds at all. Similarly, the foreign term premium, in local currency terms, is $E_t\left[rx_{t+1}^{(\infty),*}\right] = \frac{1}{2}\left(\alpha_0^*\sigma_T^*\right)^2 + \alpha_0^*\sigma_{TP}^*$. The currency risk premium is the difference in the two countries' conditional SDF entropy:

$$E_t\left[rx_{t+1}^{FX}\right] = r_t^{f,*} - r_t^f - E_t[\Delta s_{t+1}] = \frac{1}{2}\left(\alpha_0^2\sigma_T^2 + 2\alpha_0\sigma_{TP} + \sigma_P^2\right) - \frac{1}{2}\left(\left(\alpha_0^*\sigma_T^*\right)^2 + 2\alpha_0\sigma_{TP}^* + \left(\sigma_P^*\right)^2\right).$$

[6] For some countries (Australia, Canada, and Sweden), time-series regressions show that the yield slopes predict dollar bond returns with the "wrong" sign: while an increase in the yield slope decreases short-bond carry trade excess returns, it increases the dollar long-bond excess returns (cf Table 1). To match this particular evidence, one may replace Condition 1 with the following condition: if $E_t\left(rx_{t+1}^{FX}\right) = L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+1}^*}{\Lambda_t^*}\right) > 0$, then $E_t[rx_{t+1}^{(\infty),\$}] - E_t\left[rx_{t+1}^{(\infty)}\right] = L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) - L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}\right) < 0$. We do not study this stricter condition as the amount of predictability on long-bond dollar excess returns is not economically significant, as can be seen in Table 2.

As a result, the foreign term premium in dollars, given by Equation (7), is:

$$E_t\left[rx_{t+1}^{(\infty),\$}\right] = E_t\left[rx_{t+1}^{(\infty),*}\right] + E_t\left[rx_t^{FX}\right] = \frac{1}{2}\alpha_0^2\sigma_T^2 + \alpha_0\sigma_{TP} + \frac{1}{2}\left(\sigma_P^2 - (\sigma_P^*)^2\right).$$

In the Alvarez and Jermann (2005) example, Condition 1 is satisfied, provided that $\sigma_P^2 = (\sigma_P^*)^2$. Then the foreign term premium in dollars equals the domestic term premium:

$$E_t\left[rx_{t+1}^{(\infty),\$}\right] = \frac{1}{2}\alpha_0^2\sigma_T^2 + \alpha_0\sigma_{TP} = E_t\left[rx_{t+1}^{(\infty)}\right].$$

This example illustrates our theoretical and empirical findings. It shows how SDFs can deliver carry trade risk premia with Treasury bills, when $\alpha_0 \neq \alpha_0^*, \sigma_T \neq \sigma_T^*$ or $\sigma_{TP} \neq \sigma_{TP}^*$, while producing no carry trade risk premia with long-term bonds when $\sigma_P = \sigma_P^*$. It also shows that exchange rate stationarity (i.e. when $\varepsilon_{t+1}^P = \varepsilon_{t+1}^{*,P}$ and) is a sufficient but not a necessary condition to produce no carry trade risk premia with long-term bonds. This example, however, lacks the time variation in risk premia that has been extensively documented in equity, bond, and currency markets. We turn now to a second example, with time-varying risk premia.

In order to illustrate the SDF decomposition and Condition 1 in a very transparent setting, we focus here on a simple two-country version of the one-factor Cox, Ingersoll, and Ross (1985) model with only one kind of shocks per country.

Example 2. *The two-country Cox, Ingersoll, and Ross (1985) model is defined by the following law of motions for the SDFs:*

$$\begin{aligned} -\log\frac{\Lambda_{t+1}}{\Lambda_t} &= \alpha + \chi z_t + \sqrt{\gamma z_t}u_{t+1}, \\ z_{t+1} &= (1-\phi)\theta + \phi z_t - \sigma\sqrt{z_t}u_{t+1}, \end{aligned}$$

where a similar law of motion applies to the foreign SDF (with foreign variables and parameters denoted with an **), and where* z_t *and* z_t^* *are the two state variables that govern the volatilities of the normal shocks* u_{t+1} *and* u_{t+1}^*.

The domestic risk-free rate is given by $r_t^f = \alpha + (\chi - \frac{1}{2}\gamma)z_t$. The log bond prices are affine in the state variable z:

$p_t^{(n)} = -B_0^n - B_1^n z_t$, where the bond price coefficients evolve according to the second-order difference equations:

$$\begin{aligned}
B_0^n &= \alpha + B_0^{n-1} + B_1^{n-1}(1-\phi)\theta, \\
B_1^n &= \chi - \frac{1}{2}\gamma + B_1^{n-1}\phi - \frac{1}{2}\left(B_1^{n-1}\right)^2\sigma^2 + \sigma\sqrt{\gamma}B_1^{n-1}.
\end{aligned}$$

Therefore, the transitory component of the pricing kernel is $\Lambda_t^{\mathbb{T}} = \lim_{n\to\infty}\frac{\delta^{t+n}}{P_t^{(n)}} = \lim_{n\to\infty}\left(\delta^{t+n}e^{B_0^n + B_1^n z_t}\right)$, so the temporary and martingale components of the SDF are:

$$\begin{aligned}
\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} &= \delta e^{B_1^\infty[(\phi-1)(z_t-\theta)-\sigma\sqrt{z_t}u_{t+1}]}, \\
\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} &= \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = \frac{1}{\delta}e^{-\alpha-\chi z_t - B_1^\infty(\phi-1)(z_t-\theta)-\sqrt{\gamma z_t}u_{t+1}+B_1^\infty\sigma\sqrt{z_t}u_{t+1}}.
\end{aligned}$$

where the constant δ is chosen in order to satisfy: $0 < \lim_{n\to\infty}\frac{P_t^{(n)}}{\delta^n} < \infty$. The limit of $B_0^n - B_0^{n-1}$ is finite: $\lim_{n\to\infty}(B_0^n - B_0^{n-1}) = \alpha + B_1^\infty(1-\phi)\theta$, where B_1^∞ is defined implicitly in the second order equation $B_1^\infty = \chi - \frac{1}{2}\gamma + B_1^\infty\phi - \frac{1}{2}\left(B_1^\infty\right)^2\sigma^2 + \sigma\sqrt{\gamma}B_1^\infty$. As a result, B_0^n grows at a linear rate in the limit. We choose the constant δ to offset the growth in B_0^n as n becomes very large. Setting $\delta = e^{-\alpha - B_1^\infty(1-\phi)\theta}$ guarantees that Assumption 1 in Alvarez and Jermann (2005) is satisfied. The temporary component of the pricing kernel is thus equal to: $\Lambda_t^{\mathbb{T}} = \delta^t e^{B_1^\infty z_t}$.

This reduced-form model shows the difference between unconditional and conditional risk premia. In the case of a symmetric model ($\sigma = \sigma^*$, $\phi = \phi^*$, $\theta = \theta^*$), the unconditional risk currency risk premium is zero: $E[rx_{t+1}^{FX}] = \frac{1}{2}\gamma(E[z_t] - E[z_t^*]) = 0$, while the conditional currency risk premium is not: $E_t[rx_{t+1}^{FX}] = \frac{1}{2}\gamma(z_t - z_t^*) \neq 0$. The conditional risk premium moves with the two state variables z_t and z_t^*. These variables take different values across countries because countries experience different shocks. Our work is about the conditional risk premium: the portfolios are built by sorting on the level and slope of interest rates, which in this model are driven by the state variables z_t and z_t^*. When taking the average of all returns in the high interest rate portfolio for example, we are averaging over low values of the state variables z_t, not over all possible values of z_t. In this symmetric model, the average currency risk premium obtained by simply averaging all returns would be zero. Yet, as its simulation show in the previous section, the average currency risk premium obtained on the high interest rate portfolio (for example) is as large as in the data. Our portfolio and predictability tests thus are about conditional risk premia, as is our preference-free condition.

In this two-country model, Condition 1 requires that $\left(\sqrt{\gamma} - B_1^{\infty}\sigma\right) z_t = \left(\sqrt{\gamma^*} - B_1^{\infty *}\sigma^*\right) z_t^*$. We leave the study of these restrictions for the Online Appendix because this one-factor model is too simple. As noted in Campbell (2017), the Cox, Ingersoll, and Ross (1985) model implies that large yield spreads predict low term premia, contrary to the data. Condition 1, however, is a diagnostic tool that can be applied to richer models that are not subject to this criticism. In the Online Appendix, we derive parametric restrictions to implement Condition 1 in four classes of dynamic term structure models, from the simple one-factor Vasicek (1977) and Cox, Ingersoll, and Ross (1985) models to their multi-factor versions. We also sketch there a multi-factor version of the Cox, Ingersoll, and Ross (1985) model with "temporary" and "permanent" factors that can reproduce the downward term structure of carry risk premia.

In the Online Appendix, we also consider leading structural macro-finance models. We focus on models of the real SDF, because (i) there is no evidence to suggest that inflation risk can account for UIP deviations, and (ii) inflation risk does not seem to account for the cross-country variation in local currency bond excess returns, as shown in Section 1. We consider two-country versions of the habit, long-run risk, and disaster risk models. In a nutshell, among the reduced-form term structure models we consider, Condition 1 implies novel parameter restrictions for all models (and in some cases, it rules out all permanent shocks or the time-variation in the price of risk). In order to save space, we summarize the implications of Condition 1 in Table 4. All the intermediary steps to determine the two components of the pricing kernels are reported in the Appendix. In the habit model with common shocks, the carry trade risk premia and Condition 1 requires that countries exhibit the same risk-aversion and the same volatility of consumption growth shocks but they can differ in the persistence of their habit levels. The long-run risk models satisfy Condition 1 only with common shocks and for knife-edge parameter values. This model generically produces non-stationary real exchange rates. For the disaster risk models, common shocks are also necessary for Condition 1 to hold and the downward term structure of carry trade risk premia is consistent with heterogeneity in the rate of time preference, the rate of depreciation, or the country-specific growth rate, but no heterogeneity in the coefficient of risk aversion, the common and country-specific consumption drops in case of a disaster, and the probability of a disaster. Overall, the term structure of carry trade raises the bar for international finance models.

Table 4: Condition 1: Dynamic Asset Pricing Model Scorecard

	Symmetric Models with Country-specific Shocks	*Asymmetric Models with Common Shocks*
External Habit Model	Condition 1 is satisfied.	Condition 1 is satisfied, but the only heterogeneity across countries is in ϕ. No heterogeneity possible in (γ, σ^2).
Long Run Risks Model	Condition 1 is never satisfied	Condition 1 is satisfied, but only knife-edge cases.
Disaster Model	Condition 1 is never satisfied	Condition 1 is satisfied, but the only heterogeneity across countries is in (R, λ, g_w). No heterogeneity possiblein (γ, B, F, p_t).

This table summarizes whether each class of models can satisfy Condition 1 and what are the potential sources of heterogeneity across countries. The left section of the table focuses on models with only country-specific shocks in which all countries have the same parameters. The right section focuses on models with only common shocks and heterogeneity in the parameters. In the external habit model (Campbell and Cochrane, 1999; Wachter, 2006; Verdelhan, 2010; Stathopoulos, 2017), the parameters ϕ and B govern the dynamics of the surplus consumption ratio process. In the long run risk model (Bansal and Yaron, 2004; Colacito and Croce, 2011; Bansal and Shaliastovich, 2013; Engel, 2016), Condition 1 is always violated except in knife-edge cases. In the disaster model (Farhi and Gabaix, 2016; Gabaix, 2012; Wachter, 2013), the parameters R, λ, and g_w govern the rate of time preference, the rate of depreciation and the country-specific growth rate, while the parameters γ, B, F, p_t are the coefficient of risk aversion, the common consumption drop in case of a disaster, the country-specific consumption drop in case of a disaster, and the probability of a disaster. The details of these models are in the Online Appendix.

4 U.I.P. in the Long Run

Examining the conditional moments of one-period returns on long-maturity bonds, the focus of our paper, is not equivalent to studying the moments of long-maturity bond yields in tests of the long-horizon U.I.P. condition. In this section, we show the links and differences between these moments. To do so, we use again the decomposition of the pricing kernel proposed by Alvarez and Jermann (2005). Exchange rate changes can be represented as the product of two components, defined below:

$$\frac{S_{t+1}}{S_t} = \left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\frac{\Lambda^{\mathbb{P},*}_t}{\Lambda^{\mathbb{P},*}_{t+1}}\right)\left(\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\frac{\Lambda^{\mathbb{T},*}_t}{\Lambda^{\mathbb{T},*}_{t+1}}\right) = \frac{S^{\mathbb{P}}_{t+1}}{S^{\mathbb{P}}_t}\frac{S^{\mathbb{T}}_{t+1}}{S^{\mathbb{T}}_t}. \tag{11}$$

Exchange rate changes reflect differences in both the transitory and the permanent component of the two countries' pricing kernels.[7] If two countries share the same martingale component of the pricing kernel, then the resulting exchange rate is stationary and Condition 1 is trivially satisfied. However, as already mentioned, exchange rate stationarity is obviously not necessary for Condition 1 to hold. This exchange rate decomposition implies a lower bound on the cross-country correlation of the permanent components of the SDFs. In the interest of space, we present it in the Online Appendix and focus in the main text on its implications for long-run U.I.P.

The definition of the transitory component of exchange rate changes, given by $\Delta s_{t+1}^{\mathbb{T}} = \left(\lambda_{t+1}^{\mathbb{T}} - \lambda_t^{\mathbb{T}}\right) - \left(\lambda_{t+1}^{\mathbb{T},*} - \lambda_t^{\mathbb{T},*}\right)$, where $\lambda_t^{\mathbb{T}} \equiv \log \Lambda_t^{\mathbb{T}} = \lim_{k\to\infty}(t+k)\log\delta + \lim_{k\to\infty} k y_t^{(k)}$, implies that a currency experiences a temporary appreciation when its long-term interest rates increase more than the foreign ones:

$$\Delta s_{t+1}^{\mathbb{T}} = \log\delta - \log\delta^* + \lim_{k\to\infty} k\left(\Delta y_{t+1}^{(k)} - \Delta y_{t+1}^{(k),*}\right). \tag{12}$$

By backward substitution, it follows that the transitory component of the exchange rate in levels is given by the spread in long-term yields:

$$s_t^{\mathbb{T}} = s_0 + t(\log\delta - \log\delta^*) + \lim_{k\to\infty} k\left(y_t^{(k)} - y_t^{(k),*}\right) - \lim_{k\to\infty} k\left(y_0^{(k)} - y_0^{(k),*}\right). \tag{13}$$

This decomposition of exchange rates implies that deviations from the long-run UIP are due to the permanent component of exchange rates. We use $rx_{t\to t+k}^{FX} = k\left(y_t^{(k),*} - y_t^{(k)}\right) - \Delta s_{t\to t+k}$ to denote the currency excess return over a holding period of k years.

Proposition 2. *When financial markets are complete, the expected rate of transitory depreciation per period is equal to the spread in the long-term interest rates:*

$$\lim_{k\to\infty}\frac{1}{k}E_t\left[\Delta s_{t\to t+k}^{\mathbb{T}}\right] = -\lim_{k\to\infty}\left(y_t^{(k)} - y_t^{(k),*}\right).$$

Thus, the (per period) deviation from long-run U.I.P is the (per period) permanent component of the exchange

[7] Note that $S_{t+1}^{\mathbb{P}}$, the ratio of two martingales, is itself not a martingale in general. However, in the class of affine term structure models, this exchange rate component is indeed a martingale.

rate change:

$$\lim_{k\to\infty}\frac{1}{k}E_t\left[rx^{FX}_{t\to t+k}\right]=\lim_{k\to\infty}\left(y_t^{(k),*}-y_t^{(k)}\right)-\lim_{k\to\infty}\frac{1}{k}E_t\left[\Delta s_{t\to t+k}\right]=-\lim_{k\to\infty}\frac{1}{k}E_t\left[\Delta s^{\mathbb{P}}_{t\to t+k}\right].$$

If exchange rates are stationary in levels, in which case the permanent component of exchange rate changes is zero, then per period long-run U.I.P. deviations converge to zero. In this case, the slope coefficient in the regression of per period long-run exchange rate changes on yield differences converges to one and the intercept converges to zero. This result is previewed in Backus, Boyarchenko, and Chernov (2016), who show that claims to stationary cash flows earn a zero per period log risk premium over long holding periods. It follows that long-run deviations from U.I.P. are consistent with no arbitrage only if the exchange rate is not stationary in levels.

To illustrate Proposition 2, let us study two examples where exchange rates are stationary in levels. We first go back to Example 1 and assume that each country's pricing kernel has no permanent innovations (i.e., $\varepsilon_t^P=0$ for all t, so $\Lambda_t^{\mathbb{P}}=1$ for all t, and similarly for the foreign pricing kernel). In that case, the two pricing kernels and the exchange rate are stationary, due to the square summability of sequences α and α^*. Bond yields satisfy $y_t^{(k)}=-\log\beta-\frac{1}{k}\frac{\sigma_T^2}{2}\sum_{i=0}^{k-1}\alpha_i^2-\frac{1}{k}\sum_{i=0}^{\infty}(\alpha_{k+i}-\alpha_i)\varepsilon_{t-i}$ and the expected rate of per period transitory depreciation is $\frac{1}{k}E_t[\Delta s^{\mathbb{T}}_{t\to t+k}]=\frac{1}{k}E_t[\Delta s_{t\to t+k}]=\log\beta-\log\beta^*+\frac{1}{k}\sum_{i=0}^{\infty}(\alpha_{k+i}-\alpha_i)\varepsilon^T_{t-i}-\frac{1}{k}\sum_{i=0}^{\infty}(\alpha^*_{k+i}-\alpha^*_i)\varepsilon^{T,*}_{t-i}$. In the limit of $k\to\infty$, Proposition 2 holds as both sides converge to zero, using the property $\lim_{k\to\infty}\alpha_k=\lim_{k\to\infty}\alpha_k^*=0$ that arises from the square summability of sequences α and α^*.

We can also consider the symmetric two-country CIR model with country-specific factors presented in Example 2. In that model, the transitory component of the exchange rate is given by: $s_t^{\mathbb{T}}=s_0+B_1^{\infty}\left((z_t-z_0)-(z_t^*-z_0^*)\right)$. As already noted, the pricing kernel is not subject to permanent shocks when $B_1^{\infty}=\frac{\sqrt{\gamma}}{\sigma}=\frac{\chi}{1-\phi}$.[8] In that case, the exchange rate is stationary and hence $s_t=s_t^{\mathbb{T}}$. The expected rate of depreciation is then equal to

$$\lim_{k\to\infty}\frac{1}{k}E_t[\Delta s_{t\to t+k}]=-\frac{\chi}{1-\phi}(z_t-z_t^*)=-\lim_{k\to\infty}\left(y_t^{(k)}-y_t^{(k),*}\right).$$

[8]The restrictions $B_1^{\infty}=\frac{\sqrt{\gamma}}{\sigma}=\frac{\chi}{1-\phi}$ have a natural interpretation as restrictions on the long-run loadings of the exchange rate on the risk factors: $\sum_{i=1}^{\infty}E_t[\Delta s_{t+i}]=\sum_{i=1}^{\infty}E_t[m_{t+i}-m^*_{t+i}]=\sum_{i=1}^{\infty}\phi^{i-1}\chi(z_t^*-z_t)$.

Even if exchange rates are not stationary in levels, as long as unconditional SDF entropies are the same across countries, then (under some additional regularity conditions), long-run U.I.P holds on average, as the following proposition shows.

Proposition 3. *If financial markets are complete, the stochastic discount factors* $\frac{\Lambda_{t+1}}{\Lambda_t}$ *and* $\frac{\Lambda^*_{t+1}}{\Lambda^*_t}$ *are strictly stationary, and* $\lim_{k\to\infty}\frac{1}{k}L\left(E_t\left[\frac{\Lambda_{t+k}}{\Lambda_t}\right]\right)=0$ *and* $\lim_{k\to\infty}\frac{1}{k}L\left(E_t\left[\frac{\Lambda^*_{t+k}}{\Lambda^*_t}\right]\right)=0$*, then the per period long-run currency risk premium is given by:*

$$\lim_{k\to\infty}\frac{1}{k}E[rx^{FX}_{t\to t+k}]=\lim_{k\to\infty}E\left(y^{(k),*}_t-y^{(k)}_t\right)-\lim_{k\to\infty}\frac{1}{k}E[\Delta s_{t\to t+k}]=\left[L\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right)-L\left(\frac{\Lambda^{\mathbb{P},*}_{t+1}}{\Lambda^{\mathbb{P},*}_t}\right)\right].$$

This immediately implies that the per period currency risk premium converges to zero on average, and therefore long-run U.I.P. holds on average, if Condition 1 is satisfied. Importantly, unconditional long-run U.I.P. does not necessarily require stationary exchange rates, as Condition 1 is a weaker condition than exchange rate stationarity. Moreover, Condition 1 may hold while deviations from long-run U.I.P. still exist. Proposition 3 only implies that long-run U.I.P. holds on average, not at any date. As a result, the study of holding period returns of long-term bonds, the topic of this paper, is not the same as the study of U.I.P. in the long-run.

5 Conclusion

While holding period bond returns, expressed in a common currency, differ across G10 countries at the short end of the yield curve (the U.I.P. puzzle), they are rather similar at the long end. In other words, the term structure of currency carry trade risk premia is downward-sloping. Replicating such a term structure is a non-trivial for most models: recent no-arbitrage models of international finance that are able to address the U.I.P. puzzle fail to replicate the downward-sloping term structure of carry trade risk premia.

We derive a preference-free result that helps assess existing models and guides future theoretical and empirical work. In order to exhibit similar long-term bond returns when expressed in the same units, complete market models need to exhibit the same volatility of the permanent components of their pricing kernels. This condition implies novel parameter restrictions in the workhorse no-arbitrage models of international finance.

Our results show that exchange rate risk is different from equity and bond risk. In order to account for the high equity premium and the low term premium, most of the variation in the marginal utility of wealth in no-arbitrage models must come from *permanent* shocks. Yet, differences across countries in how *temporary* shocks affect the marginal utility of wealth — and thus exchange rates — appear as natural way to align no-arbitrage models with the downward-sloping term structure of carry trade risk premia.

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APPENDIX

A Data

Our zero-coupon bond dataset covers the same benchmark sample of G10 countries, but from at most 1/1975 to 12/2015. We use the entirety of the dataset in Wright (2011) and complement the sample, as needed, with sovereign zero-coupon curve data sourced from Bloomberg, estimated from government notes and bonds as well as interest rate swaps of different maturities. The panel is unbalanced: for each currency, the sample starts with the beginning of the Wright (2011) dataset. The starting dates for each country are as follows: 2/1987 for Australia, 1/1986 for Canada, 1/1973 for Germany, 1/1985 for Japan, 1/1990 for New Zealand, 1/1998 for Norway, 12/1992 for Sweden, 1/1988 for Switzerland, 1/1979 for the U.K., and 12/1971 for the U.S. For New Zealand, the data for maturities above 10 years start in 12/1994. Yields are available for bond maturities ranging from three months to 15 years, in three-month increments.

To construct averages of credit ratings, we assign each rating to a number, with a smaller number corresponding to a higher rating. In particular, a credit rating of AAA corresponds to a numerical value of 1, with each immediately lower rating getting assigned the immediately higher numerical value: AA+ corresponds to a numerical value of 2 and AA to 3, all the way down to CC- (22) and SD (23). We also construct rating series adjusted for outlook, as follows: a 'Negative' outlook corresponds to an upward adjustment of 0.5 in the numerical value of the rating, a 'Watch Negative' outlook to an upward adjustment of 0.25, a 'Stable' or 'Satisfactory' outlook to no adjustment, and a 'Positive', 'Strong' or 'Very Strong' outlook to a downward adjustment of 0.5. For example, a credit rating of BB (coded as 12) receives a numerical value of 12.5 with a 'Negative' outlook and a value of 11.5 with a 'Positive' outlook. In order to construct credit rating averages for portfolios formed before 7/1989, we backfill each country's credit rating by assuming that the country's rating before 7/1989 is equal to its rating at the first available observation.

B Proofs

- Proof of Proposition 1:

Proof. The proof builds on some results in Backus, Foresi, and Telmer (2001) and Alvarez and Jermann (2005). Specifically, Backus, Foresi, and Telmer (2001) show that the foreign currency risk premium is equal to the difference between domestic and foreign total SDF entropy:

$$(f_t - s_t) - E_t[\Delta s_{t+1}] = L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\right).$$

Furthermore, Alvarez and Jermann (2005) establish that total SDF entropy equals the sum of the entropy of the permanent pricing kernel component and the expected log term premium:

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) + E_t\left(\log \frac{R^{(\infty)}_{t+1}}{R^f_t}\right).$$

Applying the Alvarez and Jermann (2005) decomposition to the Backus, Foresi, and Telmer (2001) expression yields the desired result.

To derive the Backus, Foresi, and Telmer (2001) expression, consider a foreign investor who enters a forward position in the currency market with payoff $S_{t+1} - F_t$. The investor's Euler equation is:

$$E_t\left(\frac{\Lambda^*_{t+1}}{\Lambda^*_t}(S_{t+1} - F_t)\right) = 0.$$

In the presence of complete, arbitrage-free international financial markets, exchange rate changes equal the ratio of the domestic and foreign pricing kernels:

$$\frac{S_{t+1}}{S_t} = \frac{\Lambda_{t+1}}{\Lambda_t}\frac{\Lambda^*_t}{\Lambda^*_{t+1}},$$

Dividing the investor's Euler equation by S_t and applying the no arbitrage condition, the forward discount is:

$$f_t - s_t = \log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - \log E_t\left(\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\right).$$

The second component of the currency risk premium is expected foreign appreciation; applying logs and conditional expectations to the no arbitrage condition above leads to:

$$E_t[\Delta s_{t+1}] = E_t\left(\log \frac{\Lambda_{t+1}}{\Lambda_t}\right) - E_t\left(\log \frac{\Lambda^*_{t+1}}{\Lambda^*_t}\right).$$

Combining the two terms of the currency risk premium leads to:

$$(f_t - s_t) - E_t[\Delta s_{t+1}] = \log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - E_t\left(\log \frac{\Lambda_{t+1}}{\Lambda_t}\right) - \log E_t\left(\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\right) + E_t\left(\log \frac{\Lambda^*_{t+1}}{\Lambda^*_t}\right)$$

Applying the definition of conditional entropy in the equation above yields the Backus, Foresi, and Telmer (2001) expression.

To derive the Alvarez and Jermann (2005) result, first note that since the permanent component of the pricing kernel is a martingale, its conditional entropy can be expressed as follows:

$$L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = -E_t\left(\log \frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right).$$

The definition of conditional entropy implies the following decomposition of total pricing kernel entropy:

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - E_t\left(\log \frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right)$$

or, using the above expression for the conditional entropy of the permanent pricing kernel component:

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = -\log R^f_t - E_t\left(\log \frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\right) + L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right).$$

The Alvarez and Jermann (2005) result hinges on:

$$\lim_{k\to\infty} R^{(k)}_{t+1} = \Lambda^{\mathbb{T}}_t / \Lambda^{\mathbb{T}}_{t+1}.$$

Under the assumption that $0 < \lim_{k\to\infty} \frac{P^{(k)}_t}{\delta^k} < \infty$ for all t, one can write:

$$\lim_{k\to\infty} R^{(k)}_{t+1} = \lim_{k\to\infty} \frac{E_{t+1}\left(\frac{\Lambda_{t+k}}{\Lambda_{t+1}}\right)}{E_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right)} = \frac{\lim_{k\to\infty} \frac{E_{t+1}(\Lambda_{t+k}/\delta^{t+k})}{\Lambda_{t+1}}}{\lim_{k\to\infty} \frac{E_t(\Lambda_{t+k}/\delta^{t+k})}{\Lambda_t}} = \frac{\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda_{t+1}}}{\frac{\Lambda^{\mathbb{P}}_t}{\Lambda_t}} = \Lambda^{\mathbb{T}}_t / \Lambda^{\mathbb{T}}_{t+1}.$$

Thus, the infinite-maturity bond is exposed only to transitory pricing kernel risk. □

- Proof of Proposition 2:

Proof. The transitory component of the exchange rate in levels is given by the spread in long-term yields:

$$s^{\mathbb{T}}_t = s_0 + t(\log\delta - \log\delta^*) + \lim_{k\to\infty} k\left(y^{(k)}_t - y^{(k),*}_t\right) - \lim_{k\to\infty} k\left(y^{(k)}_0 - y^{(k),*}_0\right).$$

This follows directly from the definition of the transitory component of the pricing kernel. Note that $\lim_{k\to\infty} k\left(y^{(k)}_t - y^{(k),*}_t\right)$ is not a constant. If it was, the transitory component of the exchange rate would always be a constant.
This, in turn, implies that the rate of appreciation is given by:

$$\Delta s^{\mathbb{T}}_{t\to t+k} = k(\log\delta - \log\delta^*) + \lim_{k\to\infty} k\left(y^{(k)}_{t+k} - y^{(k),*}_{t+k}\right) - \lim_{k\to\infty} k\left(y^{(k)}_t - y^{(k),*}_t\right).$$

After expressing everything on a per period basis, taking conditional expectations and then taking the limit as the holding period goes to infinity, this expression yields:

$$\lim_{k\to\infty} \frac{1}{k} E_t\left[\Delta s^{\mathbb{T}}_{t\to t+k}\right] = -\lim_{k\to\infty}\left(y^{(k)}_t - y^{(k),*}_t\right),$$

where we have used $\log\delta = -\lim_{k\to\infty} E_t(y_{t+k}^{(k)}) = -\lim_{k\to\infty} E(y_t^{(k)})$; the average long yield in logs is equal to $\log\delta$ under regularity conditions given in Borovička, Hansen, and Scheinkman (2016) on pages 2515 and 2516. □

- Proof of Proposition 3:

Proof. We start from the following equation:

$$\lim_{k\to\infty}\frac{1}{k}E[rx_{t\to t+k}^{FX}] = \lim_{k\to\infty} E\left(y_t^{(k),*} - y_t^{(k)}\right) - \lim_{k\to\infty}\frac{1}{k}E[\Delta s_{t\to t+k}] = \lim_{k\to\infty}(1/k)E\left[L_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+k}^*}{\Lambda_t^*}\right)\right].$$

This follows from the definition of the currency risk premium $E_t[rx_{t\to t+k}] = L_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+k}^*}{\Lambda_t^*}\right)$ in equation (7), extended to longer horizons. Next, we note that $L(x_{t+1}) = EL_t(x_{t+1}) + L(E_t(x_{t+1}))$. Given the stationary of the stochastic discount factor, we know that $\lim_{k\to\infty}(1/k)L\left(E_t\frac{\Lambda_{t+k}}{\Lambda_t}\right) = 0$. Hence $\lim_{k\to\infty}(1/k)L\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) = \lim_{k\to\infty}(1/k)EL_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right)$. It then follows that

$$\lim_{k\to\infty}(1/k)E\left[L_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+k}^*}{\Lambda_t^*}\right)\right] = \lim_{k\to\infty}(1/k)\left[L\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L\left(\frac{\Lambda_{t+k}^*}{\Lambda_t^*}\right)\right] = \left[L\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) - L\left(\frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}\right)\right].$$

The last equality follows directly from the Alvarez-Jermann decomposition of the pricing kernel (see Alvarez and Jermann (2005)'s proposition 6). □

Online Appendix for
"The Term Structure of Currency Carry Trade Risk Premia"
—Not For Publication—

This Online Appendix describes additional empirical and theoretical results on foreign bond returns in U.S. dollars.

- Section A presents robustness checks on the main time-series results reported in the paper:
  - subsection A.1 reports time-series predictability results using standard asymptotic inference;
  - subsection A.2 reports time-series predictability results using inflation and sovereign credit as additional controls;
  - subsection A.3 proposes a different decomposition of the dollar bond returns into its exchange rate component ($-\Delta s_{t+1}$) and the local currency bond return difference, $r^{(10),*} - r^{(10)}$ (instead of excess returns);
  - subsection A.4 reports time-series predictability results with GBP as base currency;
  - subsection A.5 reports additional individual country time-series predictability results obtained on different time-windows (10/1983–12/2007, 1/1975–12/2007, 10/1983–12/2015) and investment horizons (three months).
- Section B presents additional robustness checks for the cross-sectional portfolio results reported in the paper.
  - subsection B.1 reports portfolio statistics for different time-windows (10/1983–12/2007, 1/1975–12/2007, 10/1983—12/2015);
  - subsection B.2 focuses on currency portfolios sorted on the deviation of interest rates from their 10-year rolling means and reports statistics for different sample periods, different holding periods and different sets of currencies;
  - subsection B.3 focuses on currency portfolios sorted on interest rate levels and reports statistics for different sample periods, different holding periods and different sets of currencies;
  - subsection B.4 focuses on currency portfolios sorted on yield curve slopes and reports statistics for different sample periods, different holding periods and different sets of currencies.
- Section C reports additional results obtained with zero-coupon bonds for our benchmark sample of G10 countries and a larger sample of developed countries.
- Section D reports additional theoretical results on dynamic term structure models, starting with the simple Vasicek (1977) model, before turning to their k-factor extensions and the model studied in Lustig, Roussanov, and Verdelhan (2014).
- Section E presents the details of pricing kernel decomposition for three classes of structural models: models with external habit formation, models with long run risks, and models with rare disasters.
- Section F reports additional proofs of preference-free results.
- Section G presents two additional preference-free implications of our findings: a lower bound on the cross-country correlations of the permanent SDF components and a new benchmark for holding bond returns.
- Section H compares finite to infinite maturity bond returns in the benchmark Joslin, Singleton, and Zhu (2011) term structure model.

A Robustness Checks on Time-Series Results

A.1 Time-Series Predictability using Standard Asymptotic Inference

Table A1 refers to the same regressions as Table 1 in the main text, with the difference being that we report Newey and West (1987) standard errors calculated with a kernel bandwidth equal to $S = 6$, the value indicated by the benchmark "textbook" rule $S = 0.75T^{1/3}$, and that we discuss statistical significance using the standard asymptotic distributions. When we use interest rate differentials as the forecasting variable (Panel A), we find no predictability for dollar bond return differentials, with the exception of Japan, consistent with our discussion of Table 1 in the main text. Jointly testing all slope coefficients of individual country regressions, we find marginal significance (at the 5%, but not the 1% level) because of Japan. However, the panel slope coefficient is not statistically significant. Finally, our findings on predictability using yield curve slope differentials (Panel B) are not materially different from those in Table 1.

A.2 Time-Series Predictability with Additional Controls

Table A2 presents additional time-series predictability results when using inflation and sovereign credit rating as additional controls. In particular, we include as regressors the difference (foreign minus domestic) in realized inflation between t and $t+1$ as well as the difference (foreign minus domestic) in the sovereign credit rating at t. These results should be compared to Table 1 in the paper. The slope coefficients are quite similar.

A.3 Time Series Regressions: Exchange Rate Changes and Local Bond Returns (Instead of Excess Returns)

Table A3 proposes a different decomposition of the dollar bond returns into its exchange rate component ($-\Delta s_{t+1}$) and the local currency bond return difference, $r^{(10),*} - r^{(10)}$. When we regress the local currency log return differential (instead of the excess returns) on the interest rate differential, there is no evidence of predictability (Panel A). This decomposition does not suffer from any mechanical link between the right- and left-hand side variables. But its drawback is that it does not show the currency excess return predictability in the middle columns. Instead, it reports the usual U.I.P slope coefficient in a regression of exchange rate changes on the interest rate differential (Panel A). There is of course a simple mapping between those coefficients and those of Table 1 in the paper. A zero slope coefficient in a regression of exchange rate changes on interest rate differences is equivalent to a slope coefficient of one in a regression of currency excess returns on interest rate differences. Table A3 shows that the slope of the yield curve predicts significantly the bond return differential (in local currencies). The predictability results on dollar bond returns are the same as in Table 1 in the paper.

A.4 Time Series Predictability with GBP as Base Currency

Table A4 presents the results obtained when using the GBP as the base currency. We start by considering the interest rate as a predictor. U.I.P. deviations are weaker when the base currency is the GBP. The panel regression coefficient is 1.60 (instead of 1.98). On the other hand, there is less predictability of the local currency bond excess return differential when using the interest rate spread as the predictor. The panel regression coefficient is -0.60 (instead of -1.34). The net effect is a slope coefficient of 1.00, which is significant only at the 10% level. However, when we use the slope of the yield curve as a predictor, the slope coefficient is -2.10 (-2.02 with USD as base currency) for the currency excess return, but 2.53 (3.96 with USD as base currency) for the local currency bond excess return differential. The net effect is a slope coefficient of 0.43, which is not statistically significantly different from zero. To summarize, the slope and interest evidence is qualitatively similar. The slope evidence is entirely in line with our hypothesis. The interest rate evidence suggests there is some predictability left in the dollar bond excess returns.

However, there is no economically significant predictability. In particular, to assess the economic significance of these results, Table A5 presents the results obtained when an investor exploits interest rate and slope predictability by going long U.K. bonds and shorts foreign bonds when the interest rate difference (slope difference) is positive (negative), and reverses the position otherwise. The equally-weighted return on the interest rate strategy in the top panel is only 1.41% per annum, not significant at conventional significance levels. The Sharpe ratio is only 0.22. Similarly, the equally-weighted return on the slope strategy reported in the bottom panel is 0.45% per annum and the annualized Sharpe ratio is 0.07. Thus, there is no evidence of economically significant time variation in GBP bond excess returns, consistent with our hypothesis, in line with (but quantitatively different than) our conclusions for USD bond returns.

Table A1: Dollar Bond Return Differential Predictability

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	p-value	
	Panel A: Short-Term Interest Rates																
Australia	0.01	[0.03]	-0.15	[0.97]	-0.20	-0.02	[0.02]	1.29	[0.62]	0.56	0.03	[0.02]	-1.44	[0.60]	1.51	0.21	492
Canada	0.02	[0.02]	-1.10	[0.69]	0.11	-0.01	[0.01]	1.22	[0.53]	0.46	0.03	[0.01]	-2.32	[0.46]	3.64	0.01	492
Germany	0.01	[0.02]	1.52	[1.21]	0.37	0.02	[0.02]	2.49	[0.99]	1.71	-0.01	[0.01]	-0.97	[0.60]	0.48	0.53	492
Japan	0.06	[0.03]	2.37	[0.84]	1.13	0.07	[0.02]	3.11	[0.67]	3.48	-0.01	[0.02]	-0.74	[0.52]	0.13	0.49	492
New Zealand	-0.03	[0.04]	0.69	[0.87]	-0.03	-0.07	[0.03]	2.23	[0.49]	3.14	0.04	[0.03]	-1.54	[0.66]	1.62	0.12	492
Norway	-0.02	[0.02]	0.72	[0.62]	0.08	-0.02	[0.02]	1.74	[0.57]	2.26	0.01	[0.01]	-1.02	[0.41]	0.97	0.23	492
Sweden	-0.00	[0.02]	-0.64	[0.91]	-0.02	-0.02	[0.02]	0.89	[0.91]	0.25	0.01	[0.01]	-1.53	[0.49]	2.02	0.23	492
Switzerland	0.02	[0.02]	1.16	[0.82]	0.33	0.05	[0.02]	2.45	[0.78]	2.43	-0.03	[0.01]	-1.29	[0.43]	1.69	0.25	492
United Kingdom	-0.02	[0.03]	1.02	[1.18]	0.04	-0.05	[0.02]	2.69	[0.95]	2.44	0.03	[0.02]	-1.67	[0.66]	1.39	0.27	492
Panel	–	–	0.65	[0.49]	-0.05	–	–	1.98	[0.44]	1.82	–	–	-1.34	[0.30]	1.37	0.00	4428
Joint zero (p-value)	0.44		0.04			0.00		0.00			0.00		0.00			0.04	
	Panel B: Yield Curve Slopes																
Australia	0.06	[0.02]	3.84	[1.56]	1.54	0.00	[0.02]	-1.00	[1.17]	-0.02	0.05	[0.02]	4.84	[0.92]	7.65	0.01	492
Canada	0.04	[0.02]	4.04	[0.98]	2.25	-0.00	[0.01]	-0.72	[0.66]	-0.07	0.04	[0.01]	4.76	[0.63]	9.09	0.00	492
Germany	0.00	[0.02]	0.50	[1.77]	-0.18	-0.01	[0.02]	-3.05	[1.37]	1.15	0.01	[0.01]	3.55	[0.97]	4.07	0.11	492
Japan	0.00	[0.02]	-0.32	[1.38]	-0.19	-0.01	[0.02]	-4.18	[1.08]	2.91	0.01	[0.01]	3.85	[0.82]	3.96	0.03	492
New Zealand	0.08	[0.04]	2.94	[2.04]	1.26	-0.01	[0.03]	-1.60	[1.18]	0.62	0.09	[0.03]	4.55	[1.19]	7.41	0.05	492
Norway	-0.00	[0.02]	0.59	[1.03]	-0.12	-0.01	[0.02]	-2.03	[0.92]	1.33	0.01	[0.01]	2.62	[0.59]	3.35	0.06	492
Sweden	0.02	[0.02]	3.12	[1.23]	2.12	-0.00	[0.02]	-0.13	[1.14]	-0.20	0.02	[0.01]	3.25	[0.71]	5.29	0.05	492
Switzerland	0.00	[0.02]	0.97	[1.17]	-0.06	-0.02	[0.02]	-3.59	[1.26]	1.97	0.02	[0.01]	4.55	[0.78]	8.82	0.01	492
United Kingdom	0.02	[0.03]	1.59	[1.53]	0.17	-0.02	[0.02]	-3.17	[1.37]	2.11	0.04	[0.01]	4.75	[0.83]	7.95	0.02	492
Panel	–	–	1.94	[0.84]	0.42	–	–	-2.02	[0.73]	0.83	–	–	3.96	[0.50]	6.08	0.00	4428
Joint zero (p-value)	0.07		0.00			0.96		0.00			0.00		0.00			0.00	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A2: Dollar Bond Return Differential Predictability – Controlling for Inflation and Credit Ratings

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	p-value	
	Panel A: Short-Term Interest Rates																
Australia	-0.02	[0.03]	0.62	[0.97]	1.72	-0.04	[0.02]	1.72	[0.61]	1.37	0.02	[0.03]	-1.10	[0.62]	2.29	0.33	492
Canada	0.02	[0.02]	-1.13	[0.73]	-0.28	-0.01	[0.02]	1.36	[0.59]	0.16	0.03	[0.01]	-2.49	[0.46]	3.54	0.01	492
Germany	0.01	[0.02]	1.81	[1.14]	0.30	0.03	[0.02]	2.75	[1.02]	1.92	-0.01	[0.01]	-0.94	[0.54]	0.38	0.54	492
Japan	0.10	[0.03]	2.96	[0.84]	1.50	0.10	[0.03]	3.37	[0.66]	3.65	0.00	[0.02]	-0.41	[0.54]	0.23	0.70	492
New Zealand	-0.10	[0.05]	1.53	[0.82]	2.36	-0.11	[0.03]	2.53	[0.53]	3.95	0.01	[0.04]	-1.00	[0.65]	2.99	0.31	492
Norway	-0.01	[0.02]	0.74	[0.61]	0.35	-0.01	[0.02]	1.78	[0.56]	3.22	0.00	[0.02]	-1.04	[0.42]	0.64	0.21	492
Sweden	-0.01	[0.02]	-0.63	[0.90]	-0.21	-0.02	[0.02]	0.94	[0.92]	0.09	0.01	[0.01]	-1.56	[0.52]	1.64	0.23	492
Switzerland	0.02	[0.03]	1.69	[0.79]	0.79	0.07	[0.03]	2.85	[0.81]	2.78	-0.05	[0.01]	-1.16	[0.46]	2.93	0.31	492
United Kingdom	-0.02	[0.03]	0.87	[1.22]	-0.30	-0.05	[0.03]	2.83	[1.00]	2.09	0.03	[0.02]	-1.96	[0.65]	1.39	0.21	492
Panel	–	–	0.81	[0.46]	0.24	–	–	2.07	[0.44]	1.98	–	–	-1.26	[0.31]	1.45	0.00	4428
Joint zero (p-value)	0.07		0.00			0.00		0.00			0.00		0.00			0.09	
	Panel B: Yield Curve Slopes																
Australia	0.03	[0.03]	3.00	[1.55]	2.62	-0.02	[0.02]	-1.53	[1.12]	0.54	0.05	[0.02]	4.52	[0.95]	7.85	0.02	492
Canada	0.05	[0.02]	4.80	[1.10]	2.24	-0.00	[0.02]	-0.99	[0.80]	-0.39	0.05	[0.01]	5.80	[0.70]	10.70	0.00	492
Germany	0.00	[0.02]	0.24	[1.74]	-0.46	0.00	[0.02]	-3.47	[1.39]	1.36	0.00	[0.01]	3.71	[0.95]	4.15	0.10	492
Japan	0.02	[0.03]	-0.91	[1.35]	-0.27	0.01	[0.02]	-4.72	[1.08]	3.23	0.01	[0.02]	3.81	[0.87]	3.65	0.03	492
New Zealand	-0.01	[0.06]	2.14	[1.96]	2.34	-0.08	[0.04]	-1.96	[1.16]	1.54	0.07	[0.04]	4.10	[1.19]	7.47	0.07	492
Norway	0.01	[0.02]	0.45	[1.02]	0.11	0.00	[0.02]	-2.20	[0.93]	2.45	0.01	[0.01]	2.65	[0.60]	3.05	0.05	492
Sweden	0.01	[0.02]	3.10	[1.20]	1.81	-0.01	[0.02]	-0.25	[1.13]	-0.38	0.02	[0.01]	3.35	[0.74]	5.00	0.04	492
Switzerland	-0.01	[0.02]	0.51	[1.19]	-0.17	-0.02	[0.02]	-3.97	[1.29]	2.03	0.01	[0.01]	4.48	[0.83]	9.47	0.01	492
United Kingdom	0.02	[0.03]	1.62	[1.53]	-0.07	-0.02	[0.02]	-3.18	[1.40]	1.81	0.04	[0.02]	4.80	[0.84]	7.67	0.02	492
Panel	–	–	1.81	[0.81]	0.54	–	–	-2.10	[0.71]	0.99	–	–	3.91	[0.50]	6.07	0.00	4428
Joint zero (p-value)	0.29		0.00			0.67		0.00			0.00		0.00			0.00	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^{f}$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). In each regression, we also include the realized inflation differential (foreign minus domestic) between t and $t+1$, as well as the credit rating differential (foreign minus domestic) at t as regressors. The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A3: Dollar Bond Return Differential Predictability: Exchange Rate Changes and Local Bond Return Differentials

	Bond dollar return difference $r^{(10),\$} - r^{(10)}$					Exchange rate change $-\Delta s_{t+1}$					Bond local currency return diff. $r^{(10),*} - r^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	p-value	
	Panel A: Short-Term Interest Rates																
Australia	0.01	[0.03]	-0.15	[0.97]	-0.20	-0.02	[0.02]	0.29	[0.62]	-0.16	0.03	[0.02]	-0.44	[0.60]	-0.04	0.70	492
Canada	0.02	[0.02]	-1.10	[0.69]	0.11	-0.01	[0.01]	0.22	[0.53]	-0.18	0.03	[0.01]	-1.32	[0.46]	1.08	0.13	492
Germany	0.01	[0.02]	1.52	[1.21]	0.37	0.02	[0.02]	1.49	[0.99]	0.49	-0.01	[0.01]	0.03	[0.60]	-0.20	0.99	492
Japan	0.06	[0.03]	2.37	[0.84]	1.13	0.07	[0.02]	2.11	[0.67]	1.53	-0.01	[0.02]	0.26	[0.52]	-0.16	0.81	492
New Zealand	-0.03	[0.04]	0.69	[0.87]	-0.03	-0.07	[0.03]	1.23	[0.49]	0.84	0.04	[0.03]	-0.54	[0.66]	0.02	0.59	492
Norway	-0.02	[0.02]	0.72	[0.62]	0.08	-0.02	[0.02]	0.74	[0.57]	0.25	0.01	[0.01]	-0.02	[0.41]	-0.20	0.98	492
Sweden	0.00	[0.02]	-0.64	[0.91]	-0.02	-0.02	[0.02]	-0.11	[0.91]	-0.20	0.01	[0.01]	-0.53	[0.49]	0.07	0.68	492
Switzerland	0.02	[0.02]	1.16	[0.82]	0.33	0.05	[0.02]	1.45	[0.78]	0.73	-0.03	[0.01]	-0.29	[0.43]	-0.11	0.80	492
United Kingdom	-0.02	[0.03]	1.02	[1.18]	0.04	-0.05	[0.02]	1.69	[0.95]	0.86	0.03	[0.02]	-0.67	[0.66]	0.06	0.66	492
Panel	–	–	0.65	[0.49]	-0.05	–	–	0.98	[0.44]	0.45	–	–	-0.34	[0.30]	0.17	0.26	4428
Joint zero (p-value)	0.44		0.04			0.00		0.00			0.00		0.19			0.95	
	Panel B: Yield Curve Slopes																
Australia	0.06	[0.02]	3.84	[1.56]	1.54	-0.01	[0.02]	0.43	[1.14]	-0.17	0.06	[0.02]	3.42	[0.95]	3.77	0.08	492
Canada	0.04	[0.02]	4.04	[0.98]	2.25	-0.00	[0.01]	0.49	[0.66]	-0.14	0.04	[0.01]	3.55	[0.66]	5.12	0.00	492
Germany	0.00	[0.02]	0.50	[1.77]	-0.18	0.00	[0.02]	-1.89	[1.37]	0.32	-0.00	[0.01]	2.39	[1.01]	1.74	0.29	492
Japan	0.00	[0.02]	-0.32	[1.38]	-0.19	0.01	[0.02]	-2.94	[1.07]	1.37	-0.01	[0.01]	2.61	[0.82]	1.72	0.13	492
New Zealand	0.08	[0.04]	2.94	[2.04]	1.26	-0.03	[0.03]	-0.39	[1.09]	-0.15	0.10	[0.03]	3.33	[1.25]	3.95	0.15	492
Norway	-0.00	[0.02]	0.59	[1.03]	-0.12	-0.02	[0.02]	-0.66	[0.91]	-0.04	0.01	[0.01]	1.25	[0.59]	0.61	0.36	492
Sweden	0.02	[0.02]	3.12	[1.23]	2.12	-0.01	[0.02]	1.05	[1.13]	0.15	0.02	[0.01]	2.07	[0.73]	2.07	0.21	492
Switzerland	0.00	[0.02]	0.97	[1.17]	-0.06	0.01	[0.02]	-2.43	[1.28]	0.81	-0.01	[0.01]	3.40	[0.82]	4.92	0.05	492
United Kingdom	0.02	[0.03]	1.59	[1.53]	0.17	-0.03	[0.02]	-2.38	[1.34]	1.12	0.05	[0.01]	3.96	[0.86]	5.53	0.05	492
Panel	–	–	1.94	[0.84]	0.42	–	–	-0.82	[0.72]	0.13	–	–	2.75	[0.52]	3.10	0.00	4428
Joint zero (p-value)	0.07		0.00			0.75		0.03			0.00		0.00			0.00	

Notes: The table reports regression results of the bond dollar return difference ($r_{t+1}^{(10),\$} - r_{t+1}^{(10)}$, left panel) or the exchange rate change ($-\Delta s_{t+1}$, middle panel) or the bond local currency return difference ($r_{t+1}^{(10),*} - r_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A4: Dollar Bond Return Differential Predictability – GBP as base currency

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	p-value	
	Panel A: Short-Term Interest Rates																
Australia	-0.01	[0.02]	1.64	[0.97]	0.41	-0.01	[0.02]	1.89	[0.69]	0.98	-0.00	[0.01]	-0.24	[0.61]	-0.15	0.84	492
Canada	0.02	[0.02]	2.33	[1.17]	0.73	0.02	[0.02]	3.54	[0.95]	2.82	-0.00	[0.01]	-1.20	[0.90]	0.54	0.43	492
Germany	0.03	[0.03]	0.96	[0.89]	0.17	0.02	[0.02]	1.16	[0.63]	0.68	0.00	[0.01]	-0.20	[0.44]	-0.15	0.85	492
Japan	0.08	[0.05]	1.76	[1.01]	0.52	0.08	[0.05]	2.11	[0.90]	1.28	-0.00	[0.01]	-0.34	[0.37]	-0.10	0.80	492
New Zealand	0.00	[0.03]	-0.33	[0.83]	-0.17	-0.01	[0.02]	1.15	[0.52]	0.58	0.01	[0.02]	-1.48	[0.58]	1.65	0.13	492
Norway	-0.00	[0.02]	1.12	[0.67]	0.70	-0.00	[0.01]	1.08	[0.45]	0.99	-0.00	[0.01]	0.04	[0.46]	-0.20	0.96	492
Sweden	-0.02	[0.02]	0.14	[1.05]	-0.20	-0.01	[0.01]	0.93	[0.77]	0.30	-0.01	[0.01]	-0.79	[0.52]	0.55	0.54	492
Switzerland	0.06	[0.03]	1.58	[0.78]	1.22	0.06	[0.03]	1.68	[0.58]	1.87	-0.00	[0.01]	-0.09	[0.33]	-0.17	0.92	492
United States	0.02	[0.03]	1.02	[1.18]	0.04	0.05	[0.02]	2.69	[0.95]	2.44	-0.03	[0.02]	-1.67	[0.66]	1.39	0.27	492
Panel	–	–	1.00	[0.54]	0.15	–	–	1.60	[0.37]	1.16	–	–	-0.60	[0.31]	0.34	0.05	4428
Joint zero (p-value)	0.41		0.03			0.06		0.00			0.82		0.03			0.86	
	Panel B: Yield Curve Slopes																
Australia	0.00	[0.02]	0.20	[1.51]	-0.20	-0.00	[0.02]	-2.48	[1.01]	0.72	0.01	[0.01]	2.68	[1.02]	2.71	0.14	492
Canada	-0.00	[0.02]	0.50	[1.47]	-0.17	-0.00	[0.02]	-2.85	[1.37]	1.53	0.00	[0.01]	3.35	[0.79]	4.91	0.10	492
Germany	0.01	[0.02]	-1.48	[1.24]	0.18	0.01	[0.02]	-2.45	[0.94]	1.48	0.00	[0.01]	0.96	[0.62]	0.37	0.54	492
Japan	0.02	[0.03]	-2.24	[1.47]	0.39	0.01	[0.02]	-3.53	[1.19]	1.93	0.00	[0.01]	1.29	[0.62]	0.57	0.50	492
New Zealand	0.05	[0.03]	3.46	[1.09]	2.75	0.01	[0.02]	-0.67	[0.59]	-0.02	0.05	[0.01]	4.13	[0.75]	9.88	0.00	492
Norway	-0.01	[0.02]	-0.71	[0.99]	-0.03	-0.00	[0.01]	-1.54	[0.66]	0.96	-0.00	[0.01]	0.82	[0.63]	0.49	0.49	492
Sweden	-0.02	[0.02]	0.93	[1.38]	0.03	-0.01	[0.01]	-1.25	[1.01]	0.42	-0.01	[0.01]	2.18	[0.65]	3.69	0.20	492
Switzerland	0.00	[0.02]	-2.73	[1.19]	1.04	0.00	[0.02]	-3.92	[0.91]	3.13	-0.00	[0.01]	1.19	[0.55]	1.24	0.43	492
United States	-0.02	[0.03]	1.59	[1.53]	0.17	0.02	[0.02]	-3.17	[1.37]	2.11	-0.04	[0.01]	4.75	[0.83]	7.95	0.02	492
Panel	–	–	0.43	[0.82]	-0.15	–	–	-2.10	[0.56]	1.06	–	–	2.53	[0.44]	3.67	0.00	4428
Joint zero (p-value)	0.67		0.01			0.98		0.00			0.03		0.00			0.00	

Notes: The table reports regression results of the bond British pound return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.K. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.K. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond pound return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.S. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A5: Dynamic Long-Short Foreign and U.S. Bond Portfolios – GBP as Base Currency

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Short-Term Interest Rates														
Australia	3.81	[2.31]	14.91	0.26	[0.16]	3.70	[1.91]	12.36	0.30	[0.16]	0.11	[1.19]	7.54	0.01	[0.16]
Canada	2.03	[1.89]	12.44	0.16	[0.16]	2.96	[1.62]	10.44	0.28	[0.16]	-0.94	[1.07]	7.17	-0.13	[0.16]
Germany	0.12	[1.76]	11.28	0.01	[0.15]	1.50	[1.36]	8.91	0.17	[0.15]	-1.38	[0.93]	5.99	-0.23	[0.15]
Japan	0.19	[2.22]	14.65	0.01	[0.16]	1.18	[1.89]	12.19	0.10	[0.16]	-0.99	[1.16]	7.40	-0.13	[0.17]
New Zealand	0.14	[2.40]	15.66	0.01	[0.15]	3.02	[1.82]	12.07	0.25	[0.16]	-2.88	[1.56]	10.10	-0.29	[0.15]
Norway	3.48	[1.65]	10.52	0.33	[0.15]	2.80	[1.38]	8.80	0.32	[0.16]	0.68	[0.93]	6.09	0.11	[0.15]
Sweden	1.30	[1.80]	11.41	0.11	[0.16]	2.52	[1.49]	9.33	0.27	[0.16]	-1.22	[0.99]	6.51	-0.19	[0.16]
Switzerland	0.73	[1.81]	11.66	0.06	[0.16]	0.72	[1.57]	10.24	0.07	[0.16]	0.01	[0.73]	4.73	0.00	[0.16]
United States	0.89	[2.00]	12.76	0.07	[0.15]	3.09	[1.59]	10.26	0.30	[0.16]	-2.20	[1.25]	8.21	-0.27	[0.15]
Equally-weighted	1.41	[0.98]	6.33	0.22	[0.16]	2.39	[0.80]	5.11	0.47	[0.17]	-0.98	[0.52]	3.41	-0.29	[0.16]
	Panel B: Yield Curve Slopes														
Australia	0.41	[2.34]	14.95	0.03	[0.16]	3.48	[1.94]	12.37	0.28	[0.17]	-3.07	[1.20]	7.48	-0.41	[0.15]
Canada	-2.42	[1.95]	12.44	-0.19	[0.16]	1.56	[1.70]	10.47	0.15	[0.16]	-3.98	[1.10]	7.08	-0.56	[0.15]
Germany	1.38	[1.72]	11.27	0.12	[0.16]	2.88	[1.37]	8.88	0.32	[0.16]	-1.50	[0.95]	5.99	-0.25	[0.15]
Japan	3.04	[2.39]	14.62	0.21	[0.16]	3.57	[1.96]	12.15	0.29	[0.16]	-0.53	[1.17]	7.40	-0.07	[0.16]
New Zealand	-2.96	[2.34]	15.64	-0.19	[0.15]	2.43	[1.85]	12.08	0.20	[0.16]	-5.39	[1.60]	10.02	-0.54	[0.13]
Norway	0.89	[1.67]	10.57	0.08	[0.16]	2.72	[1.43]	8.80	0.31	[0.16]	-1.83	[0.94]	6.07	-0.30	[0.15]
Sweden	1.53	[1.80]	11.41	0.13	[0.15]	3.66	[1.48]	9.30	0.39	[0.16]	-2.13	[1.00]	6.49	-0.33	[0.16]
Switzerland	4.88	[1.79]	11.58	0.42	[0.15]	6.38	[1.57]	10.08	0.63	[0.15]	-1.50	[0.70]	4.71	-0.32	[0.16]
United States	-2.73	[1.95]	12.73	-0.21	[0.16]	2.06	[1.61]	10.29	0.20	[0.16]	-4.79	[1.30]	8.12	-0.59	[0.16]
Equally-weighted	0.45	[1.05]	6.59	0.07	[0.16]	3.19	[0.92]	5.65	0.57	[0.16]	-2.75	[0.55]	3.45	-0.80	[0.15]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the British bond when the foreign short-term interest rate is higher than the U.K. interest rate (or the foreign yield curve slope is lower than the U.K. yield curve slope), and go long the British bond and short the foreign country bond when the U.K. interest rate is higher than the country's interest rate (or the U.K. yield curve slope is lower than the foreign yield curve slope). Results based on interest rate levels are reported in Panel A and results based on interest rate slopes are reported in Panel B. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.K. pounds ($rx^{(10),\$} - rx^{(10)}$, left panel). The holding period is one month. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 1/1975–12/2015.

A.5 Individual Country Time-Series Predictability Results

Table A6 and A7 report the time-series regression results when we end the sample in 2007: Table A6 considers the shorter 10/1983-12/2007 sample period, whereas Table A7 considers the 1/1975-12/2007 sample period. The first column looks at dollar return differential predictability. The panel slope coefficient for the interest rate regressions is 1.05 in the short sample, compared to 0.65 in the full sample, and we find marginal evidence in favor of interest rate predictability of the dollar return differential, driven mainly by Japan. The R^2s in these regressions are extremely low. However, the evidence for yield curve slope predictability is weaker in this shorter sample; the panel slope coefficient of 0.58 is no longer statistically different from zero. When we look at the 1/1975-12/2007 sample, the panel slope coefficient for the interest rate regression is 0.86 and not statistically significant, while the panel slope coefficient for the slope regression is 1.54, marginally statistically significant. As happens for our benchmark sample period, the latter coefficient for this sample period also has the opposite sign from what the standard slope carry trade would imply. Finally, Table A8 reports the predictability regression results for the sample period 10/1983-12/2105. We find that the slope coefficient in the interest rate predictability panel regression is 0.81 and non-significant, whereas the slope coefficient in the yield curve slope predictability panel regression is 1.26 and also not statistically significant.

Tables A9, A10 and A11 explore whether there is economically significant evidence of return predictability. Note that, as regards the first two of those tables, leaving out the recent financial crisis would *have to influence average returns* if one believes that carry trade returns compensate investors for taking on non-diversifiable risk (see Lustig and Verdelhan, 2007, for an early version of this perspective). In the shorter 10/1983-12/2007 sample (Table A9), the equally-weighted dollar return on the dynamic strategy that exploits interest rate predictability is 2.57% per annum, with a standard error of 1.17%. Not surprisingly, this increase in the dollar return is due to a higher currency excess return of 3.89% per annum in the sample that leaves out the crisis; the currency excess return only 2.59% in the full sample. That difference largely explains why this strategy produces statistically significant returns in the shorter sample. On the other hand, the equally-weighted dollar return on the dynamic strategy that exploits slope predictability is 1.53% per annum, with a standard error of 1.58%. In the longer 1/1975-12/2007 sample (Table A10), the equally-weighted dollar return on the dynamic strategy that exploits interest rate predictability is 1.38% per annum, with a standard error of 1.02%. Thus, in this longer sample, the dollar return differential is no longer significant. The equally-weighted dollar return on the dynamic strategy that exploits slope predictability is -0.65% per annum, also not significant, as its standard error is 1.30%. To summarize, the main difference seems to be an increase in carry trade returns if we exclude the financial crisis. Finally, in the 10/1983-12/2015 sample period (Table A11), neither the interest rate nor the yield curve slope equally-weighted strategy yields statistically significant dollar bond returns: the former has an average annualized return of 1.42% with a standard error of 1.16% and the latter has an average annualized return of 0.51% with a standard error of 1.47%. Overall, our main findings continue to hold.

Finally, we check the robustness of our time-series predictability results by considering a horizon of three months. Tables A12 and A13 report the output of three-month return predictability regressions for bond and currency excess returns over our benchmark sample period (1/1975–12/2015), for both coupon bonds (balanced sample) and zero-coupon bonds (unbalanced sample). As we can see, while we find no statistical evidence or USD bond return predictability using slopes, there is some evidence for interest rate predictability, both using coupon bonds and zero-coupon bonds. However, as seen in Tables A14 and A15 that evaluate the economic significant of interest rate and slope predictability, neither interest rate- nor slope-based portfolio strategies can achieve statistically significant USD bond returns, in line with our hypothesis.

Table A6: Dollar Bond Return Differential Predictability (10/1983 - 12/2007 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	p-value	
	Panel A: Short-Term Interest Rates																
Australia	-0.00	[0.04]	0.71	[1.16]	-0.17	-0.02	[0.03]	1.71	[0.71]	1.48	0.02	[0.02]	-1.00	[0.67]	0.85	0.46	291
Canada	0.03	[0.02]	-0.89	[0.74]	-0.04	0.01	[0.02]	1.05	[0.55]	0.50	0.02	[0.01]	-1.93	[0.48]	3.16	0.04	291
Germany	0.01	[0.03]	1.06	[1.42]	-0.05	0.03	[0.02]	2.07	[1.26]	1.08	-0.02	[0.01]	-1.01	[0.84]	0.45	0.60	291
Japan	0.09	[0.04]	3.58	[1.16]	1.64	0.12	[0.04]	4.27	[1.10]	4.15	-0.03	[0.02]	-0.69	[0.59]	-0.14	0.66	291
New Zealand	-0.05	[0.05]	1.32	[0.85]	0.47	-0.06	[0.03]	2.27	[0.54]	4.70	0.01	[0.03]	-0.96	[0.65]	0.68	0.34	291
Norway	-0.01	[0.03]	1.15	[0.84]	0.41	-0.00	[0.02]	1.50	[0.77]	1.49	-0.01	[0.02]	-0.34	[0.54]	-0.20	0.76	291
Sweden	0.02	[0.03]	-0.06	[0.99]	-0.34	0.00	[0.03]	1.20	[1.10]	0.79	0.02	[0.02]	-1.26	[0.52]	2.04	0.40	291
Switzerland	0.02	[0.03]	2.06	[1.16]	0.92	0.06	[0.04]	2.88	[1.23]	2.48	-0.04	[0.02]	-0.82	[0.70]	0.29	0.63	291
United Kingdom	-0.02	[0.03]	1.07	[1.37]	-0.08	-0.03	[0.03]	2.69	[1.23]	1.96	0.01	[0.02]	-1.61	[0.65]	1.48	0.38	291
Panel	–	–	1.05	[0.61]	0.14	–	–	2.03	[0.56]	2.14	–	–	-0.98	[0.36]	0.73	0.01	2619
Joint zero (p-value)	0.29		0.02			0.02		0.00			0.05		0.00			0.52	
	Panel B: Yield Curve Slopes																
Australia	0.04	[0.03]	1.61	[1.91]	-0.03	0.00	[0.02]	-2.11	[1.30]	0.63	0.04	[0.02]	3.72	[0.99]	5.44	0.11	291
Canada	0.04	[0.02]	2.95	[1.08]	1.54	0.02	[0.01]	-0.95	[0.73]	0.04	0.02	[0.01]	3.90	[0.66]	7.49	0.00	291
Germany	0.00	[0.02]	-0.07	[1.97]	-0.35	0.01	[0.02]	-3.22	[1.77]	1.19	-0.01	[0.01]	3.16	[1.23]	3.07	0.23	291
Japan	-0.01	[0.03]	-2.42	[1.71]	0.09	-0.02	[0.02]	-6.05	[1.56]	3.96	0.01	[0.02]	3.63	[0.96]	2.28	0.12	291
New Zealand	0.04	[0.05]	0.84	[2.66]	-0.22	-0.01	[0.04]	-2.55	[1.54]	2.07	0.06	[0.03]	3.39	[1.52]	4.56	0.27	291
Norway	0.01	[0.03]	-0.47	[1.35]	-0.30	0.01	[0.02]	-1.86	[1.37]	0.67	-0.00	[0.02]	1.39	[0.85]	0.49	0.47	291
Sweden	0.04	[0.03]	1.70	[1.30]	0.54	0.02	[0.02]	-1.09	[1.55]	0.09	0.02	[0.02]	2.79	[0.81]	5.11	0.17	291
Switzerland	-0.02	[0.02]	-0.41	[1.35]	-0.32	-0.02	[0.02]	-3.45	[1.58]	1.83	-0.00	[0.01]	3.04	[0.78]	4.36	0.14	291
United Kingdom	0.01	[0.03]	0.33	[1.70]	-0.33	-0.02	[0.03]	-3.41	[1.62]	1.49	0.03	[0.02]	3.74	[0.98]	4.50	0.11	291
Panel	–	–	0.58	[1.01]	-0.22	–	–	-2.49	[0.91]	1.30	–	–	3.08	[0.59]	3.76	0.00	2619
Joint zero (p-value)	0.29		0.20			0.87		0.00			0.03		0.00			0.00	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 10/1983–12/2007. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 5. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 5.

Table A7: Dollar Bond Return Differential Predictability (1/1975 - 12/2007 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	p-value	
	Panel A: Short-Term Interest Rates																
Australia	0.01	[0.03]	-0.14	[1.01]	-0.25	-0.02	[0.02]	1.42	[0.57]	1.07	0.03	[0.02]	-1.56	[0.63]	1.81	0.18	396
Canada	0.03	[0.02]	-1.09	[0.68]	0.16	-0.00	[0.01]	1.24	[0.50]	0.96	0.03	[0.01]	-2.32	[0.47]	3.86	0.01	396
Germany	0.02	[0.02]	1.80	[1.29]	0.62	0.04	[0.02]	2.89	[1.04]	2.77	-0.01	[0.01]	-1.08	[0.64]	0.60	0.51	396
Japan	0.11	[0.03]	3.45	[0.93]	2.21	0.12	[0.03]	4.07	[0.72]	5.55	-0.01	[0.02]	-0.62	[0.59]	-0.05	0.60	396
New Zealand	-0.05	[0.05]	1.00	[0.86]	0.15	-0.09	[0.03]	2.55	[0.46]	5.78	0.04	[0.03]	-1.54	[0.68]	1.60	0.11	396
Norway	-0.01	[0.02]	0.95	[0.60]	0.37	-0.01	[0.02]	1.92	[0.56]	3.80	0.01	[0.02]	-0.97	[0.42]	0.94	0.23	396
Sweden	0.00	[0.02]	-0.47	[0.93]	-0.14	-0.02	[0.02]	1.10	[0.92]	0.65	0.02	[0.01]	-1.57	[0.50]	2.15	0.23	396
Switzerland	0.03	[0.03]	1.28	[0.88]	0.42	0.07	[0.03]	2.84	[0.88]	3.48	-0.05	[0.02]	-1.55	[0.47]	2.34	0.21	396
United Kingdom	-0.01	[0.03]	1.15	[1.27]	0.07	-0.06	[0.03]	3.09	[1.03]	3.38	0.04	[0.02]	-1.95	[0.71]	1.79	0.23	396
Panel	–	–	0.86	[0.51]	0.06	–	–	2.26	[0.46]	2.96	–	–	-1.40	[0.33]	1.53	0.00	3564
Joint zero (p-value)	0.05		0.00			0.00		0.00			0.00		0.00			0.03	
	Panel B: Yield Curve Slopes																
Australia	0.05	[0.02]	3.87	[1.67]	1.67	-0.00	[0.02]	-1.56	[1.01]	0.36	0.05	[0.02]	5.43	[1.00]	9.37	0.01	396
Canada	0.04	[0.01]	3.44	[0.96]	2.17	0.00	[0.01]	-1.29	[0.61]	0.51	0.04	[0.01]	4.72	[0.66]	9.61	0.00	396
Germany	0.01	[0.02]	0.11	[1.90]	-0.25	0.01	[0.02]	-3.68	[1.47]	2.04	-0.00	[0.01]	3.80	[1.06]	4.62	0.11	396
Japan	0.01	[0.03]	-0.87	[1.54]	-0.18	0.01	[0.02]	-5.04	[1.18]	3.95	0.01	[0.02]	4.17	[0.95]	4.06	0.03	396
New Zealand	0.07	[0.05]	2.54	[2.13]	0.92	-0.04	[0.03]	-2.29	[1.18]	1.94	0.11	[0.03]	4.83	[1.28]	7.94	0.05	396
Norway	0.00	[0.02]	-0.23	[0.91]	-0.24	0.00	[0.02]	-2.76	[0.84]	3.44	0.00	[0.02]	2.53	[0.63]	3.31	0.04	396
Sweden	0.01	[0.02]	2.62	[1.25]	1.70	0.00	[0.02]	-0.67	[1.14]	-0.07	0.01	[0.01]	3.29	[0.73]	5.61	0.05	396
Switzerland	-0.01	[0.02]	0.88	[1.23]	-0.12	-0.02	[0.02]	-3.89	[1.32]	2.73	0.01	[0.01]	4.77	[0.85]	10.15	0.01	396
United Kingdom	0.03	[0.03]	1.37	[1.57]	0.07	-0.02	[0.03]	-3.54	[1.41]	3.05	0.05	[0.02]	4.90	[0.87]	8.75	0.02	396
Panel	–	–	1.54	[0.86]	0.21	–	–	-2.58	[0.72]	1.77	–	–	4.12	[0.54]	6.65	0.00	3564
Joint zero (p-value)	0.10		0.00			0.97		0.00			0.00		0.00			0.00	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2007. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A8: Dollar Bond Return Differential Predictability (10/1983 - 12/2015 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	p-value	
	Panel A: Short-Term Interest Rates																
Australia	-0.01	[0.03]	0.63	[1.10]	-0.15	-0.02	[0.03]	1.54	[0.74]	0.67	0.02	[0.02]	-0.91	[0.64]	0.63	0.49	387
Canada	0.02	[0.02]	-1.03	[0.76]	0.01	-0.00	[0.02]	0.98	[0.58]	0.10	0.02	[0.01]	-2.01	[0.48]	2.99	0.04	387
Germany	0.00	[0.02]	0.87	[1.34]	-0.09	0.01	[0.02]	1.74	[1.18]	0.53	-0.01	[0.01]	-0.87	[0.72]	0.29	0.62	387
Japan	0.03	[0.03]	2.13	[1.01]	0.62	0.06	[0.03]	2.80	[0.94]	2.01	-0.02	[0.02]	-0.68	[0.51]	-0.03	0.62	387
New Zealand	-0.03	[0.04]	0.99	[0.86]	0.13	-0.04	[0.03]	1.98	[0.54]	2.34	0.02	[0.03]	-0.99	[0.66]	0.79	0.33	387
Norway	-0.02	[0.03]	0.97	[0.84]	0.13	-0.02	[0.02]	1.41	[0.77]	0.86	-0.01	[0.02]	-0.44	[0.51]	-0.06	0.70	387
Sweden	0.01	[0.02]	-0.19	[0.95]	-0.24	-0.00	[0.02]	1.01	[1.05]	0.32	0.01	[0.01]	-1.20	[0.49]	1.76	0.40	387
Switzerland	0.01	[0.03]	1.89	[1.06]	0.67	0.04	[0.03]	2.46	[1.05]	1.47	-0.03	[0.01]	-0.58	[0.63]	0.04	0.70	387
United Kingdom	-0.02	[0.03]	0.82	[1.26]	-0.11	-0.03	[0.02]	2.29	[1.12]	1.36	0.01	[0.01]	-1.47	[0.56]	1.32	0.38	387
Panel	–	–	0.81	[0.58]	-0.00	–	–	1.76	[0.52]	1.16	–	–	-0.94	[0.33]	0.65	0.00	3483
Joint zero (p-value)	0.87		0.15			0.24		0.00			0.04		0.00			0.53	
	Panel B: Yield Curve Slopes																
Australia	0.05	[0.03]	2.08	[1.80]	0.20	0.01	[0.02]	-1.23	[1.49]	-0.04	0.04	[0.02]	3.31	[0.95]	4.11	0.16	387
Canada	0.03	[0.02]	3.88	[1.16]	1.90	0.01	[0.01]	-0.09	[0.82]	-0.26	0.03	[0.01]	3.97	[0.64]	6.96	0.01	387
Germany	-0.00	[0.02]	0.46	[1.80]	-0.24	-0.00	[0.02]	-2.48	[1.55]	0.47	0.00	[0.01]	2.93	[1.00]	2.57	0.22	387
Japan	-0.02	[0.03]	-1.62	[1.56]	-0.04	-0.03	[0.02]	-5.07	[1.36]	2.92	0.01	[0.01]	3.45	[0.88]	2.33	0.10	387
New Zealand	0.05	[0.05]	1.53	[2.57]	0.12	0.00	[0.03]	-1.75	[1.56]	0.57	0.05	[0.02]	3.28	[1.43]	4.40	0.28	387
Norway	0.01	[0.03]	0.91	[1.53]	-0.14	0.01	[0.02]	-0.85	[1.46]	-0.11	0.00	[0.02]	1.76	[0.80]	0.88	0.41	387
Sweden	0.04	[0.02]	2.51	[1.28]	1.19	0.01	[0.02]	-0.28	[1.50]	-0.24	0.02	[0.01]	2.79	[0.75]	4.66	0.16	387
Switzerland	-0.01	[0.02]	-0.10	[1.27]	-0.26	-0.02	[0.02]	-3.07	[1.42]	1.06	0.01	[0.01]	2.97	[0.71]	3.61	0.12	387
United Kingdom	0.00	[0.03]	0.61	[1.64]	-0.22	-0.02	[0.02]	-2.82	[1.56]	0.93	0.02	[0.02]	3.44	[0.83]	3.90	0.13	387
Panel	–	–	1.26	[1.01]	-0.00	–	–	-1.77	[0.92]	0.41	–	–	3.03	[0.54]	3.46	0.00	3483
Joint zero (p-value)	0.26		0.03			0.87		0.00			0.00		0.00			0.01	

Notes: The table reports regression results of the bond dollar return difference ($rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$, left panel) or the currency excess return (rx_{t+1}^{FX}, middle panel) or the bond local currency return difference ($rx_{t+1}^{(10),*} - rx_{t+1}^{(10)}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r_t^{f,*} - r_t^f$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y_t^{(10,*)} - y_t^{(1,*)}] - [y_t^{(10)} - y_t^{(1)}]$, Panel B). The column "Slope Diff." presents the p-value of the test for equality between the slope coefficient in the bond dollar return difference regression and the slope coefficient in the currency excess return regression for each country. The last line in each panel presents the p-value of the joint test that all individual-country regression coefficients in the respective column are zero. We use returns on 10-year coupon bonds. The holding period is one month and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 10/1983–12/2015. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A9: Dynamic Long-Short Foreign and U.S. Bond Portfolios (10/1983 - 12/2007 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Short-Term Interest Rates														
Australia	3.95	[2.86]	14.25	0.28	[0.21]	5.41	[2.18]	10.62	0.51	[0.22]	-1.46	[1.52]	7.75	-0.19	[0.20]
Canada	0.76	[1.64]	8.18	0.09	[0.21]	2.13	[1.16]	5.83	0.36	[0.21]	-1.37	[1.08]	5.28	-0.26	[0.20]
Germany	2.16	[2.41]	12.18	0.18	[0.21]	3.20	[2.13]	10.87	0.29	[0.21]	-1.04	[1.42]	7.17	-0.15	[0.21]
Japan	2.02	[2.82]	14.49	0.14	[0.20]	1.88	[2.23]	11.51	0.16	[0.20]	0.14	[1.73]	8.86	0.02	[0.20]
New Zealand	2.72	[3.46]	17.23	0.16	[0.21]	6.45	[2.43]	11.98	0.54	[0.23]	-3.73	[2.24]	11.22	-0.33	[0.19]
Norway	3.70	[2.48]	12.36	0.30	[0.22]	5.11	[2.06]	10.29	0.50	[0.22]	-1.41	[1.78]	8.52	-0.17	[0.20]
Sweden	4.22	[2.30]	11.61	0.36	[0.21]	5.68	[2.11]	10.57	0.54	[0.23]	-1.46	[1.59]	7.69	-0.19	[0.20]
Switzerland	2.20	[2.44]	12.42	0.18	[0.20]	1.00	[2.28]	11.62	0.09	[0.20]	1.20	[1.41]	6.99	0.17	[0.20]
United Kingdom	1.43	[2.50]	12.12	0.12	[0.21]	4.19	[2.08]	10.33	0.41	[0.21]	-2.76	[1.45]	6.99	-0.39	[0.21]
Equally-weighted	2.57	[1.17]	5.63	0.46	[0.22]	3.89	[0.95]	4.69	0.83	[0.24]	-1.32	[0.73]	3.56	-0.37	[0.21]
	Panel B: Yield Curve Slopes														
Australia	2.00	[2.95]	14.29	0.14	[0.21]	4.92	[2.22]	10.64	0.46	[0.21]	-2.92	[1.56]	7.71	-0.38	[0.20]
Canada	-1.16	[1.72]	8.18	-0.14	[0.21]	2.35	[1.18]	5.82	0.40	[0.21]	-3.51	[1.08]	5.20	-0.68	[0.21]
Germany	3.46	[2.28]	12.15	0.28	[0.21]	6.64	[2.07]	10.74	0.62	[0.21]	-3.18	[1.41]	7.11	-0.45	[0.20]
Japan	2.93	[2.83]	14.48	0.20	[0.21]	6.65	[2.17]	11.36	0.59	[0.22]	-3.72	[1.81]	8.80	-0.42	[0.21]
New Zealand	2.53	[3.63]	17.23	0.15	[0.21]	6.22	[2.51]	11.99	0.52	[0.23]	-3.69	[2.30]	11.22	-0.33	[0.19]
Norway	1.06	[2.58]	12.40	0.09	[0.20]	3.19	[2.07]	10.35	0.31	[0.21]	-2.14	[1.75]	8.51	-0.25	[0.20]
Sweden	0.77	[2.42]	11.67	0.07	[0.20]	4.44	[2.13]	10.62	0.42	[0.21]	-3.67	[1.52]	7.63	-0.48	[0.20]
Switzerland	2.42	[2.61]	12.42	0.19	[0.20]	5.11	[2.29]	11.53	0.44	[0.21]	-2.69	[1.47]	6.95	-0.39	[0.20]
United Kingdom	-0.20	[2.45]	12.13	-0.02	[0.20]	2.63	[2.10]	10.37	0.25	[0.21]	-2.82	[1.42]	6.99	-0.40	[0.20]
Equally-weighted	1.53	[1.58]	7.54	0.20	[0.20]	4.68	[1.23]	6.11	0.77	[0.22]	-3.15	[1.06]	5.18	-0.61	[0.21]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate (or the foreign yield curve slope is lower than the U.S. yield curve slope), and go long the U.S. bond and short the foreign country bond when the U.S. interest rate is higher than the country's interest rate (or the U.S. yield curve slope is lower than the foreign yield curve slope). Results based on interest rate levels are reported in Panel A and results based on interest rate slopes are reported in Panel B. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). The holding period is one month. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 10/1983–12/2007.

Table A10: Dynamic Long-Short Foreign and U.S. Bond Portfolios (1/1975 - 12/2007 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Short-Term Interest Rates														
Australia	1.39	[2.61]	14.44	0.10	[0.18]	4.06	[1.83]	10.24	0.40	[0.18]	-2.67	[1.54]	9.03	-0.30	[0.17]
Canada	0.02	[1.48]	8.48	0.00	[0.17]	1.61	[1.00]	5.65	0.29	[0.17]	-1.60	[1.00]	5.75	-0.28	[0.18]
Germany	2.12	[2.21]	12.82	0.17	[0.17]	3.99	[1.86]	10.95	0.36	[0.17]	-1.87	[1.37]	7.76	-0.24	[0.18]
Japan	1.48	[2.68]	15.13	0.10	[0.17]	2.20	[2.06]	11.60	0.19	[0.18]	-0.72	[1.69]	9.47	-0.08	[0.17]
New Zealand	0.46	[3.04]	17.20	0.03	[0.18]	4.30	[1.99]	11.29	0.38	[0.19]	-3.84	[2.13]	12.35	-0.31	[0.17]
Norway	2.31	[2.24]	12.68	0.18	[0.18]	4.97	[1.73]	9.97	0.50	[0.18]	-2.66	[1.66]	9.38	-0.28	[0.17]
Sweden	1.10	[2.27]	12.84	0.09	[0.18]	4.19	[1.82]	10.59	0.40	[0.19]	-3.10	[1.66]	9.28	-0.33	[0.17]
Switzerland	1.34	[2.21]	12.94	0.10	[0.17]	1.62	[2.05]	12.14	0.13	[0.17]	-0.28	[1.39]	7.98	-0.03	[0.17]
United Kingdom	2.19	[2.28]	13.00	0.17	[0.18]	4.57	[1.80]	10.40	0.44	[0.18]	-2.38	[1.55]	8.77	-0.27	[0.18]
Equally-weighted	1.38	[1.02]	5.65	0.24	[0.18]	3.50	[0.81]	4.55	0.77	[0.19]	-2.12	[0.64]	3.66	-0.58	[0.18]
	Panel B: Yield Curve Slopes														
Australia	-2.52	[2.53]	14.43	-0.17	[0.17]	3.24	[1.81]	10.27	0.32	[0.19]	-5.77	[1.59]	8.91	-0.65	[0.16]
Canada	-1.99	[1.51]	8.46	-0.23	[0.17]	2.07	[1.00]	5.64	0.37	[0.17]	-4.06	[0.95]	5.65	-0.72	[0.18]
Germany	2.80	[2.19]	12.80	0.22	[0.18]	6.94	[1.88]	10.83	0.64	[0.18]	-4.14	[1.32]	7.69	-0.54	[0.17]
Japan	-0.13	[2.79]	15.14	-0.01	[0.17]	5.95	[2.09]	11.49	0.52	[0.18]	-6.07	[1.63]	9.31	-0.65	[0.19]
New Zealand	-0.52	[3.08]	17.20	-0.03	[0.17]	3.85	[2.02]	11.30	0.34	[0.19]	-4.37	[2.22]	12.34	-0.35	[0.17]
Norway	0.76	[2.14]	12.69	0.06	[0.17]	4.56	[1.72]	9.99	0.46	[0.18]	-3.80	[1.60]	9.34	-0.41	[0.17]
Sweden	-2.98	[2.23]	12.82	-0.23	[0.17]	2.60	[1.84]	10.63	0.24	[0.18]	-5.58	[1.62]	9.18	-0.61	[0.17]
Switzerland	0.23	[2.17]	12.94	0.02	[0.17]	5.54	[2.08]	12.04	0.46	[0.18]	-5.30	[1.32]	7.83	-0.68	[0.17]
United Kingdom	-1.50	[2.29]	13.00	-0.12	[0.17]	3.68	[1.81]	10.42	0.35	[0.17]	-5.18	[1.52]	8.66	-0.60	[0.17]
Equally-weighted	-0.65	[1.30]	7.28	-0.09	[0.18]	4.27	[1.04]	5.79	0.74	[0.18]	-4.92	[0.89]	5.12	-0.96	[0.18]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate (or the foreign yield curve slope is lower than the U.S. yield curve slope), and go long the U.S. bond and short the foreign country bond when the U.S. interest rate is higher than the country's interest rate (or the U.S. yield curve slope is lower than the foreign yield curve slope). Results based on interest rate levels are reported in Panel A and results based on interest rate slopes are reported in Panel B. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). The holding period is one month. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 1/1975–12/2007.

Table A11: Dynamic Long-Short Foreign and U.S. Bond Portfolios (10/1983 - 12/2015 Sample Period)

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Short-Term Interest Rates														
Australia	3.17	[2.52]	14.09	0.23	[0.18]	4.42	[2.11]	11.95	0.37	[0.19]	-1.25	[1.30]	7.26	-0.17	[0.17]
Canada	-0.03	[1.60]	9.07	-0.00	[0.18]	1.04	[1.25]	7.40	0.14	[0.18]	-1.07	[0.88]	5.06	-0.21	[0.17]
Germany	2.24	[2.06]	11.86	0.19	[0.18]	3.23	[1.92]	11.12	0.29	[0.18]	-0.99	[1.21]	6.69	-0.15	[0.17]
Japan	1.19	[2.40]	13.57	0.09	[0.17]	1.11	[1.94]	11.16	0.10	[0.17]	0.07	[1.49]	8.42	0.01	[0.17]
New Zealand	2.40	[3.02]	16.85	0.14	[0.18]	5.33	[2.20]	12.98	0.41	[0.19]	-2.93	[1.89]	10.27	-0.29	[0.17]
Norway	1.29	[2.29]	12.87	0.10	[0.18]	2.79	[1.88]	10.99	0.25	[0.18]	-1.49	[1.48]	8.21	-0.18	[0.18]
Sweden	1.54	[2.16]	11.97	0.13	[0.18]	2.91	[1.94]	11.29	0.26	[0.18]	-1.37	[1.30]	7.22	-0.19	[0.17]
Switzerland	1.00	[2.16]	12.43	0.08	[0.17]	0.54	[2.09]	11.86	0.05	[0.18]	0.46	[1.20]	6.74	0.07	[0.18]
United Kingdom	-0.03	[2.21]	12.02	-0.00	[0.18]	2.40	[1.76]	10.17	0.24	[0.18]	-2.44	[1.23]	6.63	-0.37	[0.17]
Equally-weighted	1.42	[1.16]	6.53	0.22	[0.18]	2.64	[1.01]	5.88	0.45	[0.19]	-1.22	[0.66]	3.68	-0.33	[0.17]
	Panel B: Yield Curve Slopes														
Australia	1.70	[2.45]	14.11	0.12	[0.18]	4.05	[2.12]	11.96	0.34	[0.18]	-2.35	[1.27]	7.24	-0.33	[0.18]
Canada	-1.47	[1.59]	9.06	-0.16	[0.18]	1.21	[1.34]	7.40	0.16	[0.18]	-2.68	[0.89]	5.01	-0.53	[0.18]
Germany	2.25	[2.13]	11.86	0.19	[0.17]	4.47	[1.98]	11.09	0.40	[0.18]	-2.22	[1.20]	6.66	-0.33	[0.17]
Japan	1.43	[2.43]	13.57	0.11	[0.17]	4.77	[1.97]	11.08	0.43	[0.18]	-3.34	[1.48]	8.37	-0.40	[0.18]
New Zealand	2.20	[3.05]	16.85	0.13	[0.18]	5.17	[2.39]	12.99	0.40	[0.19]	-2.97	[1.81]	10.27	-0.29	[0.18]
Norway	-0.69	[2.25]	12.88	-0.05	[0.18]	1.35	[2.01]	11.01	0.12	[0.18]	-2.04	[1.45]	8.20	-0.25	[0.18]
Sweden	-0.98	[2.11]	11.97	-0.08	[0.17]	2.37	[2.03]	11.30	0.21	[0.18]	-3.34	[1.31]	7.17	-0.47	[0.18]
Switzerland	2.18	[2.30]	12.42	0.18	[0.18]	4.28	[2.23]	11.80	0.36	[0.18]	-2.10	[1.23]	6.72	-0.31	[0.18]
United Kingdom	-2.07	[2.13]	12.00	-0.17	[0.18]	0.83	[1.80]	10.19	0.08	[0.17]	-2.91	[1.17]	6.61	-0.44	[0.18]
Equally-weighted	0.51	[1.47]	8.15	0.06	[0.18]	3.17	[1.30]	7.10	0.45	[0.19]	-2.66	[0.92]	5.05	-0.53	[0.19]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate (or the foreign yield curve slope is lower than the U.S. yield curve slope), and go long the U.S. bond and short the foreign country bond when the U.S. interest rate is higher than the country's interest rate (or the U.S. yield curve slope is lower than the foreign yield curve slope). Results based on interest rate levels are reported in Panel A and results based on interest rate slopes are reported in Panel B. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). The holding period is one month. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 10/1983–12/2015.

Table A12: Dollar Bond Return Differential Predictability, Interest Rates, Three-month Horizon

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	α	s.e.	β	s.e.	$R^2(\%)$	p-value	
	Panel A: Coupon Bonds																
Australia	-0.02	[0.03]	0.94	[0.81]	0.56	-0.03	[0.02]	1.37	[0.52]	2.23	0.00	[0.02]	-0.43	[0.57]	0.29	0.65	490
Canada	0.01	[0.02]	-0.37	[0.56]	-0.08	-0.01	[0.01]	1.21	[0.47]	1.91	0.02	[0.01]	-1.57	[0.34]	7.31	0.03	490
Germany	0.01	[0.02]	1.34	[1.08]	1.19	0.01	[0.02]	1.77	[0.88]	2.57	-0.00	[0.01]	-0.43	[0.53]	0.21	0.76	490
Japan	0.06	[0.02]	2.48	[0.78]	4.09	0.06	[0.02]	2.71	[0.61]	7.06	-0.00	[0.01]	-0.22	[0.52]	-0.11	0.82	490
New Zealand	-0.06	[0.04]	1.26	[0.75]	1.22	-0.06	[0.03]	1.94	[0.46]	6.94	0.00	[0.03]	-0.68	[0.58]	0.71	0.44	490
Norway	-0.02	[0.02]	1.02	[0.57]	1.36	-0.02	[0.02]	1.51	[0.54]	4.69	-0.00	[0.01]	-0.49	[0.37]	0.64	0.53	490
Sweden	-0.01	[0.02]	-0.46	[0.92]	0.04	-0.01	[0.02]	0.33	[0.98]	-0.04	0.00	[0.01]	-0.78	[0.45]	1.47	0.56	490
Switzerland	0.02	[0.02]	1.36	[0.77]	2.05	0.04	[0.02]	1.99	[0.69]	4.82	-0.02	[0.01]	-0.63	[0.41]	1.09	0.54	490
United Kingdom	-0.04	[0.03]	1.78	[1.11]	1.71	-0.03	[0.02]	2.06	[0.92]	3.86	-0.00	[0.01]	-0.29	[0.61]	-0.07	0.84	490
Panel	–	–	1.06	[0.46]	0.91	–	–	1.63	[0.43]	3.64	–	–	-0.57	[0.28]	0.90	0.04	4410
Joint zero (p-value)	0.12		0.00			0.01		0.00			0.39		0.00			0.66	
	Panel B: Zero-Coupon Bonds																
Australia	-0.04	[0.03]	2.17	[1.28]	2.88	-0.01	[0.03]	1.45	[0.83]	1.55	-0.03	[0.02]	0.72	[0.91]	0.54	0.64	344
Canada	0.00	[0.02]	0.49	[0.77]	-0.12	-0.00	[0.02]	1.47	[0.53]	2.15	0.01	[0.01]	-0.98	[0.56]	1.56	0.29	357
Germany	0.01	[0.02]	1.53	[0.93]	1.36	0.01	[0.02]	1.72	[0.81]	2.42	0.00	[0.01]	-0.19	[0.59]	-0.16	0.88	490
Japan	0.02	[0.03]	1.88	[1.02]	1.61	0.06	[0.03]	2.66	[0.94]	4.73	-0.03	[0.02]	-0.78	[0.57]	0.48	0.57	369
New Zealand	-0.02	[0.06]	1.31	[2.02]	0.18	0.03	[0.06]	0.28	[2.03]	-0.30	-0.05	[0.03]	1.03	[1.20]	0.49	0.72	309
Norway	-0.04	[0.03]	1.90	[1.68]	1.22	0.00	[0.03]	0.43	[1.78]	-0.36	-0.05	[0.02]	1.47	[1.03]	2.05	0.55	213
Sweden	-0.01	[0.02]	1.95	[1.28]	1.82	-0.01	[0.02]	1.52	[1.18]	1.33	0.00	[0.01]	0.43	[0.92]	-0.13	0.81	274
Switzerland	-0.00	[0.02]	1.91	[0.97]	2.42	0.02	[0.02]	2.51	[1.08]	4.59	-0.03	[0.01]	-0.60	[0.74]	0.29	0.68	333
United Kingdom	-0.05	[0.03]	2.28	[1.32]	2.36	-0.03	[0.02]	1.84	[1.04]	3.18	-0.03	[0.02]	0.45	[0.80]	-0.04	0.79	441
Panel	–	–	1.81	[0.63]	1.83	–	–	1.72	[0.64]	2.31	–	–	0.08	[0.35]	0.05	0.81	3130
Joint zero (p-value)	0.53		0.02			0.58		0.00			0.02		0.38			0.98	

Notes: The table reports regression results obtained when regressing the bond dollar return difference, defined as the difference between the log return on foreign bonds (expressed in U.S. dollars) and the log return of U.S. bonds in U.S. dollars, or the currency excess return, defined as the difference between the log return on foreign Treasury bills (expressed in U.S. dollars) and the log return of U.S. Treasury bills in U.S. dollars, or the bond local currency return difference, defined as the difference between the log return on foreign bonds (expressed in local currency terms) and the log return of U.S. bonds in U.S. dollars, on the corresponding interest rate differential, defined as the difference between the foreign nominal interest rate and the U.S. nominal interest rate. Panel A uses 10-year coupon bonds, whereas Panel B uses zero-coupon bonds. The holding period is three months and returns are sampled monthly. The log returns and the interest rate differentials are annualized. The sample period is 1/1975–12/2015. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A13: Dollar Bond Return Differential Predictability, Yield Curve Slopes, Three-Month Horizon

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$					Slope Diff.	Obs.
	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	α	s.e.	β	s.e.	R^2(%)	p-value	
	Panel A: Coupon Bonds																
Australia	0.01	[0.02]	0.71	[1.28]	-0.03	-0.00	[0.02]	-1.52	[0.96]	1.01	0.02	[0.01]	2.24	[0.77]	5.18	0.16	490
Canada	0.02	[0.01]	2.18	[0.76]	2.34	-0.00	[0.01]	-0.99	[0.57]	0.62	0.02	[0.01]	3.17	[0.44]	17.47	0.00	490
Germany	-0.00	[0.02]	0.15	[1.53]	-0.20	-0.00	[0.02]	-2.06	[1.14]	1.55	0.00	[0.01]	2.21	[0.85]	4.99	0.25	490
Japan	-0.00	[0.02]	-1.25	[1.21]	0.31	-0.01	[0.02]	-3.82	[0.95]	6.57	0.01	[0.01]	2.56	[0.75]	5.75	0.10	490
New Zealand	0.05	[0.04]	2.10	[2.06]	1.69	-0.00	[0.03]	-1.11	[1.13]	0.91	0.06	[0.02]	3.21	[1.16]	9.72	0.17	490
Norway	-0.01	[0.02]	-0.48	[0.91]	-0.05	-0.01	[0.02]	-1.80	[0.86]	2.96	-0.00	[0.01]	1.32	[0.52]	2.60	0.29	490
Sweden	0.01	[0.02]	2.73	[1.22]	4.64	0.01	[0.02]	0.73	[1.25]	0.25	0.01	[0.01]	2.01	[0.66]	5.85	0.25	490
Switzerland	-0.01	[0.02]	0.04	[0.99]	-0.20	-0.02	[0.02]	-2.76	[1.00]	3.49	0.01	[0.01]	2.80	[0.50]	9.66	0.05	490
United Kingdom	0.01	[0.02]	0.44	[1.45]	-0.13	-0.01	[0.02]	-2.27	[1.26]	2.91	0.02	[0.01]	2.70	[0.64]	7.40	0.16	490
Panel	–	–	0.85	[0.78]	0.17	–	–	-1.55	[0.67]	1.55	–	–	2.41	[0.41]	6.90	0.00	4410
Joint zero (p-value)	0.78		0.07			0.98		0.00			0.00		0.00			0.00	
	Panel B: Zero-Coupon Bonds																
Australia	0.02	[0.03]	-0.28	[2.08]	-0.27	0.01	[0.03]	-1.49	[1.80]	0.37	0.01	[0.02]	1.21	[1.33]	0.52	0.66	344
Canada	0.02	[0.02]	1.43	[1.05]	0.43	0.00	[0.01]	-1.39	[0.67]	0.86	0.02	[0.01]	2.82	[0.68]	7.74	0.02	357
Germany	0.01	[0.02]	0.58	[0.98]	-0.06	-0.01	[0.02]	-1.55	[0.86]	1.16	0.01	[0.01]	2.12	[0.74]	3.77	0.10	490
Japan	-0.03	[0.03]	-1.77	[1.33]	0.60	-0.04	[0.02]	-5.01	[1.22]	9.02	0.01	[0.02]	3.24	[0.85]	6.53	0.07	369
New Zealand	0.04	[0.04]	1.82	[2.35]	0.34	0.05	[0.04]	0.83	[2.47]	-0.16	-0.00	[0.02]	0.98	[1.09]	0.19	0.77	309
Norway	-0.02	[0.03]	-0.57	[1.83]	-0.37	0.01	[0.03]	0.29	[1.99]	-0.44	-0.03	[0.02]	-0.86	[1.22]	0.09	0.75	213
Sweden	0.01	[0.02]	1.20	[2.03]	0.03	-0.00	[0.02]	-0.51	[2.06]	-0.28	0.02	[0.02]	1.71	[1.21]	1.49	0.56	274
Switzerland	-0.03	[0.02]	-0.49	[1.03]	-0.18	-0.03	[0.02]	-2.59	[1.26]	3.31	0.00	[0.01]	2.10	[0.84]	4.82	0.20	333
United Kingdom	0.00	[0.03]	0.12	[1.64]	-0.22	-0.01	[0.02]	-1.64	[1.42]	1.39	0.02	[0.02]	1.76	[0.95]	1.51	0.42	441
Panel	–	–	0.10	[0.80]	-0.01	–	–	-1.75	[0.91]	1.33	–	–	1.85	[0.48]	2.48	0.00	3130
Joint zero (p-value)	0.65		0.81			0.56		0.00			0.29		0.00			0.12	

Notes: The table reports regression results obtained when regressing the bond dollar return difference, defined as the difference between the log return on foreign bonds (expressed in U.S. dollars) and the log return of U.S. bonds in U.S. dollars, or the currency excess return, defined as the difference between the log return on foreign Treasury bills (expressed in U.S. dollars) and the log return of U.S. Treasury bills in U.S. dollars, or the bond local currency return difference, defined as the difference between the log return on foreign bonds (expressed in local currency terms) and the log return of U.S. bonds in U.S. dollars, on the corresponding yield curve slope differential, defined as the difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope. Panel A uses 10-year coupon bonds, whereas Panel B uses zero-coupon bonds. The holding period is three months and returns are sampled monthly. The log returns and the yield curve slope differentials are annualized. The sample period is 1/1975–12/2015. In individual country regressions, standard errors are obtained with a Newey and West (1987) approximation of the spectral density matrix, with the lag truncation parameter (kernel bandwidth) equal to 6. Panel regressions include country fixed effects, and standard errors are obtained using the Driscoll and Kraay (1998) methodology, with the lag truncation parameter (kernel bandwidth) equal to 6.

Table A14: Dynamic Long-Short Interest Rate Foreign and U.S. Bond Portfolios, Three-Month Holding Period

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Coupon Bonds														
Australia	2.63	[2.02]	14.36	0.18	[0.16]	3.31	[1.68]	11.68	0.28	[0.17]	-0.69	[1.19]	8.21	-0.08	[0.15]
Canada	0.14	[1.29]	8.38	0.02	[0.15]	1.07	[1.04]	6.67	0.16	[0.16]	-0.93	[0.68]	4.61	-0.20	[0.15]
Germany	2.50	[1.95]	12.22	0.20	[0.16]	3.25	[1.85]	11.35	0.29	[0.16]	-0.75	[1.15]	7.15	-0.10	[0.16]
Japan	0.55	[2.18]	14.51	0.04	[0.16]	1.14	[1.92]	12.14	0.09	[0.16]	-0.59	[1.39]	8.70	-0.07	[0.15]
New Zealand	-0.02	[2.90]	18.42	-0.00	[0.15]	3.31	[1.91]	12.62	0.26	[0.16]	-3.33	[1.87]	12.28	-0.27	[0.15]
Norway	2.21	[2.15]	13.55	0.16	[0.16]	3.43	[1.75]	11.24	0.31	[0.16]	-1.23	[1.49]	8.85	-0.14	[0.15]
Sweden	1.35	[2.09]	13.53	0.10	[0.16]	2.64	[1.84]	11.69	0.23	[0.16]	-1.29	[1.47]	8.89	-0.15	[0.15]
Switzerland	-0.09	[2.01]	12.79	-0.01	[0.15]	0.60	[2.03]	12.49	0.05	[0.16]	-0.69	[1.23]	7.75	-0.09	[0.16]
United Kingdom	1.56	[1.97]	13.81	0.11	[0.15]	2.58	[1.69]	10.97	0.23	[0.15]	-1.02	[1.27]	8.41	-0.12	[0.16]
Equally-weighted	1.20	[1.01]	6.63	0.18	[0.16]	2.37	[0.90]	5.94	0.40	[0.17]	-1.17	[0.58]	3.54	-0.33	[0.15]
	Panel B: Zero-Coupon Bonds														
Australia	4.51	[2.50]	13.50	0.33	[0.19]	5.13	[2.14]	11.78	0.44	[0.20]	-0.62	[1.58]	8.80	-0.07	[0.19]
Canada	0.17	[1.77]	9.70	0.02	[0.19]	1.31	[1.35]	7.43	0.18	[0.19]	-1.15	[0.96]	5.67	-0.20	[0.18]
Germany	2.86	[2.10]	13.05	0.22	[0.16]	3.40	[1.82]	11.27	0.30	[0.16]	-0.54	[1.44]	9.15	-0.06	[0.16]
Japan	0.07	[2.49]	13.98	0.00	[0.18]	-0.31	[2.26]	12.12	-0.03	[0.18]	0.38	[1.70]	9.22	0.04	[0.18]
New Zealand	2.24	[2.70]	12.92	0.17	[0.20]	4.08	[2.20]	11.73	0.35	[0.21]	-1.84	[1.61]	8.01	-0.23	[0.20]
Norway	-0.17	[3.36]	13.48	-0.01	[0.24]	0.68	[2.79]	11.88	0.06	[0.24]	-0.85	[1.96]	8.60	-0.10	[0.24]
Sweden	3.86	[2.65]	12.70	0.30	[0.21]	4.47	[2.25]	11.17	0.40	[0.22]	-0.60	[1.68]	8.46	-0.07	[0.21]
Switzerland	1.67	[2.33]	11.66	0.14	[0.19]	1.70	[2.24]	11.36	0.15	[0.19]	-0.03	[1.56]	7.85	-0.00	[0.19]
United Kingdom	2.04	[2.43]	15.57	0.13	[0.17]	2.75	[1.76]	10.88	0.25	[0.17]	-0.71	[1.86]	11.32	-0.06	[0.17]
Equally-weighted	1.56	[1.14]	6.68	0.23	[0.19]	2.28	[1.17]	6.58	0.35	[0.22]	-0.72	[0.67]	3.76	-0.19	[0.18]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign short-term interest rate is higher than the U.S. interest rate, and go long the U.S. bond and short the foreign country bond when the U.S. interest rate is higher than the country's interest rate. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). Panel A uses 10-year coupon bonds, whereas Panel B uses zero-coupon bonds. The holding period is three months. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 1/1975–12/2015 (or largest subset available), with the exception of the equally-weighted portfolio of zero-coupon bonds, which refers to the sample period 4/1985–12/2015.

Table A15: Dynamic Long-Short Yield Curve Slope Foreign and U.S. Bond Portfolios, Three-Month Holding Period

	Bond dollar return difference $rx^{(10),\$} - rx^{(10)}$					Currency excess return rx^{FX}					Bond local currency return diff. $rx^{(10),*} - rx^{(10)}$				
	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.	Mean	s.e.	Std.	SR	s.e.
	Panel A: Coupon Bonds														
Australia	0.71	[2.08]	14.42	0.05	[0.16]	2.58	[1.77]	11.73	0.22	[0.16]	-1.87	[1.19]	8.16	-0.23	[0.15]
Canada	-0.93	[1.29]	8.37	-0.11	[0.16]	1.48	[1.05]	6.65	0.22	[0.16]	-2.41	[0.66]	4.47	-0.54	[0.15]
Germany	1.08	[1.96]	12.28	0.09	[0.16]	3.36	[1.89]	11.35	0.30	[0.16]	-2.27	[1.07]	7.07	-0.32	[0.15]
Japan	0.65	[2.17]	14.51	0.04	[0.16]	4.22	[1.90]	11.97	0.35	[0.16]	-3.57	[1.31]	8.52	-0.42	[0.16]
New Zealand	-0.23	[2.84]	18.42	-0.01	[0.15]	3.11	[1.93]	12.63	0.25	[0.16]	-3.34	[1.83]	12.27	-0.27	[0.15]
Norway	0.40	[2.12]	13.59	0.03	[0.16]	2.54	[1.78]	11.30	0.23	[0.16]	-2.14	[1.41]	8.80	-0.24	[0.16]
Sweden	-2.32	[2.03]	13.49	-0.17	[0.15]	0.53	[1.85]	11.76	0.05	[0.16]	-2.86	[1.50]	8.79	-0.33	[0.14]
Switzerland	1.70	[1.95]	12.76	0.13	[0.16]	4.66	[1.96]	12.27	0.38	[0.16]	-2.96	[1.21]	7.62	-0.39	[0.15]
United Kingdom	-1.55	[2.07]	13.81	-0.11	[0.15]	1.48	[1.74]	11.02	0.13	[0.15]	-3.03	[1.35]	8.29	-0.36	[0.16]
Equally-weighted	-0.05	[1.26]	8.04	-0.01	[0.15]	2.66	[1.13]	7.10	0.38	[0.16]	-2.72	[0.79]	4.71	-0.58	[0.14]
	Panel B: Zero-Coupon Bonds														
Australia	3.81	[2.53]	13.55	-0.28	[0.19]	5.16	[2.10]	11.77	-0.44	[0.20]	-1.34	[1.58]	8.78	0.15	[0.19]
Canada	-0.57	[1.76]	9.70	0.06	[0.18]	1.69	[1.33]	7.41	-0.23	[0.20]	-2.26	[0.94]	5.59	0.40	[0.18]
Germany	1.08	[2.11]	13.12	-0.08	[0.16]	3.81	[1.83]	11.23	-0.34	[0.16]	-2.73	[1.42]	9.05	0.30	[0.16]
Japan	2.00	[2.52]	13.94	-0.14	[0.18]	4.89	[2.25]	11.87	-0.41	[0.19]	-2.89	[1.67]	9.11	0.32	[0.18]
New Zealand	0.66	[2.69]	12.96	-0.05	[0.20]	3.18	[2.23]	11.80	-0.27	[0.20]	-2.52	[1.59]	7.97	0.32	[0.21]
Norway	-0.86	[3.36]	13.47	0.06	[0.24]	-0.16	[2.80]	11.88	0.01	[0.24]	-0.70	[1.92]	8.60	0.08	[0.25]
Sweden	0.82	[2.70]	12.84	-0.06	[0.21]	2.25	[2.29]	11.33	-0.20	[0.21]	-1.42	[1.70]	8.43	0.17	[0.21]
Switzerland	1.78	[2.33]	11.65	-0.15	[0.20]	4.28	[2.20]	11.19	-0.38	[0.20]	-2.50	[1.55]	7.75	0.32	[0.19]
United Kingdom	-0.52	[2.45]	15.60	0.03	[0.16]	2.17	[1.76]	10.91	-0.20	[0.17]	-2.69	[1.85]	11.25	0.24	[0.17]
Equally-weighted	1.60	[1.40]	8.35	-0.19	[0.18]	4.40	[1.38]	7.84	-0.56	[0.21]	-2.80	[0.95]	5.62	0.50	[0.19]

Notes: For each country, the table presents summary return statistics of investment strategies that go long the foreign country bond and short the U.S. bond when the foreign yield curve slope is lower than the U.S. yield curve slope, and go long the U.S. bond and short the foreign country bond when the U.S. yield curve slope is lower than the foreign yield curve slope. The table reports the mean, standard deviation and Sharpe ratio (denoted SR) for the currency excess return (rx^{FX}, middle panel), for the foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*} - rx^{(10)}$, right panel) and for the foreign bond excess return on 10-year government bond indices in U.S. dollars ($rx^{(10),\$} - rx^{(10)}$, left panel). Panel A uses 10-year coupon bonds, whereas Panel B uses zero-coupon bonds. The holding period is three months. The table also presents summary return statistics for the equally-weighted average of the individual country strategies. The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The log returns are annualized. The data are monthly and the sample is 1/1975–12/2015 (or largest subset available), with the exception of the equally-weighted portfolio of zero-coupon bonds, which refers to the sample period 4/1985–12/2015.

B Robustness Checks on Cross-sectional Portfolio Results

This section consider further robustness checks for the cross-sectional results by extending the sample of countries, by sorting on the level of interest rates, and by sorting on the slope of the yield curve.

B.1 Portfolio Cross-Sectional Evidence: Different Sample Periods

We start by considering different sample periods. Table A16 reports the results for the pre-crisis 10/1983-12/2007 sample, Table A17 for the pre-crisis 1/1975-12/2007 sample and Table A18 for the 10/1983-12/2015 sample. In all three tables, we focus on the benchmark set of G-10 countries and we consider currency portfolios sorted either on deviations of the short-term interest rate from its 10-year rolling mean or on the level of the yield curve slope. The results are consistent across sample periods and also consistent with the findings reported in the benchmark sample: the long-short portfolios do not produce statistically significant dollar bond returns.

Table A16: Cross-sectional Predictability: Bond Portfolios (10/1983 - 12/2007 Sample Period)

		Sorted by Short-Term Interest Rates				Sorted by Yield Curve Slopes			
Portfolio		1	2	3	$3-1$	1	2	3	$1-3$
		Panel A: Portfolio Characteristics							
Inflation rate	Mean	2.23	2.39	3.79	1.56	4.01	2.70	1.70	2.30
	s.e.	[0.17]	[0.18]	[0.21]	[0.23]	[0.22]	[0.19]	[0.16]	[0.22]
	Std	0.84	0.89	1.07	1.13	1.08	0.93	0.81	1.12
Rating	Mean	1.59	1.36	1.51	-0.09	1.58	1.47	1.40	0.18
	s.e.	[0.03]	[0.02]	[0.03]	[0.06]	[0.02]	[0.03]	[0.03]	[0.04]
Rating (adj. for outlook)	Mean	1.64	1.40	1.69	0.05	1.73	1.53	1.46	0.27
	s.e.	[0.04]	[0.02]	[0.03]	[0.07]	[0.03]	[0.03]	[0.03]	[0.05]
$y_t^{(10),*} - r_t^{*,f}$	Mean	1.20	0.70	-0.55	-1.74	-0.99	0.67	1.68	-2.67
		Panel B: Currency Excess Returns							
$-\Delta s_{t+1}$	Mean	0.46	2.18	1.73	1.26	1.30	2.02	1.06	0.25
$r_t^{f,*} - r_t^f$	Mean	0.42	0.73	2.84	2.42	3.94	0.81	-0.75	4.69
rx_{t+1}^{FX}	Mean	0.88	2.92	4.57	3.69	5.24	2.83	0.30	4.94
	s.e.	[1.55]	[1.82]	[1.87]	[1.51]	[1.86]	[1.68]	[1.66]	[1.64]
	SR	0.12	0.33	0.50	0.50	0.56	0.34	0.04	0.61
		Panel C: Local Currency Bond Excess Returns							
$rx_{t+1}^{(10),*}$	Mean	3.32	2.72	0.12	-3.19	-0.57	2.32	4.42	-4.98
	s.e.	[0.85]	[0.86]	[0.98]	[1.03]	[0.99]	[0.81]	[0.92]	[1.02]
	SR	0.78	0.64	0.03	-0.62	-0.12	0.58	0.96	-1.01
		Panel D: Dollar Bond Excess Returns							
$rx_{t+1}^{(10),\$}$	Mean	4.20	5.64	4.69	0.49	4.67	5.15	4.72	-0.05
	s.e.	[1.89]	[2.08]	[2.09]	[1.78]	[2.04]	[1.96]	[2.01]	[1.93]
	SR	0.45	0.55	0.45	0.06	0.45	0.54	0.47	-0.01
$rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$	Mean	0.32	1.77	0.82	0.49	0.79	1.27	0.84	-0.05
	s.e.	[2.02]	[2.08]	[2.29]	[1.78]	[2.20]	[2.07]	[1.99]	[1.93]

Notes: The countries are sorted by the level of their short term interest rates in deviation from the 10-year mean into three portfolios (left section) or the slope of their yield curves (right section). The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The table reports the average inflation rate, the standard deviation of the inflation rate, the average credit rating, the average credit rating adjusted for outlook, the average slope of the yield curve ($y^{(10),*} - r^{*,f}$), the average change in exchange rates (Δs), the average interest rate difference ($r^{f,*} - r^f$), the average currency excess return (rx^{FX}), the average foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*}$) and in U.S. dollars ($rx^{(10),\$}$), as well as the difference between the average foreign bond excess return in U.S. dollars and the average U.S. bond excess return ($rx^{(10),\$} - rx^{(10)}$). For the excess returns, the table also reports their Sharpe ratios (denoted SR). The holding period is one month. The log returns are annualized. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The data are monthly and the sample is 10/1983–12/2007.

Table A17: Cross-sectional Predictability: Bond Portfolios (1/1975 - 12/2007 Sample Period)

		Sorted by Short-Term Interest Rates				Sorted by Yield Curve Slopes			
Portfolio		1	2	3	$3-1$	1	2	3	$1-3$
		Panel A: Portfolio Characteristics							
Inflation rate	Mean	3.15	3.94	5.79	2.65	5.79	3.95	3.15	2.64
	s.e.	[0.18]	[0.21]	[0.25]	[0.23]	[0.25]	[0.20]	[0.21]	[0.21]
	Std	1.05	1.25	1.44	1.31	1.39	1.14	1.23	1.27
Rating	Mean	1.44	1.26	1.37	-0.06	1.43	1.35	1.30	0.13
	s.e.	[0.03]	[0.02]	[0.02]	[0.04]	[0.02]	[0.02]	[0.02]	[0.03]
Rating (adj. for outlook)	Mean	1.48	1.41	1.77	0.29	1.79	1.49	1.38	0.41
	s.e.	[0.03]	[0.02]	[0.03]	[0.05]	[0.02]	[0.02]	[0.02]	[0.04]
$y_t^{(10),*} - r_t^{*,f}$	Mean	1.50	0.86	-0.68	-2.18	-1.09	0.78	1.99	-3.08
		Panel B: Currency Excess Returns							
$-\Delta s_{t+1}$	Mean	0.17	0.74	-0.12	-0.29	-0.41	0.83	0.37	-0.78
$r_t^{f,*} - r_t^f$	Mean	-0.53	0.37	2.96	3.49	3.65	0.39	-1.24	4.89
rx_{t+1}^{FX}	Mean	-0.37	1.11	2.84	3.21	3.24	1.21	-0.87	4.11
	s.e.	[1.43]	[1.51]	[1.51]	[1.31]	[1.55]	[1.47]	[1.52]	[1.48]
	SR	-0.04	0.13	0.33	0.43	0.37	0.15	-0.10	0.52
		Panel C: Local Currency Bond Excess Returns							
$rx_{t+1}^{(10),*}$	Mean	3.62	2.18	-1.09	-4.71	-1.77	1.99	4.49	-6.25
	s.e.	[0.75]	[0.74]	[0.86]	[0.92]	[0.81]	[0.73]	[0.75]	[0.85]
	SR	0.84	0.51	-0.22	-0.89	-0.38	0.47	0.97	-1.23
		Panel D: Dollar Bond Excess Returns							
$rx_{t+1}^{(10),\$}$	Mean	3.25	3.29	1.75	-1.50	1.47	3.20	3.61	-2.14
	s.e.	[1.75]	[1.77]	[1.79]	[1.55]	[1.80]	[1.71]	[1.85]	[1.71]
	SR	0.32	0.32	0.17	-0.17	0.15	0.33	0.34	-0.23
$rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$	Mean	0.84	0.88	-0.66	-1.50	-0.94	0.79	1.20	-2.14
	s.e.	[1.82]	[1.80]	[1.97]	[1.55]	[2.02]	[1.81]	[1.90]	[1.71]

Notes: The countries are sorted by the level of their short term interest rates in deviation from the 10-year mean into three portfolios (left section) or the slope of their yield curves (right section). The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The table reports the average inflation rate, the standard deviation of the inflation rate, the average credit rating, the average credit rating adjusted for outlook, the average slope of the yield curve ($y^{(10),*} - r^{*,f}$), the average change in exchange rates (Δs), the average interest rate difference ($r^{f,*} - r^f$), the average currency excess return (rx^{FX}), the average foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*}$) and in U.S. dollars ($rx^{(10),\$}$), as well as the difference between the average foreign bond excess return in U.S. dollars and the average U.S. bond excess return ($rx^{(10),\$} - rx^{(10)}$). For the excess returns, the table also reports their Sharpe ratios (denoted SR). The holding period is one month. The log returns are annualized. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The data are monthly and the sample is 1/1975–12/2007.

Table A18: Cross-sectional Predictability: Bond Portfolios (10/1983 - 12/2015 Sample Period)

		Sorted by Short-Term Interest Rates				Sorted by Yield Curve Slopes			
Portfolio		1	2	3	$3-1$	1	2	3	$1-3$
		Panel A: Portfolio Characteristics							
Inflation rate	Mean	2.15	2.16	3.04	0.89	3.30	2.33	1.71	1.59
	s.e.	[0.15]	[0.16]	[0.19]	[0.20]	[0.18]	[0.16]	[0.15]	[0.20]
	Std	0.86	0.90	1.11	1.11	1.07	0.97	0.86	1.12
Rating	Mean	1.57	1.32	1.63	0.06	1.68	1.48	1.36	0.32
	s.e.	[0.03]	[0.02]	[0.02]	[0.04]	[0.02]	[0.02]	[0.02]	[0.04]
Rating (adj. for outlook)	Mean	1.62	1.36	1.79	0.17	1.81	1.53	1.43	0.38
	s.e.	[0.03]	[0.02]	[0.03]	[0.05]	[0.02]	[0.02]	[0.02]	[0.04]
$y_t^{(10),*} - r_t^{*,f}$	Mean	1.30	0.81	-0.28	-1.58	-0.66	0.79	1.71	-2.37
		Panel B: Currency Excess Returns							
$-\Delta s_{t+1}$	Mean	-0.38	1.02	0.66	1.04	0.19	1.15	-0.05	0.24
$r_t^{f,*} - r_t^f$	Mean	0.65	0.87	2.47	1.82	3.49	0.91	-0.41	3.90
rx_{t+1}^{FX}	Mean	0.26	1.89	3.13	2.86	3.68	2.06	-0.46	4.14
	s.e.	[1.48]	[1.70]	[1.64]	[1.27]	[1.78]	[1.54]	[1.54]	[1.34]
	SR	0.03	0.20	0.34	0.40	0.37	0.24	-0.05	0.55
		Panel C: Local Currency Bond Excess Returns							
$rx_{t+1}^{(10),*}$	Mean	3.28	3.13	0.90	-2.38	0.10	2.62	4.59	-4.49
	s.e.	[0.78]	[0.80]	[0.82]	[0.83]	[0.84]	[0.72]	[0.82]	[0.82]
	SR	0.75	0.69	0.20	-0.51	0.02	0.63	0.99	-0.99
		Panel D: Dollar Bond Excess Returns							
$rx_{t+1}^{(10),\$}$	Mean	3.55	5.02	4.03	0.48	3.79	4.68	4.13	-0.34
	s.e.	[1.69]	[1.83]	[1.81]	[1.46]	[1.88]	[1.68]	[1.76]	[1.50]
	SR	0.37	0.49	0.39	0.06	0.36	0.48	0.41	-0.04
$rx_{t+1}^{(10),\$} - rx_{t+1}^{(10)}$	Mean	-0.43	1.04	0.05	0.48	-0.19	0.70	0.15	-0.34
	s.e.	[1.82]	[1.87]	[1.94]	[1.46]	[2.05]	[1.78]	[1.83]	[1.50]

Notes: The countries are sorted by the level of their short term interest rates in deviation from the 10-year mean into three portfolios (left section) or the slope of their yield curves (right section). The slope of the yield curve is measured by the difference between the 10-year yield and the one-month interest rate. The standard errors (denoted s.e. and reported between brackets) were generated by bootstrapping 10,000 samples of non-overlapping returns. The table reports the average inflation rate, the standard deviation of the inflation rate, the average credit rating, the average credit rating adjusted for outlook, the average slope of the yield curve ($y^{(10),*} - r^{*,f}$), the average change in exchange rates (Δs), the average interest rate difference ($r^{f,*} - r^f$), the average currency excess return (rx^{FX}), the average foreign bond excess return on 10-year government bond indices in foreign currency ($rx^{(10),*}$) and in U.S. dollars ($rx^{(10),\$}$), as well as the difference between the average foreign bond excess return in U.S. dollars and the average U.S. bond excess return ($rx^{(10),\$} - rx^{(10)}$). For the excess returns, the table also reports their Sharpe ratios (denoted SR). The holding period is one month. The log returns are annualized. The balanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The data are monthly and the sample is 10/1983–12/2015.

B.2 Sorting by Interest Rate Deviations

This section reports results for currency portfolios sorted on the deviation of the short-term interest rate from its 10-year rolling mean. We first consider the benchmark G-10 sample, but then we consider a more extended sample of developed and emerging market countries.

B.2.1 Benchmark G-10 Sample

Figure A1 plots the composition of the three currency portfolios sorted on interest rate deviations, ranked from low (Portfolio 1) to high (Portfolio 3), for the long 1/1951–12/2015 sample period.

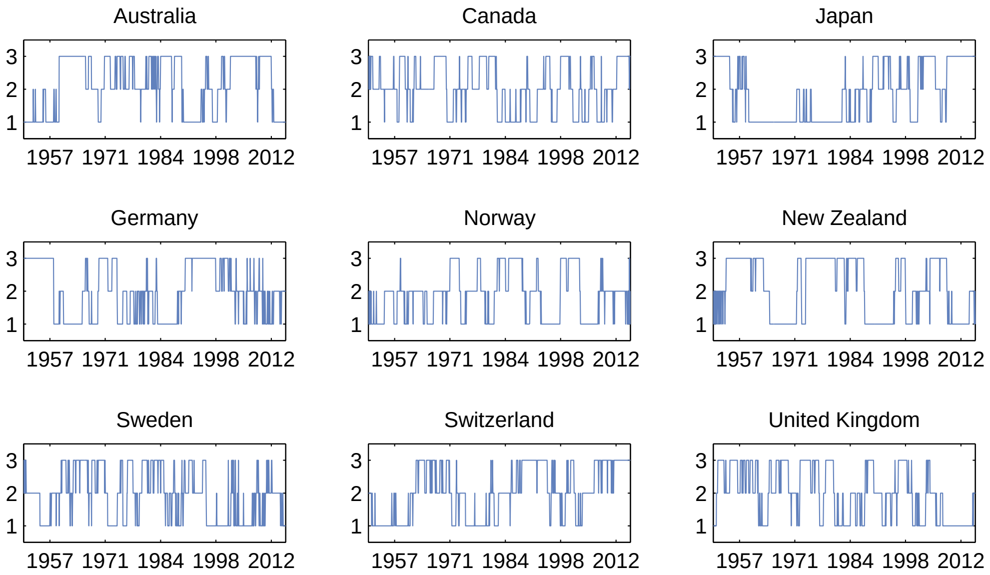

Figure A1: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 9 currencies sorted by the deviation of their short-term interest rates from the corresponding 10-year rolling mean. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

Figure A2 corresponds to the top right panel of Figure 1 in the main text. It presents the cumulative one-month log excess returns on investments in foreign Treasury bills and foreign 10-year bonds. Over the entire 1/1951 – 12/2015 sample period, the average currency log excess return of the carry trade strategy (long Portfolio 3, short Portfolio 1) is 2.52% per year, whereas the local currency bond log excess return is −3.81% per year. Thus, the interest rate carry trade implemented using 10-year bonds yields an average annualized dollar return of −1.29%, which is not statistically significant (bootstrap standard error of 0.94%). The average inflation rate of Portfolio 1 is 3.56% and its average credit rating is 1.44 (1.51 when adjusted for outlook), while the average inflation rate of Portfolio 3 is 4.72% and its average credit rating is 1.46 (1.81 when adjusted for outlook). Therefore, countries with high local currency bond term premia have low inflation and high credit ratings on average, whereas countries with low term premia have high average inflation rates and low average credit ratings, which suggests that the offsetting effect of the local currency bond excess returns is not due to compensation for inflation or credit risk. As seen in Table 3, our findings are very similar when we consider only the post-Bretton Woods period (1/1975 – 12/2015). Finally, we turn to the 7/1989 – 12/2015 period. The one-month average currency excess return of the carry trade strategy is 2.33%, largely offset by the local currency bond excess return of −1.33%. As a result, the average dollar bond excess return is 1.00%, which is not statistically significant, as its bootstrap standard error is 1.47%. Portfolio 1 has an average inflation rate of 1.91% and an average credit rating of 1.67 (1.72 when adjusted for outlook), whereas Portfolio 3 has an average inflation rate of 2.05% and an average credit rating of 1.67 (1.73 when adjusted for outlook).

We find very similar results when we increase the holding period k from 1 to 3 or 12 months: there is no evidence of statistically significant differences in dollar bond premia across the currency portfolios. In particular, for the entire 1/1951 – 12/2015 period, the

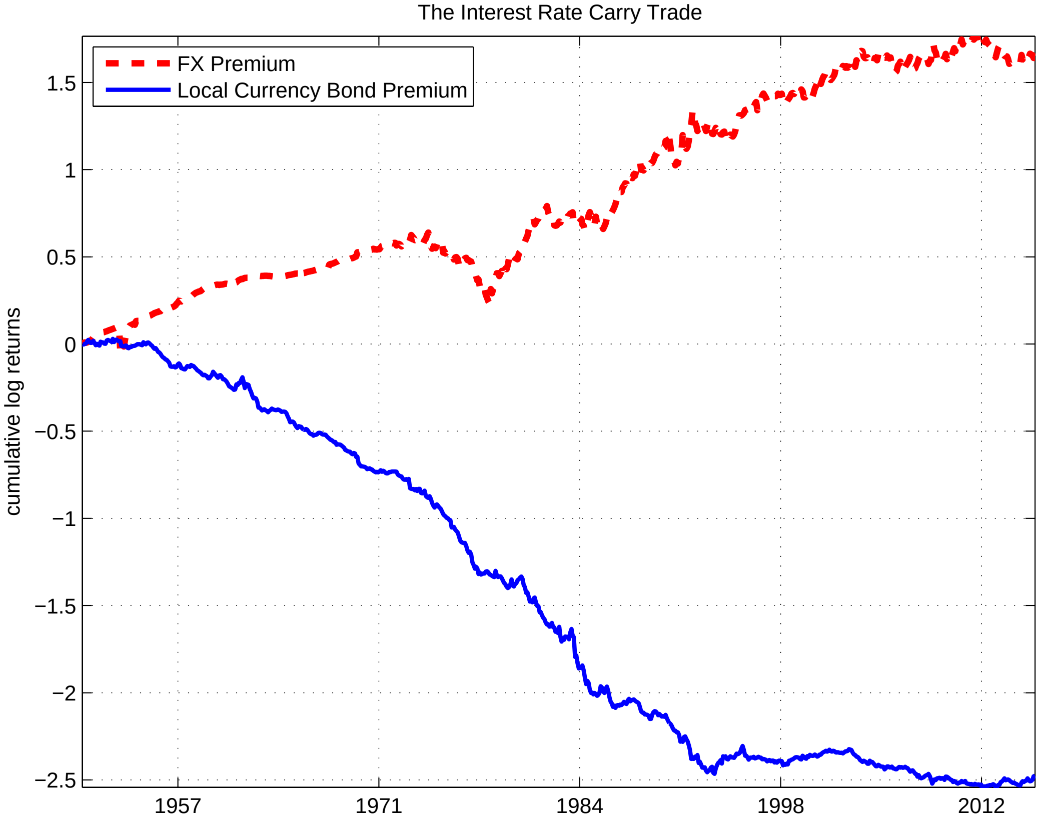

Figure A2: The Carry Trade and Term Premia – The figure presents the cumulative one-month log excess returns on investments in foreign Treasury bills and foreign 10-year bonds. The benchmark panel of countries includes Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. Countries are sorted every month into three portfolios by the level of the deviation of their one-month interest rate from its 10-year rolling mean. The returns correspond to a strategy going long in the Portfolio 3 and short in Portfolio 1. The sample period is 1/1951–12/2015.

annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant -0.68% (bootstrap standard error of 1.12%) for the 3-month holding period, as the average currency risk premium of 2.04% is offset by the average local currency bond premium of -2.72%. For the 12-month horizon, the average currency risk premium is 1.52%, which is almost fully offset by the average local currency bond premium of -1.68%, yielding an average dollar bond premium of -0.15% (bootstrap standard error of 1.08%). The corresponding average dollar bond premium for the post-Bretton Woods sample (1/1975 – 12/2015) is -0.88% for the 3-month holding period (average currency risk premium of 1.81%, average local currency bond premium of -2.68%) and -0.57% for the 12-month holding period (average currency risk premium of 1.28%, average local currency bond premium of -1.85%), neither of which is statistically significant (the bootstrap standard error is 1.39% and 1.55%, respectively). Finally, we consider the 7/1989 – 12/2015 period. The average dollar bond premium is 0.68% for the 3-month horizon (average currency risk premium of 1.39%, average local currency bond premium of -0.71%) and 0.86% for the 12-month horizon (average currency risk premium of 1.37%, average local currency bond premium of -0.51%). Neither of those average dollar bond premia is statistically significant, as their bootstrap standard error is 1.58% and 1.62%, respectively.

B.2.2 Developed Countries

Very similar patterns of risk premia emerge using larger sets of countries. In the sample of 20 developed countries (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, and the U.K.), we sort currencies in four portfolios, the composition of which is plotted in Figure A3.

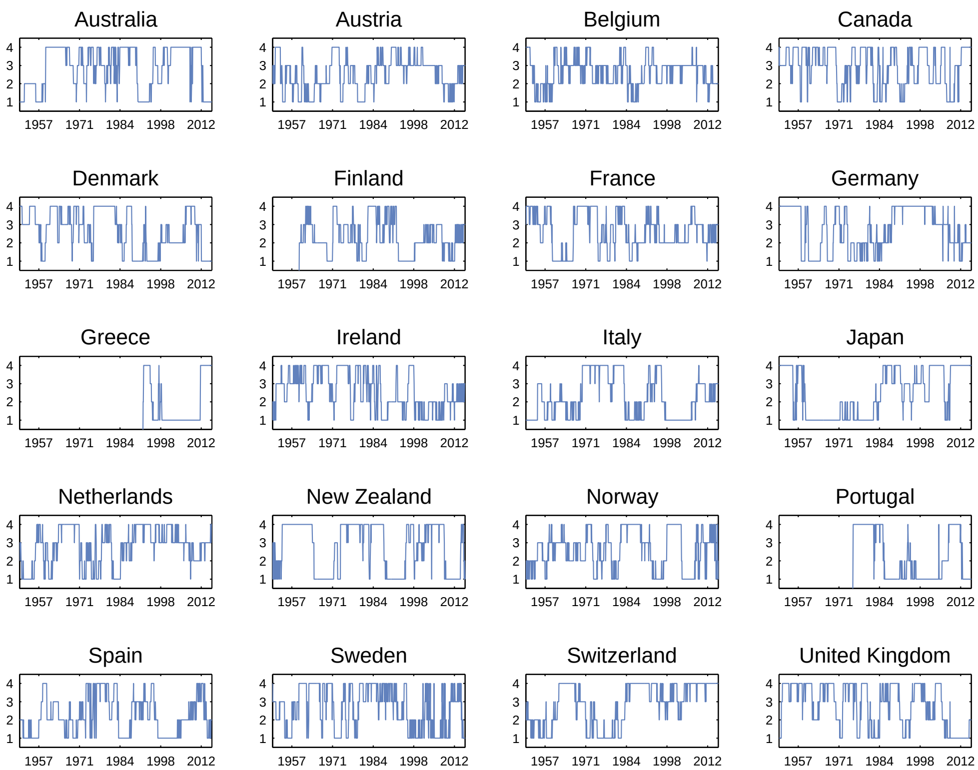

Figure A3: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

We start with 1-month holding period returns. Over the long sample period (1/1951 – 12/2015), the average currency log excess return of the carry trade is 1.32% per year, whereas the local currency bond log excess return is -4.77% per year. Therefore, the 10-year bond carry trade strategy yields a marginally significant average annualized return of -3.45% (bootstrap standard error of

1.97%). The average inflation rate of Portfolio 1 is 4.04% and its average credit rating is 2.68 (2.58 when adjusted for outlook); in comparison, the average inflation rate of Portfolio 4 is 5.05% and its average credit rating is 2.24 (2.41 when adjusted for outlook). We find similar results when we focus on the post-Bretton Woods sample: the average currency log excess return is 1.38% per year, offset by a local currency bond log excess return of −2.85%, so the 10-year bond carry trade strategy yields a statistically not significant annualized dollar excess return of −1.47% (bootstrap standard error of 1.15%). The average inflation rate of Portfolio 1 is 3.72% and its average credit rating is 2.71 (2.64 adjusted for outlook), whereas the average inflation rate of Portfolio 4 is 5.11% and its average credit rating is 2.31 (2.49 adjusted for outlook).

We now turn to longer holding periods. For the 1/1951 – 12/2015 sample, the annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant −1.15% (bootstrap standard error of 2.02%) for the 3-month holding period and a non-significant 0.45% (bootstrap standard error of 2.17%) for the 12-month holding period. The corresponding dollar excess returns for the post-Bretton Woods period are −0.11% for the 3-month holding period and 0.26% for the 12-month holding period, neither of which is statistically significant, as the bootstrap standard error is 3.21% and 1.61%, respectively.

B.2.3 Developed and Emerging Countries

Finally, we consider the sample of developed and emerging countries (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, India, Ireland, Italy, Japan, Mexico, Malaysia, the Netherlands, New Zealand, Norway, Pakistan, the Philippines, Poland, Portugal, South Africa, Singapore, Spain, Sweden, Switzerland, Taiwan, Thailand, and the United Kingdom), and sort currencies into five portfolios.

In particular, at the one-month horizon the average currency log excess return of the carry trade is 2.40% per year over the long sample period (1/1951 – 12/2015), which is more than offset by the local currency bond log excess return of −7.05% per year. As a result, the carry trade implemented using 10-year bonds yields a statistically significant average annualized return of 4.65% (the bootstrap standard error is 2.01%). The average inflation rate of Portfolio 1 is 4.59% and its average credit rating is 5.51 (4.96 when adjusted for outlook), whereas the average inflation rate of Portfolio 5 is 5.66% and its average credit rating is 4.70 (4.89 when adjusted for outlook). When we consider the post-Bretton Woods period (1/1975 – 12/2015), we get very similar results: the average currency log excess return is 3.04% per year, which is offset by a local currency bond log excess return of −6.36%, so the 10-year bond carry trade strategy yields a statistically significant annualized dollar return of −3.33% (bootstrap standard error of 1.29%). The average inflation rate of Portfolio 1 is 4.47% and its average credit rating is 5.45 (5.06 adjusted for outlook), whereas the average inflation rate of Portfolio 5 is 6.43% and its average credit rating is 4.78 (4.84 adjusted for outlook).

When we increase the holding period to 3 or 12 months, similar results emerge. For the long sample (1/1951 – 12/2015), the annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant −2.11% (bootstrap standard error of 2.07%) for the 3-month horizon and a non-significant −0.63% (bootstrap standard error of 2.18%) for the 12-month horizon. The corresponding dollar excess returns for the post-Bretton Woods period are −1.63% for the 3-month holding period and −0.70% for the 12-month holding period, both of which are marginally significant (bootstrap standard error of 1.47% and 1.62%, respectively).

B.3 Sorting by Interest Rate Levels

We now turn to currency portfolios sorted on interest rate levels (*not in deviation from the 10-year rolling mean*). We first consider the benchmark G-10 sample, but then we consider a more extended sample of developed and emerging market countries.

B.3.1 Benchmark Sample

Figure A4 plots the composition of the three interest rate-sorted currency portfolios, ranked from low (Portfolio 1) to high (Portfolio 3) interest rate currencies, for the long 1/1951–12/2015 sample period. Typically, Switzerland and Japan (after 1970) are funding currencies in Portfolio 1, while Australia and New Zealand are the carry trade investment currencies in Portfolio 3. The other currencies switch between portfolios quite often.

Over the entire 1/1951 – 12/2015 period, the average currency log excess return of the carry trade is 3.23% per year, whereas the local currency bond log excess return is −2.55% per year. As a result, the interest rate carry trade implemented using 10-year bonds yields an average annualized return of 0.68%, which is not statistically significant, as its bootstrap standard error is 1.07%. The average inflation rate of Portfolio 1 is 2.81% and its average credit rating is 1.33 (1.39 when adjusted for outlook), whereas the average inflation rate of Portfolio 3 is 5.15% and its average credit rating is 1.57 (1.92 when adjusted for outlook). Our findings are very similar when we consider only the post-Bretton Woods period (1/1975 – 12/2015): the average currency log excess return is 3.50% per year, largely offset by a local currency bond log excess return of −2.51%, so the 10-year bond carry trade strategy yields a statistically not significant annualized dollar return of 0.99% (bootstrap standard error of 1.57%). The average inflation

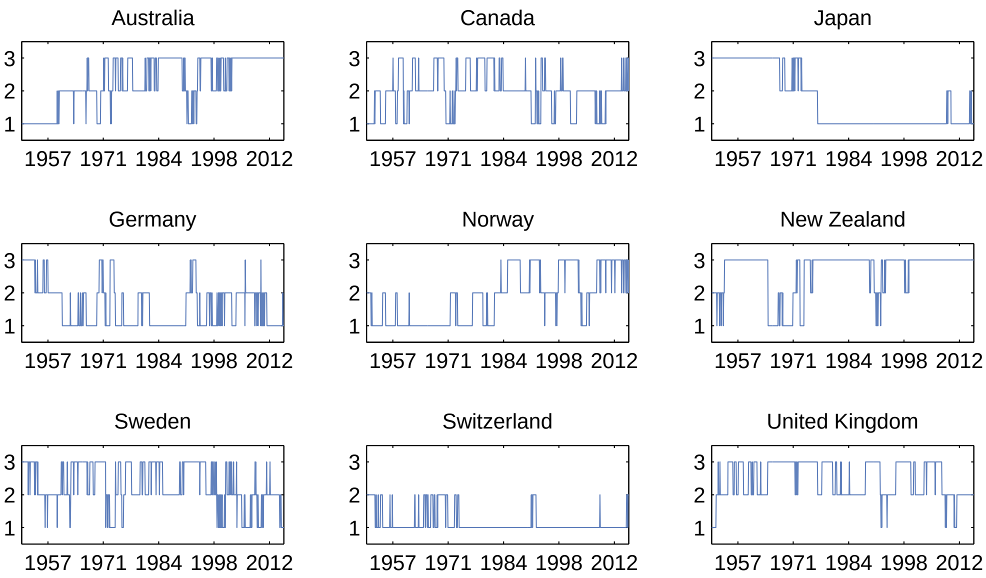

Figure A4: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 9 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

rate of Portfolio 1 is 2.00% and its average credit rating is 1.36 (1.41 when adjusted for outlook), whereas the average inflation rate of Portfolio 3 is 5.32% and its average credit rating is 1.60 (1.93 when adjusted for outlook).

We find very similar results when we increase the holding period: there is no evidence of statistically significant differences in dollar bond risk premia across the currency portfolios. In particular, for the entire 1/1951 – 12/2015 period, the annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant 1.03% (bootstrap standard error of 1.12%) for the 3-month holding period and a non-significant 1.23% (bootstrap standard error of 1.20%) for the 12-month holding period. The corresponding dollar excess returns for the post-Bretton Woods period are 1.15% for the 3-month holding period and 1.18% for the 12-month holding period, neither of which is statistically significant (bootstrap standard error of 1.65% and 1.69%, respectively).

B.3.2 Developed Countries

With coupon bonds, we consider two additional sets of countries: first, a larger sample of 20 developed countries (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, and the U.K.), and second, a large sample of 30 developed and emerging countries (the same as above, plus India, Mexico, Malaysia, the Netherlands, Pakistan, the Philippines, Poland, South Africa, Singapore, Taiwan, and Thailand). We also construct an extended version of the zero-coupon dataset which, in addition to the countries of the benchmark sample, includes the following countries: Austria, Belgium, the Czech Republic, Denmark, Finland, France, Hungary, Indonesia, Ireland, Italy, Malaysia, Mexico, the Netherlands, Poland, Portugal, Singapore, South Africa, and Spain. The data for the aforementioned extra countries are sourced from Bloomberg. The starting dates for the additional countries are as follows: 12/1994 for Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Portugal, Singapore, and Spain, 12/2000 for the Czech Republic, 3/2001 for Hungary, 5/2003 for Indonesia, 9/2001 for Malaysia, 8/2003 for Mexico, 12/2000 for Poland, and 1/1995 for South Africa.

We now turn to the sample of 20 developed countries. Figure A5 plots the composition of the four interest rate-sorted currency portfolios. As we can see, Switzerland and Japan (after 1970) are funding currencies in Portfolio 1, while Australia and New Zealand are carry trade investment currencies in Portfolio 4.

We start with 1-month holding period returns. Over the long sample period (1/1951 – 12/2015), the average currency log excess return of the carry trade is 2.73% per year, whereas the local currency bond log excess return is −2.15% per year. Therefore,

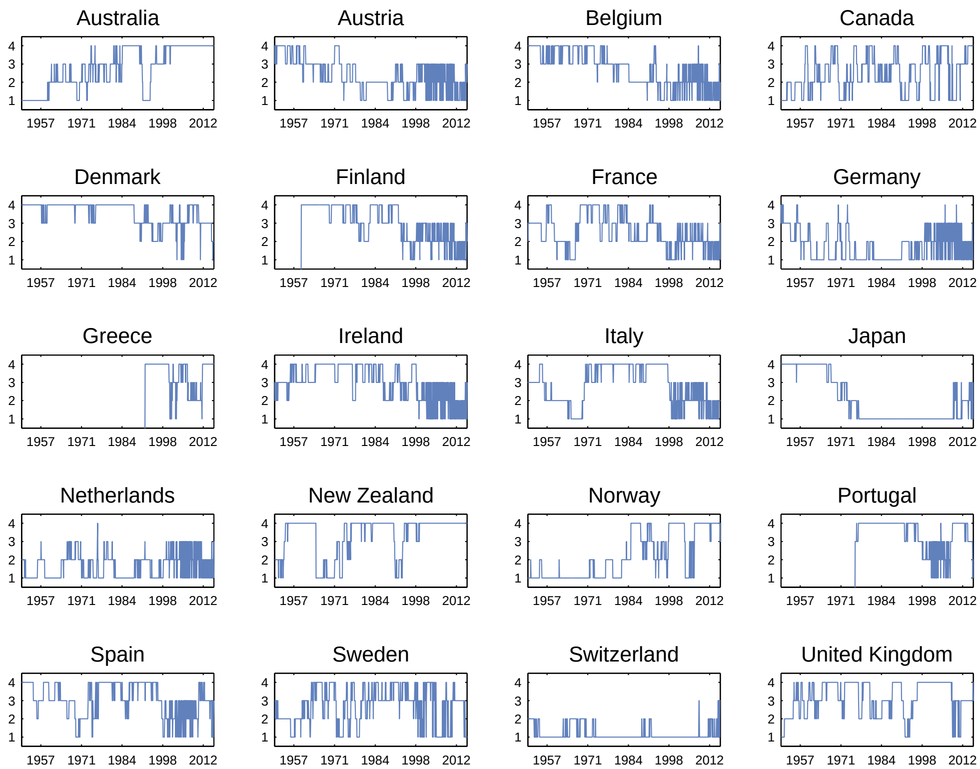

Figure A5: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

the interest rate carry trade implemented using 10-year bonds yields a non-statistically significant average dollar annualized return of 0.58% (the bootstrap standard error is 0.90%). The average inflation rate of Portfolio 1 is 3.04% and its average credit rating is 1.50 (1.54 when adjusted for outlook); the average inflation rate of Portfolio 4 is 5.73% and its average credit rating is 2.93 (3.02 when adjusted for outlook). We get very similar results when we focus on the post-Bretton Woods sample: the average currency log excess return is 2.81% per year, offset by a local currency bond log excess return of -1.37%, so the 10-year bond carry trade strategy yields a statistically not significant annualized return of 1.44% (bootstrap standard error of 1.33%). The average inflation rate of Portfolio 1 is 2.30% and its average credit rating is 1.55 (1.61 adjusted for outlook), whereas the average inflation rate of Portfolio 4 is 6.07% and its average credit rating is 2.97 (3.03 adjusted for outlook).

When we increase the holding period, we get very similar results. For the 1/1951 – 12/2015 sample, the annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant 1.15% (bootstrap standard error of 0.94%) for the 3-month holding period and a non-significant 1.48% (bootstrap standard error of 0.99%) for the 12-month holding period. The corresponding dollar excess returns for the post-Bretton Woods period are 1.92% for the 3-month holding period and 1.90% for the 12-month holding period, neither of which is statistically significant, as the bootstrap standard errors are 1.37% and 1.50%, respectively.

B.3.3 Developed and Emerging Countries

Finally, we consider the sample of developed and emerging countries and sort currencies into five portfolios.

We start by focusing on one-month returns. Over the long sample period (1/1951 – 12/2015), the average currency log excess return of the carry trade is 4.92% per year, largely offset by the local currency bond log excess return of -4.18% per year. As a result, the interest rate carry trade implemented using 10-year bonds yields a non-statistically significant average annualized return of 0.74% (the bootstrap standard error is 0.90%). The average inflation rate of Portfolio 1 is 3.17% and its average credit rating is 2.91 (2.75 when adjusted for outlook), whereas the average inflation rate of Portfolio 5 is 6.82% and its average credit rating is 6.59 (6.07 when adjusted for outlook). When we focus on the post-Bretton Woods sample, our findings are very similar: the average currency log excess return is 5.73% per year, which is offset by a local currency bond log excess return of −3.80%, so the 10-year bond carry trade strategy yields a statistically non-significant annualized return of 1.92% (the bootstrap standard error is 1.33%). The average inflation rate of Portfolio 1 is 2.49% and its average credit rating is 2.95 (2.90 adjusted for outlook); the average inflation rate of Portfolio 5 is 7.78% and its average credit rating is 6.60 (6.03 adjusted for outlook).

We now consider longer holding periods. For the long sample (1/1951 – 12/2015), the annualized dollar excess return of the carry trade strategy implemented using 10-year bonds is a non-significant 1.33% (bootstrap standard error of 1.01%) for the 3-month horizon and a marginally significant 1.94% (bootstrap standard error of 1.10%) for the 12-month horizon. The corresponding dollar excess returns for the post-Bretton Woods period are 2.56% for the 3-month holding period and 2.80% for the 12-month holding period, both of which are marginally significant (bootstrap standard error of 1.50% and 1.69%, respectively).

B.4 Sorting by Yield Curve Slopes

This section presents additional evidence on slope-sorted currency portfolios. We first consider the benchmark G-10 sample, but then we consider a more extended sample of developed and emerging market countries.

B.4.1 Benchmark Sample

Figure A6 presents the composition over time of the slope-sorted currency portfolios for the long sample period of 1/1951–12/2015.

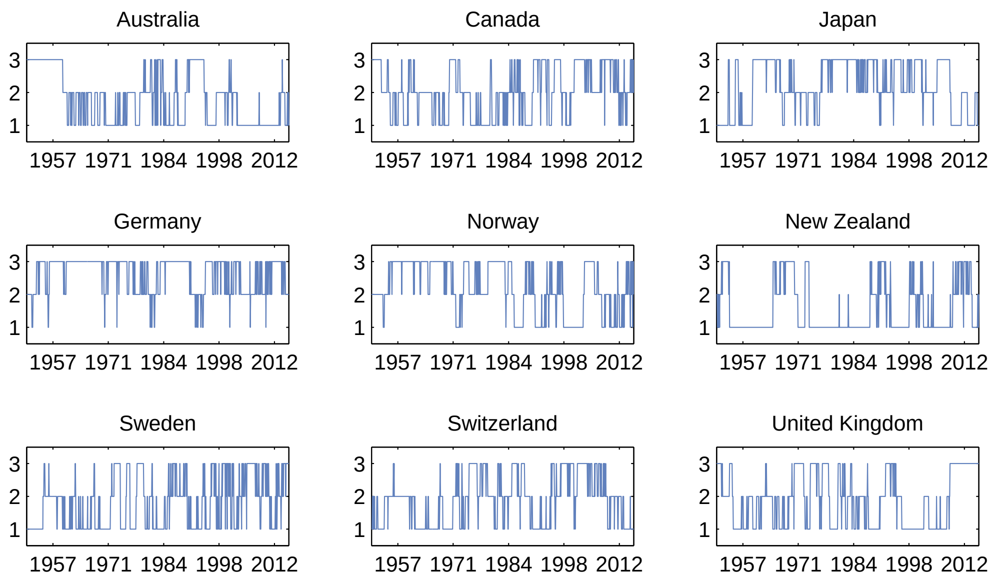

Figure A6: Composition of Slope-Sorted Portfolios — The figure presents the composition of portfolios of the currencies in the benchmark sample sorted by the slope of their yield curves. The portfolios are rebalanced monthly. The slope of the yield curve is measured by the spread between the 10-year bond yield and the one-month interest rate. Data are monthly, from 1/1951 to 12/2015.

Figure A7 corresponds to the lower left panel of Figure 1 in the main text. It presents the cumulative one-month log excess returns on investments in foreign Treasury bills and foreign 10-year bonds, starting in 1951. The returns correspond to an investment strategy going long in Portfolio 1 (flat yield curves, mostly high short-term interest rates) and short in Portfolio 3 (steep yield curves, mostly low short-term interest rates). Over the entire 1/1951 – 12/2015 period, the average currency log excess return of the slope carry trade is 3.01% per year, whereas the local currency bond log excess return is −5.46% per year. Therefore, the slope carry trade implemented using 10-year bonds results in an average return of −2.45% per year, which is statistically significant (bootstrap standard error of 0.98%). It is worth noting that neither inflation risk nor credit risk seem to be able to explain this offsetting effect: the average inflation rate of Portfolio 1, which has a low average term premium, is 4.71% and its average credit rating is 1.52 (1.84 when adjusted for outlook), whereas the average inflation rate of Portfolio 3, which has a high average term premium, is 3.51% and its average credit rating is 1.28 (1.37 when adjusted for outlook). As seen in Table 3, we get similar results when we focus only on the post-Bretton Woods period (1/1975 – 12/2015). Finally, we consider the 7/1989 – 12/2015 sample period. The one-month average currency excess return of the slope carry trade strategy is 4.41%, largely offset by the local currency bond excess return of −3.40%. As a result, the average dollar bond excess return is 1.02%, which is not statistically significant, as its bootstrap standard error is 1.32%. Portfolio 1 has an average inflation rate of 2.31% and an average credit rating of 1.71 (1.75 when adjusted for outlook), whereas Portfolio 3 has an average inflation rate of 1.51% and an average credit rating of 1.43 (1.49 when adjusted for outlook).

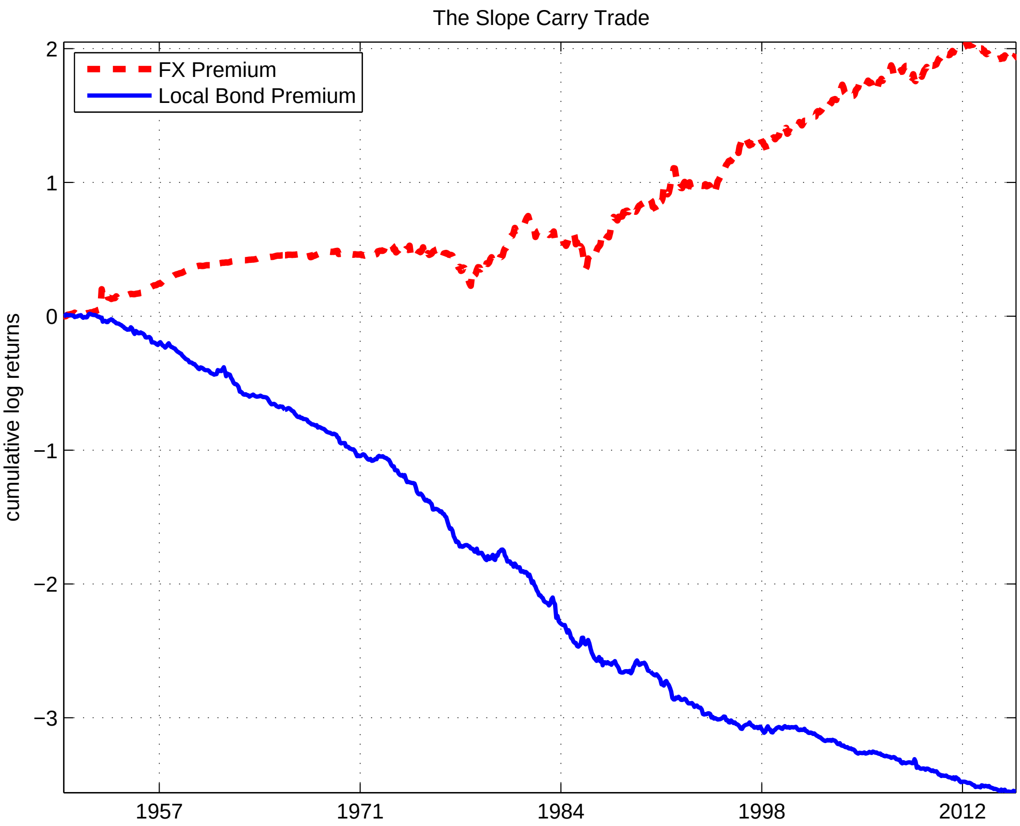

Figure A7: The Carry Trade and Term Premia: Conditional on the Slope of the Yield Curve – The figure presents the cumulative one-month log returns on investments in foreign Treasury bills and foreign 10-year bonds. The benchmark panel of countries includes Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. Countries are sorted every month by the slope of their yield curves into three portfolios. The slope of the yield curve is measured by the spread between the 10-year bond yield and the one-month interest rate. The returns correspond to an investment strategy going long in Portfolio 1 and short in the Portfolio 3. The sample period is 1/1951–12/2015.

We now consider longer holding periods. Overall, we find no evidence of statistically significant differences in dollar bond risk premia across the currency portfolios. For the full 1/1951 – 12/2015 period, the annualized dollar excess return of the slope carry trade strategy implemented using 10-year bonds is a non-significant −1.58% (bootstrap standard error of 0.99%) for the 3-month holding period, as the average currency risk premium of 2.53% is more than offset by the average local currency term premium of −4.12%. For the 12-month holding period, the average currency risk premium is 1.98%, which is offset by the average local currency term premium of −3.15%, yielding an average non-significant dollar term premium of −1.17% (bootstrap standard error of 1.00%).

The corresponding dollar excess returns for the post-Bretton Woods period (1/1975 – 12/2015) are −0.88% for the 3-month holding period (average currency risk premium of 2.95%, average local currency term premium of −3.83%) and −0.50% for the 12-month holding period (average currency risk premium of 2.19%, average local currency term premium of −2.68%), neither which are is significant, as the bootstrap standard error is 1.43% and 1.46%, respectively. Finally, we turn to the 7/1989 – 12/2015 period. The average dollar bond premium is 0.98% for the 3-month horizon (average currency risk premium of 3.14%, average local currency bond premium of −2.16%) and 1.35% for the 12-month horizon (average currency risk premium of 2.75%, average local currency bond premium of −1.39%). Both of those dollar bond premia are non-significant, as their bootstrap standard error is 1.52% and 1.71%, respectively.

B.4.2 Developed Countries

In the sample of developed countries, the flat-slope currencies (Portfolio 1) are typically those of Australia, New Zealand, Denmark and the U.K., while the steep-slope currencies (Portfolio 4) are typically those of Germany, the Netherlands, and Japan. The portfolio compositions are plotted in Figure A8.

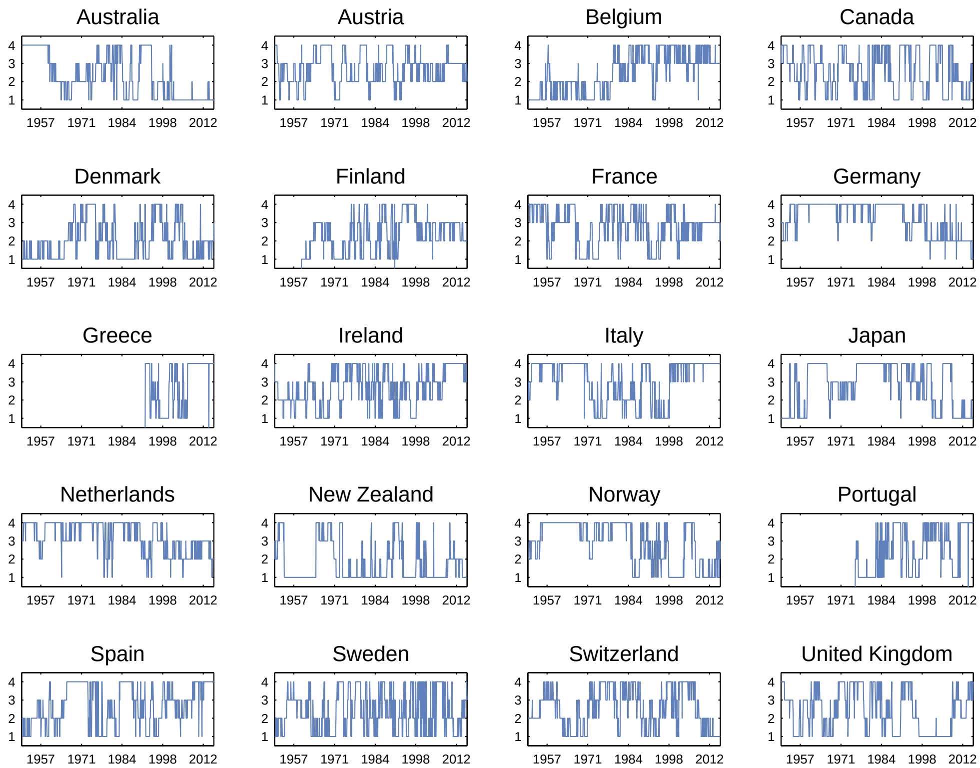

Figure A8: Composition of Slope-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their yield curve slopes. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

At the one-month horizon, the 2.50% spread in currency excess returns obtained in the full sample period (1/1951 – 12/2015) is more than offset by the −6.73% spread in local term premia. This produces a statistically significant average dollar excess return of −4.22% (bootstrap standard error of 1.02%) on a position that is long in the high yielding, low slope currencies (Portfolio 1) and short in the low yielding, high slope currencies (Portfolio 4). The average inflation rate of Portfolio 1 is 5.13% and its average credit rating is 2.20 (2.34 when adjusted for outlook), whereas the average inflation rate of Portfolio 4 is 3.97% and its average credit rating is 2.88 (2.97 when adjusted for outlook). Those results are essentially unchanged in the post-Bretton Woods period: the average currency excess return is 3.04%, more than offset by the average local currency bond excess return of −7.60%, so the slope carry trade yields an average excess return of −4.56%, which is statistically significant (bootstrap standard error of 1.48%).

The average inflation rate of Portfolio 1 is 5.36% and its average credit rating is 2.21 (2.34 when adjusted for outlook), whereas the average inflation rate of Portfolio 4 is 3.49% and its average credit rating is 3.04 (3.16 when adjusted for outlook).

We now turn to longer holding periods. In the 3-month horizon, investing in Portfolio 1 and shorting Portfolio 4 during the long sample period (1/1951 – 12/2015) yields an average currency excess return of 2.03% and an average local currency bond excess return of −5.13%, resulting in a statistically significant dollar bond excess return of −3.10% (bootstrap standard error of 1.11%). In the same period, the 12-month average currency excess return is 1.86% and the average local currency bond excess return is −3.53%, so the average dollar bond excess return is a non-significant −1.67% (bootstrap standard error of 1.42%). Similar results emerge when we focus on the post-Bretton Woods period. In the 3-month horizon, the average currency excess return is 2.31% and the average local currency bond excess return is −5.32%, yielding an average dollar bond excess return of −3.00%, which is marginally statistically significant (bootstrap standard error of 1.63%). In the 12-month horizon, the average currency excess return is 1.90% and the average local currency bond excess return is −3.42%, so the average dollar bond excess return is a non-significant −1.52% (bootstrap standard error of 2.22%).

B.4.3 Developed and Emerging Countries

In the entire sample of countries, the difference in currency risk premia at the one-month horizon is 3.44% per year, which is more than offset by a −9.84% difference in local currency term premia. As a result, investors earn a statistically significant −6.41% per annum (the bootstrap standard error is 1.06%) on a long-short bond position. As before, this involves going long the bonds of flat-yield-curve currencies (Portfolio 1), typically high interest rate currencies, and shorting the bonds of the steep-slope currencies (Portfolio 5), typically the low interest rate ones. The average inflation rate of Portfolio 1 is 5.77% and its average credit rating is 4.77 (4.74 when adjusted for outlook), whereas the average inflation rate of Portfolio 5 is 4.54% and its average credit rating is 5.62 (5.33 when adjusted for outlook). When we focus on the post-Bretton Woods period (1/1975 – 12/2015), we get very similar results: the average currency log excess return is 4.59% per year, which is more than offset by a local currency bond log excess return difference of −11.53%, so the 10-year bond carry trade strategy yields a statistically significant annualized return of −6.94% (bootstrap standard error of 1.51%). The average inflation rate of Portfolio 1 is 6.16% and its average credit rating is 4.79 (4.69 adjusted for outlook), whereas the average inflation rate of Portfolio 5 is 4.43% and its average credit rating is 5.73 (5.55 adjusted for outlook).

When we increase the holding period to 3 or 12 months, similar results emerge. For the long sample (1/1951 – 12/2015), the average annualized dollar excess return of the slope carry trade strategy (long Portfolio 1, short Portfolio 5) implemented using 10-year bonds is a statistically significant −5.32% (bootstrap standard error of 1.17%) for the 3-month horizon: the average currency excess return is 2.76%, more than offset by the average local currency bond excess return of −8.08%. For the 12-month horizon, the average currency excess return is 2.47% and the local currency bond excess return is −5.48%, so the average dollar excess return for the slope carry trade is −3.01% (statistically significant, as the bootstrap standard error is 1.29%). Finally, for the post-Bretton Woods period, the average 3-month currency excess return is 3.55% and the average local currency bond excess return is −9.22%, so the dollar excess return of the slope carry trade is −5.66% (statistically significant, as the bootstrap standard error is 1.73%). For the same period, the average 12-month currency excess return is 3.06% and the average local currency bond excess return is −5.83%, resulting in an average dollar excess return of −2.78% (not significant, given a bootstrap standard error of 1.97%).

C Foreign Bond Returns Across Maturities

This section reports additional results obtained with zero-coupon bonds. We start with the bond risk premia in our benchmark sample of G10 countries and then turn to a larger set of developed countries. We then show that holding period returns on zero-coupon bonds, once converted to a common currency (the U.S. dollar, in particular), become increasingly similar as bond maturities approach infinity.

C.1 Benchmark Countries

Figure A9 reports results for all maturities. The figure shows the local currency bond log excess returns in the top panels, the currency log excess returns in the middle panels, and the dollar bond log excess returns in the bottom panels. The top panels show that countries with the steepest local yield curves (Portfolio 3, center) exhibit local bond excess returns that are higher, and increase faster with maturity, than the flat yield curve countries (Portfolio 1, left). Thus, ignoring the effect of exchange rates, investors should invest in the short-term and long-term bonds of steep yield curve currencies.

Including the effect of currency fluctuations, by focusing on dollar returns, radically alters the results. The bottom panels of Figure A9 show that the dollar excess returns of Portfolio 1 are on average higher than those of Portfolio 3 at the short end of the yield curve, consistent with the carry trade results of Ang and Chen (2010). Yet, an investor who would attempt to replicate the short-maturity carry trade strategy at the long end of the maturity curve would incur losses on average: the long-maturity excess returns of flat yield curve currencies are lower than those of steep yield curve currencies, as currency risk premia more than offset term premia. This result is apparent in the lower panel on the right, which is the same as Figure 2 in the main text.

Figure A10 shows the results when sorting by the level of interest rates. The term structure is flat but not statistically significantly different from zero at longer horizons. The term structure is flat but not statistically significantly different from zero at longer horizons: the carry premium is 3.71% per annum (with a standard error of 1.80%), while the local currency 15-year bond premium is only -0.21% per annum (with a standard error of 1.76%), so the long-maturity dollar bond premium is 3.50% (with a standard error of 2.32%). Interest rates (in levels) do not predict bond excess returns in the cross-section in the second half of our sample (see top panels in Figure 1).

C.2 Developed Countries

When we tuning to the entire sample of developed countries, the results are very similar to those attained in our benchmark sample. An investor who buys the short-term bonds of flat-yield curve currencies and shorts the short-term bonds of steep-yield-curve currencies realizes a statistically significant dollar excess return of 4.20% per year on average (bootstrap standard error of 1.50%). However, at the long end of the maturity structure, this strategy generates negative and insignificant excess returns: the average annualized dollar excess return of an investor who pursues this strategy using 15-year bonds is −2.30% (bootstrap standard error of 2.49%). Our findings are presented graphically in Figure A11, which shows the local currency bond log excess returns in the top panels, the currency log excess returns in the middle panels, and the dollar bond log excess returns in the bottom panels as a function of maturity.

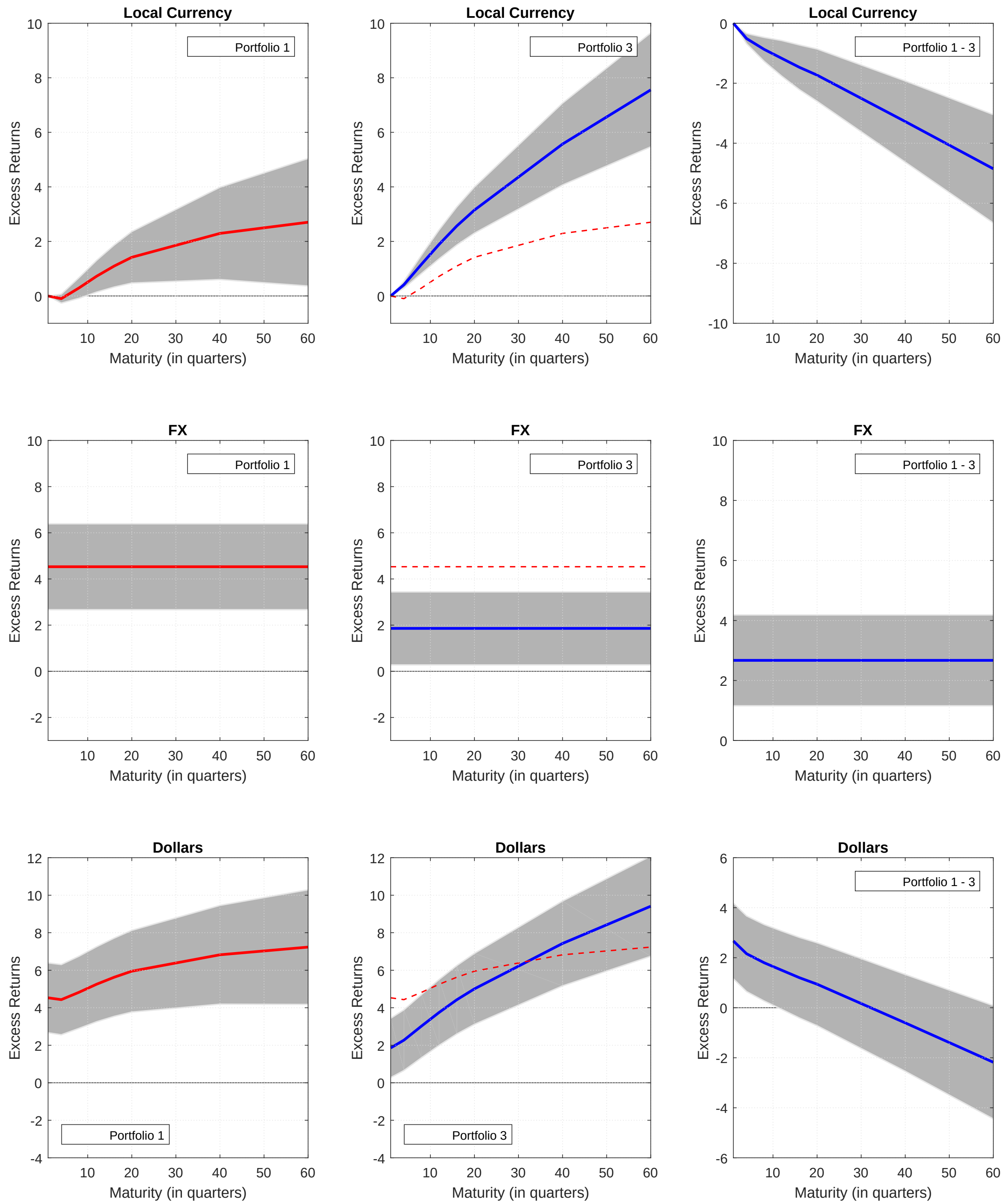

Figure A9: Dollar Bond Risk Premia Across Maturities— The figure shows the log excess returns on foreign bonds in local currency in the top panel, the currency excess return in the middle panel, and the log excess returns on foreign bonds in U.S. dollars in the bottom panel as a function of the bond maturities. The left panel focuses on Portfolio 1 (flat yield curve currencies) excess returns, while the middle panel reports Portfolio 3 (steep yield curve currencies) excess returns. The middle panels also report the Portfolio 1 excess returns in dashed lines for comparison. The right panel reports the difference. Data are monthly, from the zero-coupon dataset, and the sample window is 4/1985–12/2015. The unbalanced panel consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. The countries are sorted by the slope of their yield curves into three portfolios. The slope of the yield curve is measured by the difference between the 10-year yield and the 3-month interest rate at date t. The holding period is one quarter. The returns are annualized. The shaded areas correspond to one standard deviation above and below each point estimate. Standard deviations are obtained by bootstrapping 10,000 samples of non-overlapping returns.

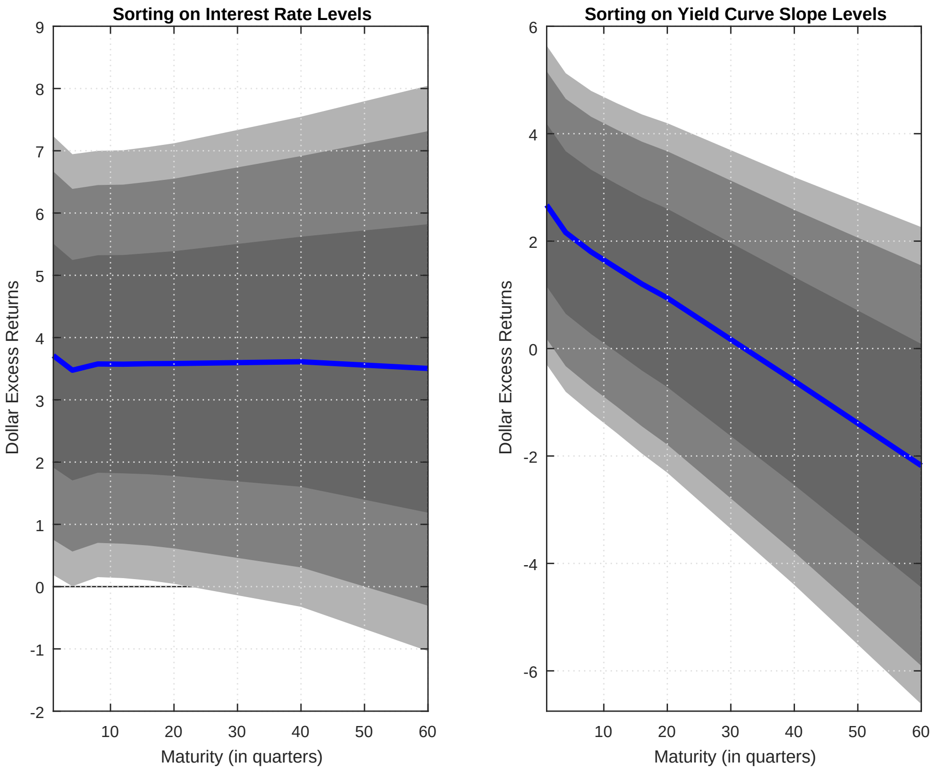

Figure A10: Long-Minus-Short Foreign Bond Risk Premia in U.S. Dollars— The figure shows the dollar log excess returns as a function of the bond maturities. Dollar excess returns correspond to the holding period returns expressed in U.S. dollars of investment strategies that go long and short foreign bonds of different countries. The unbalanced panel of countries consists of Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the U.K. At each date t, the countries are sorted by the slope of their yield curves into three portfolios. The first portfolio contains countries with flat yield curves while the last portfolio contains countries with steep yield curves. The slope of the yield curve is measured by the difference between the 10-year yield and the 3-month interest rate at date t. The level of interest rates is measured by the difference between the 10-year yield and the 3-month interest rate at date t.The holding period is one quarter. The returns are annualized. The dark (light) shaded area corresponds to the 90% (95%) confidence interval. Standard deviations are obtained by bootstrapping 10,000 samples of non-overlapping returns. Zero-coupon data are monthly, and the sample window is 4/1985–12/2015.

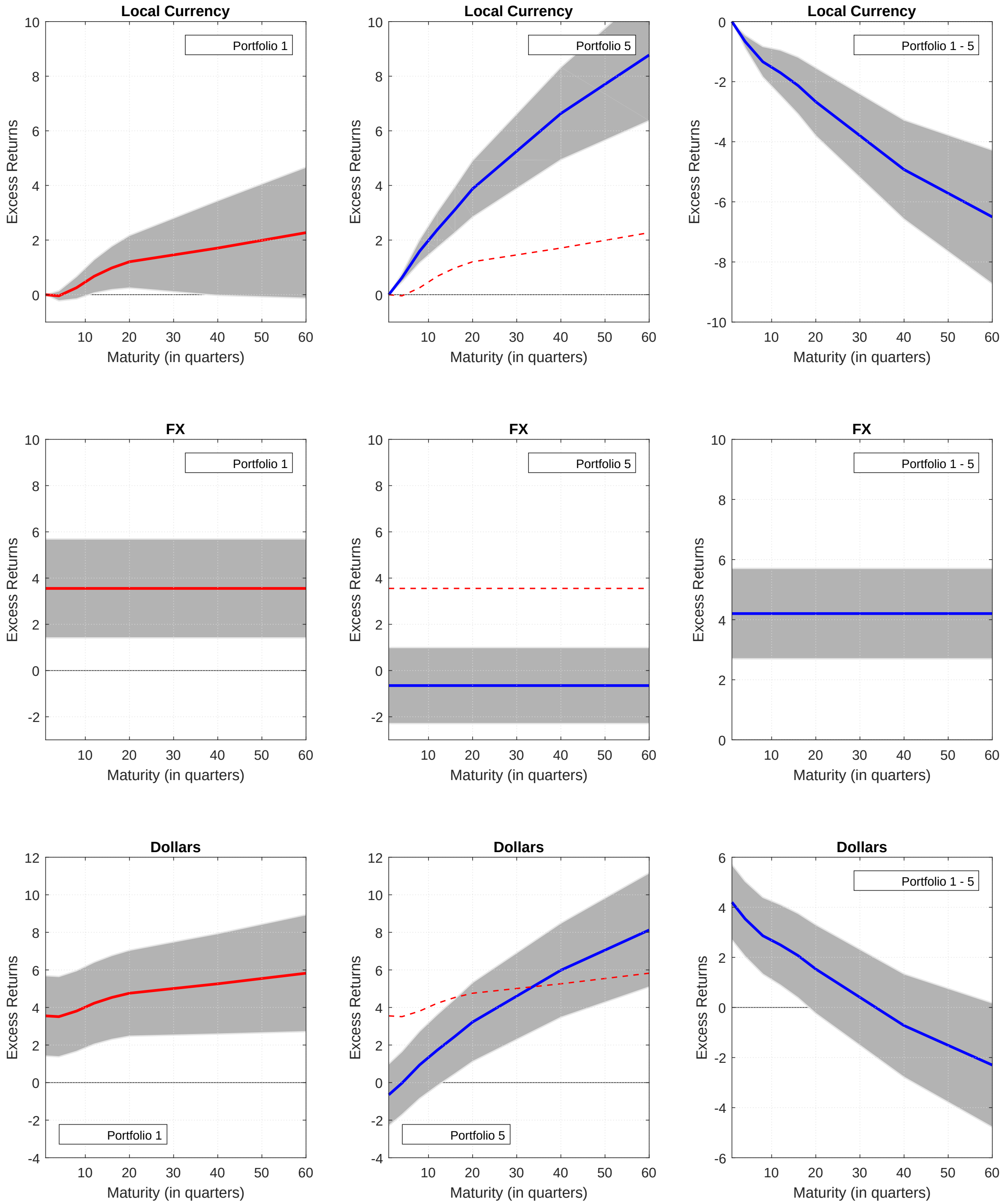

Figure A11: Dollar Bond Risk Premia Across Maturities: Extended Sample — The figure shows the local currency log excess returns in the top panel, and the dollar log excess returns in the bottom panel as a function of the bond maturities. The left panel focuses on Portfolio 1 (flat yield curve currencies) excess returns, while the middle panel reports Portfolio 5 (steep yield curve currencies) excess returns. The middle panels also report the Portfolio 1 excess returns in dashed lines for comparison. The right panel reports the difference. Data are monthly, from the zero-coupon dataset, and the sample window is 5/1987–12/2015. The unbalanced sample includes Australia, Austria, Belgium, Canada, the Czech Republic, Denmark, Finland, France, Germany, Hungary, Indonesia, Ireland, Italy, Japan, Malaysia, Mexico, the Netherlands, New Zealand, Norway, Poland, Portugal, Singapore, South Africa, Spain, Sweden, Switzerland, and the U.K. The countries are sorted by the slope of their yield curves into five portfolios. The slope of the yield curve is measured by the difference between the 10-year yield and the 3-month interest rate at date t. The holding period is one quarter. The returns are annualized. The shaded areas correspond to one standard deviation above and below each point estimate. Standard deviations are obtained by bootstrapping 10,000 samples of non-overlapping returns.

D Dynamic Term Structure Models

This section of the Appendix presents the details of pricing kernel decomposition for four classes of dynamic term structure models. Condition 1 is a diagnostic tool that can be applied to richer models. We apply it to several reduced-form term structure models, from the simple one-factor Vasicek (1977) and Cox, Ingersoll, and Ross (1985) models to their multi-factor versions. In order to save space, we summarize the restrictions implied by Condition 1 in Table A19.

D.1 Vasicek (1977)

In the Vasicek model, the log SDF evolves as:

$$-m_{t+1} = y_{1,t} + \frac{1}{2}\lambda^2\sigma^2 + \lambda\varepsilon_{t+1},$$

where $y_{1,t}$ denotes the short-term interest rate. It is affine in a single factor:

$$\begin{aligned} x_{t+1} &= \rho x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, \sigma^2\right) \\ y_{1,t} &= \delta + x_t. \end{aligned}$$

In this model, x_t is the level factor and ε_{t+1} are shocks to the level of the term structure. The Jensen term is there to ensure that $E_t\left(M_{t+1}\right) = \exp\left(-y_{1,t}\right)$. Bond prices are exponentially affine. For any maturity n, bond prices are equal to $P_t^{(n)} = \exp\left(-B_0^n - B_1^n x_t\right)$. The price of the one-period risk-free note ($n = 1$) is naturally:

$$P_t^{(1)} = \exp\left(-y_{1,t}\right) = \exp\left(-B_0^1 - B_1^1 x_t\right),$$

with $B_0^1 = \delta, B_1^1 = 1$, where the coefficients satisfy the following recursions:

$$\begin{aligned} B_0^n &= \delta + B_0^{n-1} - \frac{1}{2}\sigma^2(B_1^{n-1})^2 - \lambda B_1^{n-1}\sigma^2, \\ B_1^n &= 1 + B_1^{n-1}\rho. \end{aligned}$$

We first implement the Alvarez and Jermann (2005) approach. The temporary pricing component of the pricing kernel is:

$$\Lambda_t^{\mathbb{T}} = \lim_{n\to\infty} \frac{\beta^{t+n}}{P_t^n} = \lim_{n\to\infty} \beta^{t+n} e^{B_0^n + B_1^n x_t},$$

where the constant β is chosen in order to satisfy Assumption 1 in Alvarez and Jermann (2005):

$$0 < \lim_{n\to\infty} \frac{P_t^n}{\beta^n} < \infty.$$

The limit of $B_0^n - B_0^{n-1}$ is finite: $\lim_{n\to\infty} B_0^n - B_0^{n-1} = \delta - \frac{1}{2}\sigma^2(B_1^\infty)^2 - \lambda B_1^\infty \sigma^2$, where B_1^∞ is $1/(1-\rho)$. As a result, B_0^n grows at a linear rate in the limit. We choose the constant β to offset the growth in B_0^n as n becomes very large. Setting $\beta = e^{-\delta + \frac{1}{2}\sigma^2(B_1^\infty)^2 + \lambda B_1^\infty \sigma^2}$ guarantees that Assumption 1 in Alvarez and Jermann (2005) is satisfied. The temporary pricing component of the pricing kernel is thus equal to:

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = \beta e^{B_1^\infty (x_{t+1} - x_t)} = \beta e^{\frac{1}{1-\rho}(\rho - 1)x_t + \frac{1}{1-\rho}\varepsilon_{t+1}} = \beta e^{-x_t + \frac{1}{1-\rho}\varepsilon_{t+1}}.$$

The martingale component of the pricing kernel is then:

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = \beta^{-1} e^{x_t - \frac{1}{1-\rho}\varepsilon_{t+1} - \delta - x_t - \frac{1}{2}\lambda^2\sigma^2 - \lambda\varepsilon_{t+1}} = \beta^{-1} e^{-\delta - \frac{1}{2}\lambda^2\sigma^2 - (\frac{1}{1-\rho} + \lambda)\varepsilon_{t+1}}.$$

In the case of $\lambda = -B_1^\infty = -\frac{1}{1-\rho}$, the martingale component of the pricing kernel is constant and all the shocks that affect the pricing kernel are transitory.
The expected log excess return of an infinite maturity bond is then:

$$E_t[rx_{t+1}^{(\infty)}] = -\frac{1}{2}\sigma^2(B_1^\infty)^2 - \lambda B_1^\infty \sigma^2.$$

The first term is a Jensen term. The risk premium is constant and positive if λ is negative. The SDF is homoskedastic. The

Table A19: Condition 1: Dynamic Term Structure Model Scorecard

	Vasicek (1977)	Cox, Ingersoll, and Ross (1985)
Model	$-m_{t+1} = y_{1,t} + \frac{1}{2}\lambda^2\sigma^2 + \lambda\varepsilon_{t+1}$ $x_{t+1} = \rho x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, \sigma^2\right)$ $y_{1,t} = \delta + x_t$	$-m_{t+1} = \alpha + \chi z_t + \sqrt{\gamma z_t} u_{t+1},$ $z_{t+1} = (1-\phi)\theta + \phi z_t - \sigma\sqrt{z_t} u_{t+1},$
Condition 1	$\left(\frac{1}{1-\rho} + \lambda\right)\sigma^2 = \left(\frac{1}{1-\rho^*} + \lambda^*\right)\sigma^{*,2}$	$\left(\sqrt{\gamma} - B_1^\infty\sigma\right) z_t = \left(\sqrt{\gamma^*} - B_1^{\infty*}\sigma^*\right) z_t^*$
Restriction	$B_1^\infty = \text{-}\lambda = \frac{1}{1-\rho}$	$B_1^\infty = \frac{\chi}{1-\phi} = \sqrt{\gamma}/\sigma$

	Multi-factor Vasicek	Gaussian DTSM
Model	$-m_{t+1} = y_{1,t} + \frac{1}{2}\Lambda_t'\Sigma\Lambda_t + \Lambda_t'\varepsilon_{t+1}$ $x_{t+1} = \Gamma x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, \Sigma\right)$ $\Lambda_t = \Lambda_0 + \Lambda_1 x_t$ $y_{1,t} = \delta_0 + \delta_1' x_t$	$-m_{t+1} = y_{1,t} + \frac{1}{2}\Lambda' V(x_t)\Lambda + \Lambda' V(x_t)^{1/2}\varepsilon_{t+1}$ $x_{t+1} = \Gamma x_t + V(x_t)^{1/2}\varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, I\right)$ $V_{ii}(x_t) = \alpha_i + \beta_i' x_t$ $y_{1,t} = \delta_0 + \delta_1' x_t$
Condition 1	$(\Lambda_t' + B_1^{\infty\prime})\Sigma(\Lambda_t + B_1^\infty) = (\Lambda_t^{\star\prime} + B_1^{\star\infty\prime})\Sigma(\Lambda_t^\star + B_1^{\infty\star})$	$(\Lambda' + B_1^{\infty\prime})V(0)(\Lambda + B_1^\infty) = (\Lambda^{\star\prime} + B_1^{\star\infty\prime})V(0)(\Lambda^\star + B_1^{\infty\star})$ $(\Lambda' + B_1^{\infty\prime})V_x(\Lambda + B_1^\infty) = (\Lambda^{\star\prime} + B_1^{\star\infty\prime})V_x(\Lambda^\star + B_1^{\infty\star})$
Restriction	$B_1^\infty = -\Lambda_0; \Lambda_1 = 0$ $B_1^{\infty\prime} = (I-\Gamma)^{-1}\delta_1' = -\Lambda_0'$	$B_1^\infty = -\Lambda$ $B_1^\infty = \left(\delta_1' + \frac{1}{2}\Lambda' V_x \Lambda\right)(1-\Gamma)^{-1}$

The top panel considers single-factor models in which each country has its own factor. The bottom panel considers multi-factor versions in which all factors are common. Multi-factor Vasicek is a multi-factor extension of the Vasicek (1985) model. Gaussian Dynamic Term Structure Models (DTSM) are extensions of the Cox, Ingersoll, and Ross (1985) model. The details are in section D of the Online Appendix. The parameter restrictions for the Gaussian DTSM rules out all permanent shocks. In the Appendix, we discuss milder conditions that allow some shocks to have permanent, identical effects in both countries. In the multi-factor Vasicek model, we need to eliminate time-variation in the price of risk to impose Condition 1.

expected log currency excess return is therefore constant:

$$E_t[-\Delta s_{t+1}] + y_t^* - y_t = \frac{1}{2}Var_t(m_{t+1}) - \frac{1}{2}Var_t(m_{t+1}^*) = \frac{1}{2}\lambda\sigma^2 - \frac{1}{2}\lambda^*\sigma^{*2}.$$

When $\lambda = -B_1^\infty = -\frac{1}{1-\rho}$, the martingale component of the pricing kernel is constant and all the shocks that affect the pricing kernel are transitory. By using the expression for the bond risk premium in Equation (**??**), it is straightforward to verify that the expected log excess return of an infinite maturity bond is in this case:

$$E_t[rx_{t+1}^{(\infty)}] = \frac{1}{2}\sigma^2\lambda^2.$$

We start by examining the case in which each country has its own factor. We assume the foreign pricing kernel has the same structure, but it is driven by a different factor with different shocks:

$$\begin{aligned} -\log M_{t+1}^* &= y_{1,t}^* + \frac{1}{2}\lambda^{*2}\sigma^{*2} + \lambda^*\varepsilon_{t+1}^*, \\ x_{t+1}^* &= \rho x_t^* + \varepsilon_{t+1}^*, \quad \varepsilon_{t+1}^* \sim \mathcal{N}\left(0, \sigma^{*2}\right) \\ y_{1,t} &= \delta^* + x_t^*. \end{aligned}$$

Equation (**??**) shows that the expected log currency excess return is constant: $E_t[rx_{t+1}^{FX}] = \frac{1}{2}Var_t(m_{t+1}) - \frac{1}{2}Var_t(m_{t+1}^*) = \frac{1}{2}\lambda^2\sigma^2 - \frac{1}{2}\lambda^{2*}\sigma^{*2}$. In a Vasicek model with country-specific factors, the long bond uncovered return parity holds only if the model parameters satisfy the following restriction: $\lambda = -\frac{1}{1-\rho}$. Under these conditions, there is no martingale component in the pricing kernel and the foreign term premium on the long bond expressed in home currency is simply $E_t[rx_{t+1}^{(*,\infty)}] = \frac{1}{2}\lambda^2\sigma^2$. This expression equals the domestic term premium. The nominal exchange rate is stationary. [9]

D.2 Multi-Factor Vasicek Models

Under some conditions, the previous results can be extended to a more k-factor model. The standard $k-$factor essentially affine model in discrete time generalizes the Vasicek (1977) model to multiple risk factors. The log SDF is given by:

$$-\log M_{t+1} = y_{1,t} + \frac{1}{2}\Lambda_t'\Sigma\Lambda_t + \Lambda_t'\varepsilon_{t+1}$$

To keep the model affine, the law of motion of the risk-free rate and of the market price of risk are:

$$\begin{aligned} y_{1,t} &= \delta_0 + \delta_1' x_t, \\ \Lambda_t &= \Lambda_0 + \Lambda_1 x_t, \end{aligned}$$

where the state vector ($x_t \in R^k$) is:

$$x_{t+1} = \Gamma x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, \Sigma\right).$$

x_t is a $k \times 1$ vector, and so are ε_{t+1}, δ_1, Λ_t, and Λ_0, while Γ, Λ_1, and Σ are $k \times k$ matrices.[10]
We assume that the market price of risk is constant ($\Lambda_1 = \mathbf{0}$), so that we can define orthogonal temporary shocks. We decompose the shocks into two groups: the first $h < k$ shocks affect both the temporary and the permanent pricing kernel components and the last $k - h$ shocks are temporary.[11] The parameters of the temporary shocks satisfy $B_{1k-h}^{\infty\prime} = (I_{k-h} - \Gamma_{k-h})^{-1}\delta_{1k-h}' = -\Lambda_{0k-h}'$. This ensures that these shocks do not affect the permanent component of the pricing kernel.

[9] Alternatively, we can assume that the single state variable x_t is global. In this case, the countries trivially have the same pricing kernels.

[10] Note that if $k = 1$ and $\Lambda_1 = 0$, we are back to the Vasicek (1977) model with one factor and a constant market price of risk. The Vasicek (1977) model presented in the first section is a special case where $\Lambda_0 = \lambda$, $\delta_0' = \delta$, $\delta_0' = 1$ and $\Gamma = \rho$.

[11] A block-diagonal matrix whose blocks are invertible is invertible, and its inverse is a block diagonal matrix (with the inverse of each block on the diagonal). Therefore, if Γ is block-diagonal and $(I - \Gamma)$ is invertible, we can decompose the shocks as described

Now we assume that x_t is a global state variable:

$$
\begin{aligned}
-\log M^*_{t+1} &= y^*_{1,t} + \frac{1}{2}\Lambda^{*\prime}_t \Sigma \Lambda^*_t + \Lambda^{*\prime}_t \varepsilon_{t+1}, \\
y_{1,t} &= \delta^*_0 + \delta^{*\prime}_1 x_t, \\
\Lambda^*_t &= \Lambda^*_0, \\
x_{t+1} &= \Gamma x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}(0, \Sigma).
\end{aligned}
$$

In a multi-factor Vasicek model with global factors and constant risk prices, long bond uncovered return parity obtains only if countries share the same Λ_h and δ_{1h}, which govern exposure to the permanent, global shocks.
This condition eliminates any differences in permanent risk exposure across countries.[12] The nominal exchange rate has no permanent component $\left(\frac{S^{\mathbb{P}}_t}{S^{\mathbb{P}}_{t+1}} = 1\right)$. From equation (**??**) , the expected log currency excess return is equal to:

$$
E_t[rx^{FX}_{t+1}] = \frac{1}{2}Var_t(m_{t+1}) - \frac{1}{2}Var_t(m^*_{t+1}) = \frac{1}{2}\Lambda_0'\Sigma\Lambda_0 - \frac{1}{2}\Lambda^{*\prime}_0\Sigma\Lambda^*_0.
$$

Non-zero currency risk premia will be only due to variation in the exposure to transitory shocks (Λ^*_{0k-h}).

D.3 Cox, Ingersoll, and Ross (1985) Model

The Cox, Ingersoll, and Ross (1985) model (denoted CIR) is defined by the following two equations:

$$
\begin{aligned}
-\log M_{t+1} &= \alpha + \chi z_t + \sqrt{\gamma z_t} u_{t+1}, \\
z_{t+1} &= (1-\phi)\theta + \phi z_t - \sigma\sqrt{z_t} u_{t+1},
\end{aligned} \tag{14}
$$

where M denotes the stochastic discount factor. In this model, log bond prices are affine in the state variable z: $p^{(n)}_t = -B^n_0 - B^n_1 z_t$. The price of a one period-bond is: $P^{(1)} = E_t(M_{t+1}) = e^{-\alpha-(\chi-\frac{1}{2}\gamma)z_t}$. Bond prices are defined recursively by the Euler equation: $P^{(n)}_t = E_t(M_{t+1}P^{(n-1)}_{t+1})$. Thus the bond price coefficients evolve according to the following second-order difference equations:

$$
\begin{aligned}
B^n_0 &= \alpha + B^{n-1}_0 + B^{n-1}_1(1-\phi)\theta, \\
B^n_1 &= \chi - \frac{1}{2}\gamma + B^{n-1}_1\phi - \frac{1}{2}\left(B^{n-1}_1\right)^2\sigma^2 + \sigma\sqrt{\gamma}B^{n-1}_1.
\end{aligned} \tag{15}
$$

We first implement the Alvarez and Jermann (2005) approach. The temporary pricing component of the pricing kernel is:

$$
\Lambda^{\mathbb{T}}_t = \lim_{n\to\infty} \frac{\beta^{t+n}}{P^{(n)}_t} = \lim_{n\to\infty} \beta^{t+n} e^{B^n_0 + B^n_1 z_t},
$$

where the constant β is chosen in order to satisfy Assumption 1 in Alvarez and Jermann (2005):

$$
0 < \lim_{n\to\infty} \frac{P^{(n)}_t}{\beta^n} < \infty.
$$

The limit of $B^n_0 - B^{n-1}_0$ is finite: $\lim_{n\to\infty} B^n_0 - B^{n-1}_0 = \alpha + B^\infty_1(1-\phi)\theta$, where B^∞_1 is defined implicitly in a second-order equation above. As a result, B^n_0 grows at a linear rate in the limit. We choose the constant β to offset the growth in B^n_0 as n becomes very large. Setting $\beta = e^{-\alpha - B^\infty_1(1-\phi)\theta}$ guarantees that Assumption 1 in Alvarez and Jermann (2005) is satisfied. The temporary pricing component of the pricing kernel is thus equal to:

$$
\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t} = \beta e^{B^\infty_1(z_{t+1}-z_t)} = \beta e^{B^\infty_1\left[(\phi-1)(z_t-\theta) - \sigma\sqrt{z_t}u_{t+1}\right]}.
$$

As a result, the martingale component of the pricing kernel is then:

$$
\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\right)^{-1} = \beta^{-1} e^{-\alpha - \chi z_t - \sqrt{\gamma z_t}u_{t+1}} e^{-B^\infty_1\left[(\phi-1)(z_t-\theta) - \sigma\sqrt{z_t}u_{t+1}\right]}. \tag{16}
$$

[12] The terms δ_1' and $\delta^{*\prime}_{1h}$ do not appear in the single-factor Vasicek (1977) model of the first section because that single-factor model assumes $\delta_1 = \delta^*_{1h} = 1$.

The expected log excess return is thus given by:

$$E_t[rx_{t+1}^{(n)}] = [-\frac{1}{2}\left(B_1^{n-1}\right)^2\sigma^2 + \sigma\sqrt{\gamma}B_1^{n-1}]z_t.$$

The expected log excess return of an infinite maturity bond is then:

$$\begin{aligned} E_t[rx_{t+1}^{(\infty)}] &= [-\frac{1}{2}\left(B_1^{\infty}\right)^2\sigma^2 + \sigma\sqrt{\gamma}B_1^{\infty}]z_t, \\ &= [B_1^{\infty}(1-\phi) - \chi + \frac{1}{2}\gamma]z_t. \end{aligned}$$

The $-\frac{1}{2}\left(B_1^{\infty}\right)^2\sigma^2$ is a Jensen term. The term premium is driven by $\sigma\sqrt{\gamma}B_1^{\infty}z_t$, where B_1^{∞} is defined implicitly in the second order equation $B_1^{\infty} = \chi - \frac{1}{2}\gamma + B_1^{\infty}\phi - \frac{1}{2}\left(B_1^{\infty}\right)^2\sigma^2 + \sigma\sqrt{\gamma}B_1^{\infty}$.
Consider the special case of $B_1^{\infty}(1-\phi) = \chi$. In this case, the expected term premium is simply $E_t[rx_{t+1}^{(\infty)}] = \frac{1}{2}\gamma z_t$, which is equal to one-half of the variance of the log stochastic discount factor.
Suppose that the foreign pricing kernel is specified as in Equation (14) with the same parameters. However, the foreign country has its own factor z^*. As a result, the difference between the domestic and foreign log term premia is equal to the log currency risk premium, which is given by $E_t[rx_{t+1}^{FX}] = \frac{1}{2}\gamma(z_t - z_t^*)$. In other words, the expected foreign log holding period return on a foreign long bond converted into U.S. dollars is equal to the U.S. term premium: $E_t[rx_{t+1}^{(\infty),*}] + E_t[rx_{t+1}^{FX}] = \frac{1}{2}\gamma z_t$.
This special case corresponds to the absence of permanent shocks to the pricing kernel: when $B_1^{\infty}(1-\phi) = \chi$, the permanent component of the stochastic discount factor is constant. To see this result, let us go back to the implicit definition of B_1^{∞} in Equation (16):

$$\begin{aligned} 0 &= \frac{1}{2}\left(B_1^{\infty}\right)^2\sigma^2 + (1-\phi-\sigma\sqrt{\gamma})B_1^{\infty} - \chi + \frac{1}{2}\gamma, \\ 0 &= \frac{1}{2}\left(B_1^{\infty}\right)^2\sigma^2 - \sigma\sqrt{\gamma}B_1^{\infty} + \frac{1}{2}\gamma, \\ 0 &= \left(\sigma B_1^{\infty} - \sqrt{\gamma}\right)^2. \end{aligned}$$

In this special case, $B_1^{\infty} = \sqrt{\gamma}/\sigma$. Using this result in Equation (16), the permanent component of the pricing kernel reduces to:

$$\frac{M_{t+1}^{\mathbb{P}}}{M_t^{\mathbb{P}}} = \frac{M_{t+1}}{M_t}\left(\frac{M_{t+1}^{\mathbb{T}}}{M_t^{\mathbb{T}}}\right)^{-1} = \beta^{-1}e^{-\alpha-\chi z_t - \sqrt{\gamma z_t}u_{t+1}}e^{-B_1^{\infty}\left[(\phi-1)(z_t-\theta)-\sigma\sqrt{z_t}u_{t+1}\right]} = \beta^{-1}e^{-\alpha-\chi\theta},$$

which is a constant.[13]

The expected bond excess return is:

$$E_t(r_{t+1}^{(n)}) - r_t^f + \frac{1}{2}var_t(r_{t+1}^{(n)}) = -cov_t(p_{t+1}^{(n-1)}, m_{t+1}) = B_1^{n-1}\sigma\sqrt{\gamma}z_t. \tag{17}$$

Recall that the risk-free rate is:

$$r_t^f = -E_t(\log M_{t+1}) - \frac{1}{2}Var_t(\log M_{t+1}) = \alpha + (\chi - \frac{1}{2}\gamma)z_t. \tag{18}$$

In order to replicate the U.I.P. puzzle, risk-free rates must be low when stochastic discount factors are volatile, implying that $\chi < \frac{1}{2}\gamma$: the risk-free rate is decreasing in the state variable. Since $\chi < \frac{1}{2}\gamma$, Equation (16) implies recursively that all B_1^n coefficients are negative. The bond risk premium is thus decreasing in the state variable. The slope of the yield curve is $y_t^{(n)} - r_t^f = -p_t^{(n)}/n - r_t^f = B_0^n/n + B_1^n/n\ z_t - \alpha - (\chi - \frac{1}{2}\gamma)z_t$. Its cyclicality is not immediately obvious from this expression, but we verified that in our simulations the slope and the level of the yield curve move in opposite directions. This property appears clearly when considering infinite-maturity bonds. In the limit of long-term bonds, the slope of the yield curve is $y_t^{(\infty)} - r_t^f = -B_1^{\infty}(1-\phi)\theta - (\chi - \frac{1}{2}\gamma)z_t$. As the infinite maturity yield is constant, the infinite maturity slope moves in opposite direction as the risk-free rate. Bringing

[13] Alternatively, we assume that all the shocks are global and that z_t is a global state variable (and thus $\sigma = \sigma^*$, $\phi = \phi^*$, $\theta = \theta^*$). Condition 1 requires that:

$$\sqrt{\gamma} + B_1^{\infty}\sigma = \sqrt{\gamma^*} + B_1^{\infty *}\sigma$$

Note that B_1^{∞} depends on χ and γ, as well as on the global parameters ϕ and σ. The two countries have perfectly correlated pricing kernels.

everything together, note that when the state variable is low, the risk-free rate is high, the slope of the yield curve is low, and the bond risk premium (Equation (17)) is high. In the data, the risk-free rate and slope of the yield move in opposite directions across countries, but high-slope portfolios correspond to high bond risk premia. This simple one-factor model cannot reproduce our empirical evidence.

We now consider the two-country version of the Cox, Ingersoll, and Ross (1985) model. It is defined by the following law of motions for the SDFs:

$$\begin{aligned} -\log \frac{\Lambda_{t+1}}{\Lambda_t} &= \alpha + \chi z_t + \sqrt{\gamma z_t} u_{t+1}, \\ z_{t+1} &= (1-\phi)\theta + \phi z_t - \sigma\sqrt{z_t} u_{t+1}, \\ -\log \frac{\Lambda^*_{t+1}}{\Lambda^*_t} &= \alpha^* + \chi^* z^*_t + \sqrt{\gamma^* z^*_t} u^*_{t+1}, \\ z^*_{t+1} &= (1-\phi^*)\theta^* + \phi^* z^*_t - \sigma^*\sqrt{z^*_t} u^*_{t+1}, \end{aligned}$$

where z_t and z^*_t are the two state variables that govern the volatilities of the normal shocks u_{t+1} and u^*_{t+1}. In this two-country model, Condition 1 requires that

$$\left(\sqrt{\gamma} - B_1^{\infty}\sigma\right) z_t = \left(\sqrt{\gamma^*} - B_1^{\infty*}\sigma^*\right) z^*_t.$$

There are two ways to ensure that this condition is satisfied, depending on whether the shocks are either country-specific or common.

First, let us consider a model with only country-specific shocks and factors. Let us assume that these countries share all of the parameters. Since z_t and z^*_t will differ, a necessary and sufficient condition is that $B_1^{\infty} = \sqrt{\gamma}/\sigma$, and $B_1^{\infty,*} = \sqrt{\gamma^*}/\sigma^*$. In this case, there are no permanent shocks to the pricing kernel. Long bond prices absorb the full, cumulative effect of the shock the pricing kernel. To see why, note that in this case $B_1^{\infty} = \chi/(1-\phi)$. The log currency risk premium is given by $E_t[rx^{FX}_{t+1}] = \frac{1}{2}\gamma(z_t - z^*_t)$ and the expected term premium is simply $E_t[rx^{(\infty)}_{t+1}] = \frac{1}{2}\gamma z_t$. The expected foreign log holding period return on a foreign long bond converted into U.S. dollars is equal to the U.S. term premium: $E_t[rx^{(\infty),*}_{t+1}] + E_t[rx^{FX}_{t+1}] = \frac{1}{2}\gamma z_t$. In a two-country Cox, Ingersoll, and Ross (1985) model with country-specific shocks, Condition 1 implies some restrictions on the model parameters, and more crucially, the absence of permanent shocks in the SDFs and thus in exchange rates: exchange rates are stationary in levels. The case of country-specific shocks, however, is not the most interesting as such shocks can be diversified away.

Second, let us consider a model with common shocks and common factors: $z_t = z^*_t$ is a global state variable. In this case, the two countries share the parameters $\sigma = \sigma^*$, $\phi = \phi^*$, $\theta = \theta^*$ which govern the dynamics of z_t and z^*_t. Condition 1 then requires that $\sqrt{\gamma} + B_1^{\infty}\sigma = \sqrt{\gamma^*} + B_1^{\infty*}\sigma$. Note that B_1^{∞} depends on χ and γ, as well as on the global parameters ϕ and σ. Hence, we also need $\gamma = \gamma^*$ and $\chi = \chi^*$. In this case, Condition 1 requires that both countries have the same pricing kernel. This case illustrates the tension between the carry trade at the short and the long end of the yield curve: in order to replicate the carry trade on Treasury bills, the two-country Cox, Ingersoll, and Ross (1985) model needs to feature heterogeneous exposure to common shocks; yet, in order to replicate the absence of carry trade returns on long term bonds, this model needs to satisfy Condition 1 that prohibits such heterogeneous exposure to common shocks.

Long-Run U.I.P

Result 2. *In the two-country CIR model, the transitory component of the exchange rate is given by:*

$$s^{\mathbb{T}}_t = s_0 + \left(B_1^{\infty}(z_t - z_0) - B_1^{\infty,*}(z^*_t - z^*_0)\right).$$

When the pricing kernel is not subject to permanent shocks, $B_1^{\infty} = \frac{\sqrt{\gamma}}{\sigma} = \frac{\chi}{1-\phi}$, the exchange rate is stationary and hence $s_t = s^{\mathbb{T}}_t$:

$$s_t = s_0 + \left(\frac{\chi}{1-\phi}(z_t - z_0) - \frac{\chi^*}{1-\phi^*}(z^*_t - z^*_0)\right).$$

The expected rate of depreciation is

$$\lim_{k\to\infty} E_t[\Delta s_{t\to t+k}] = \frac{\chi}{1-\phi} z_t - \frac{\chi^*}{1-\phi^*} z^*_t = -\lim_{k\to\infty} k\left(y^{(k)}_t - y^{(k),*}_t\right).$$

Long-run U.I.P. holds for all transitory shocks the pricing kernel: the long-run response of the exchange rate to transitory innovations equals the response of the long rate today, and hence this response can be read off the yield curve.

Our analysis sheds light on the recent empirical findings of Engel (2016), Valchev (2016), and Dahlquist and Penasse (2016). Engel (2016) finds that an increase in the short-term interest rate initially cause exchange rates to appreciate, but they subsequently depreciate on average. Because the risk premia on long bonds are equalized, shocks to the quantity or price of risk (e.g., an increase in risk aversion) cannot have long-run effects; long-run U.I.P. holds for these shocks. As a result, our preference-free condition constrains the long-run response of exchange rates to transitory shocks to be equal to the instantaneous response of long-term interest rates. For example, countries which have experienced an adverse transitory shock, with higher than average long-term interest rates, always have stronger currencies (the level of the exchange rate is temporarily high), because their exchange rates are expected to revert back to the mean and depreciate in the long run by the long run interest rate difference (see Dornbusch, 1976; Frankel, 1979, for early contributions on the relation between the level of the exchange rate and interest rates). Thus, an increase in a country's short and long interest rates which causes an appreciation in the short run has to be more than offset by future depreciations.

To develop some intuition, consider a symmetric version of the two-country CIR model in which the 2 countries share all of the parameters. The restrictions $B_1^\infty = \frac{\sqrt{\gamma}}{\sigma} = \frac{\chi}{1-\phi}$ have a natural interpretation as restrictions on the long-run loadings of the exchange rate on the risk factors: $\sum_{i=1}^\infty E_t[\Delta s_{t+i}] = \sum_{i=1}^\infty E_t[m_{t+i} - m^*_{t+i}] = \sum_{i=1}^\infty \phi^{i-1}\chi(z^*_t - z_t)$. As can easily be verified, these two restrictions imply that the long-run loading of the exchange rate on the factors equals the loading of long-term interest rates:

$$\lim_{k\to\infty} E_t[\Delta s_{t\to t+k}] = \frac{\chi}{(1-\phi)}(z^*_t - z_t) = \lim_{k\to\infty} k\left(y_t^{(k),*} - y_t^{(k)}\right).$$

Hence, in the context of this model, our restrictions enforce long-run U.I.P. An increase in risk abroad causes the long rates to go up abroad and the foreign exchange rate to depreciate in the long run, but given these long-run restrictions, the initial expected exchange rate impact has to have the same sign ($\chi > 0$), thus violating the empirical evidence, as we explain below.

Our preference-free conditions constrains the sum of slope regression coefficients in a regression of future exchange rate changes Δs_{t+i} on the current interest rate spread $r_t^{f,\$,*} - r_t^{f,\$}$ to be equal to the response of long-term interest rates. Engel (2016), Valchev (2016), and Dahlquist and Penasse (2016) study these slope coefficients and find that they switch signs with the horizon i: an increase in the short-term interest rate initially cause exchange rates to appreciate, but they subsequently depreciate on average.

Result 3. *In the symmetric two-country CIR model without permanent shocks* $B_1^\infty = \frac{\sqrt{\gamma}}{\sigma} = \frac{\chi}{1-\phi}$*, the slope coefficients in a regression of* Δs_{t+i} *on the* $r_t^{f,\$,*} - r_t^{f,\$}$*, given by* $\frac{\phi^{i-1}\chi}{\chi - \frac{1}{2}\gamma}$ *decline geometrically as* i *increases, and their infinite sum equals* $\frac{B_1^\infty}{\chi - \frac{1}{2}\gamma}$.

When $(\chi - \frac{1}{2}\gamma) < 0$, the model can match the short-run forward premium puzzle: when the foreign short rate increases, the currency subsequently appreciates, but it continues to appreciate as long rates decline abroad. As a result, this model cannot match the sign switch in these regression coefficients. A richer version of the factor model with multiple country-specific risk factors can generate richer dynamics.Consider the same model with two country-specific risk factors. The long-run impulse responses of the exchange rate to short-term interest rate shocks is driven by:

$$\sum_{i=1}^\infty E_t[\Delta s_{t+i}] = \sum_{i=1}^\infty E_t[m_{t+i} - m^*_{t+i}] = \sum_{i=1}^\infty \left[\phi_1^{i-1}\chi_1(z_t^{1,*} - z_t^1) + \phi_2^{i-1}\chi_2(z_t^{2,*} - z_t^2)\right].$$

The slope coefficients in a regression of future exchange rate changes on the current interest rate spread $r_t^{f,\$,*} - r_t^{f,\$}$ are given by

$$E_t\Delta s_{t+i} = \frac{\phi_1^{i-1}\chi_1(\chi_1 - \frac{1}{2}\gamma_1) + \phi_2^{i-1}\chi_2(\chi_2 - \frac{1}{2}\gamma_2)}{(\chi_1 - \frac{1}{2}\gamma_1)^2 + (\chi_2 - \frac{1}{2}\gamma_2)^2}\left(r_t^{f,\$,*} - r_t^{f,\$}\right).$$

These coefficients can switch signs as we increase the maturity i if the risk factors have sufficiently heterogeneous persistence (ϕ_1, ϕ_2), and provided that $(\chi_1 - \frac{1}{2}\gamma_1)$ and $(\chi_2 - \frac{1}{2}\gamma_2)$ have opposite signs.

D.4 Gaussian Dynamic Term Structure Models

The $k-$factor heteroskedastic Gaussian Dynamic Term Structure Model (DTSM) generalizes the CIR model. When market prices of risk are constant, the log SDF is given by:

$$\begin{aligned}
-m_{t+1} &= y_{1,t} + \frac{1}{2}\Lambda' V(x_t)\Lambda + \Lambda' V(x_t)^{1/2}\varepsilon_{t+1}, \\
x_{t+1} &= \Gamma x_t + V(x_t)^{1/2}\varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}(0, I), \\
y_{1,t} &= \delta_0 + \delta_1' x_t,
\end{aligned}$$

where $V(x)$ is a diagonal matrix with entries $V_{ii}(x_t) = \alpha_i + \beta_i' x_t$. To be clear, x_t is a $k \times 1$ vector, and so are ε_{t+1}, Λ, δ_1, and β_i. But Γ and V are $k \times k$ matrices. A restricted version of the model would impose that β_i is a scalar and $V_{ii}(x_t) = \alpha_i + \beta_i x_{it}$— this

is equivalent to assuming that the price of shock i only depends on the state variable i.
The price of a one period-bond is:

$$P_t^{(1)} = E_t(M_{t+1}) = e^{-\delta_0 - \delta_1' x_t}.$$

For any maturity n, bond prices are exponentially affine, $P_t^{(n)} = \exp\left(-B_0^n - B_1^{n\prime} x_t\right)$. Note that B_0^n is a scalar, while B_1^n is a $k \times 1$ vector. The one period-bond corresponds to $B_0^1 = \delta_0$, $B_1' = \delta_1'$, and the bond price coefficients satisfy the following difference equation:

$$\begin{aligned} B_0^n &= \delta_0 + B_0^{n-1} - \frac{1}{2} B_1^{n-1\prime} V(0) B_1^{n-1} - \Lambda' V(0) B_1^{n-1}, \\ B_1^{n\prime} &= \delta_1' + B_1^{n-1\prime}\Gamma - \frac{1}{2} B_1^{n-1\prime} V_x B_1^{n-1} - \Lambda' V_x B_1^{n-1}, \end{aligned}$$

where V_x denotes all the diagonal slope coefficients β_i of the V matrix.
The CIR model studied in the previous pages is a special case of this model. It imposes that $k = 1$, $\sigma = -\sqrt{\beta}$, and $\Lambda = -\frac{1}{\sigma}\sqrt{\gamma}$. Note that the CIR model has no constant term in the square root component of the log SDF, but that does not imply $V(0) = 0$ here as the CIR model assumes that the state variable has a non-zero mean (while it is zero here).
From there, we can define the Alvarez and Jermann (2005) pricing kernel components as for the Vasicek model. The limit of $B_0^n - B_0^{n-1}$ is finite: $\lim_{n\to\infty} B_0^n - B_0^{n-1} = \delta_0 - \frac{1}{2} B_1^{\infty\prime} V(0) B_1^{\infty} - \Lambda_0' V(0) B_1^{\infty}$, where $B_1^{\infty\prime}$ is the solution to the second-order equation above. As a result, B_0^n grows at a linear rate in the limit. We choose the constant β to offset the growth in B_0^n as n becomes very large. Setting $\beta = e^{-\delta_0 + \frac{1}{2} B_1^{\infty\prime} V(0) B_1^{\infty} + \Lambda' V(0) B_1^{\infty}}$ guarantees that Assumption 1 in Alvarez and Jermann (2005) is satisfied. The temporary pricing component of the pricing kernel is thus equal to:

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = \beta e^{B_1^{\infty\prime}(x_{t+1} - x_t)} = \beta e^{B_1^{\infty\prime}(\Gamma - 1)x_t + B_1^{\infty\prime} V(x_t)^{1/2} \varepsilon_{t+1}}.$$

The martingale component of the pricing kernel is then:

$$\begin{aligned} \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t} \left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} &= \beta^{-1} e^{-B_1^{\infty}(\Gamma-1)x_t - B_1^{\infty\prime} V(x_t)^{1/2}\varepsilon_{t+1} - y_{1,t} - \frac{1}{2}\Lambda' V(x_t)\Lambda_t - \Lambda' V(x_t)^{1/2}\varepsilon_{t+1}} \\ &= \beta^{-1} e^{-B_1^{\infty}(\Gamma-1)x_t - \delta_0 - \delta_1' x_t - \frac{1}{2}\Lambda' V(x_t)\Lambda - (\Lambda' + B_1^{\infty\prime}) V(x_t)^{1/2}\varepsilon_{t+1}}. \end{aligned}$$

For the martingale component to be constant, we need that $\Lambda' = -B_1^{\infty\prime}$ and $B_1^{\infty}(\Gamma - 1) + \delta_1' + \frac{1}{2}\Lambda' V_x \Lambda = 0$. Note that the second condition is automatically satisfied if the first one holds: this result comes from the implicit value of $B_1^{\infty\prime}$ implied by the law of motion of B_1. As a result, the martingale component is constant as soon as $\Lambda = -B_1^{\infty}$.
The expected log holding period excess return is:

$$E_t[rx_{t+1}^{(n)}] = -\delta_0 + \left(-B_1^{n-1\prime}\Gamma + B_1^{n\prime} - \delta_1'\right) x_t.$$

The term premium on an infinite-maturity bond is therefore:

$$E_t[rx_{t+1}^{(\infty)}] = -\delta_0 + \left((1 - \Gamma) B_1^{\infty\prime} - \delta_1'\right) x_t.$$

The expected log currency excess return is equal to:

$$E_t[-\Delta s_{t+1}] + y_t^* - y_t = \frac{1}{2} Var_t(m_{t+1}) - \frac{1}{2} Var_t(m_{t+1}^*) = \frac{1}{2}\Lambda' V(x_t)\Lambda - \frac{1}{2}\Lambda^{*\prime} V(x_t^*)\Lambda^*.$$

We assume that all the shocks are global and that x_t is a global state variable ($\Gamma = \Gamma^*$ and $V = V^*$, no country-specific parameters in the V matrix— cross-country differences will appear in the vectors Λ). Let us decompose the shocks into two groups: the first $h < k$ shocks affect both the temporary and the permanent pricing kernel components and the last $k - h$ shocks are temporary. Temporary shocks are such that $\Lambda_{k-h} = -B_{1,k-h}^{\infty}$ (i.e., they do not affect the value of the permanent component of the pricing kernel).
The risk premia on the domestic and foreign infinite-maturity bonds (once expressed in the same currency) will be the same provided that the entropy of the domestic and foreign permanent components is the same:

$$\begin{aligned} (\Lambda_h' + B_{1h}^{\infty\prime}) V(0) (\Lambda_h + B_{1h}^{\infty}) &= (\Lambda_h^{*\prime} + B_{1h}^{*\infty\prime}) V(0) (\Lambda_h^* + B_{1h}^{\infty *}), \\ (\Lambda_h' + B_{1h}^{\infty\prime}) V_x (\Lambda_h + B_{1h}^{\infty}) &= (\Lambda_h^{*\prime} + B_{1h}^{*\infty\prime}) V_x (\Lambda_h^* + B_{1h}^{\infty *}). \end{aligned}$$

To compare these conditions to the results obtained in the one-factor CIR model, recall that $\sigma^{CIR} = -\sqrt{\beta}$, and $\Lambda = -\frac{1}{\sigma^{CIR}}\sqrt{\gamma^{CIR}}$.

Differences in Λ_h in the k-factor model are equivalent to differences in γ in the CIR model: in both cases, they correspond to different loadings of the log SDF on the "permanent" shocks. As in the CIT model, differences in term premia can also come form differences in the sensitivity of infinite-maturity bond prices to the global "permanent" state variable ($B_{1h}^{\infty\prime}$), which can be traced back to differences in the sensitivity of the risk-free rate to the "permanent" state variable (i.e., different δ_1 parameters).

Let us start with the special case of no permanent innovations: $h = 0$, the martingale component is constant. Two conditions need to be satisfied for the martingale component to be constant: $\Lambda' = -B_1^{\infty\prime}$ and $B_1^{\infty}(\Gamma - 1) + \delta_1' + \frac{1}{2}\Lambda' V_x \Lambda = 0$. The second condition imposes that the cumulative impact on the pricing kernel of an innovation today given by $\left(\delta_1' + \frac{1}{2}\Lambda' V_x \Lambda\right)(1-\Gamma)^{-1}$ equals the instantaneous impact of the innovation on the long bond price. The second condition is automatically satisfied if the first one holds, as can be verified from the implicit value of $B_1^{\infty\prime}$ implied by the law of motion of B_1. As a result, the martingale component is constant as soon as $\Lambda = -B_1^{\infty}$.

As implied by Equation (**??**), the term premium on an infinite-maturity zero coupon bond is:

$$E_t[rx_{t+1}^{(\infty)}] = -\delta_0 + \left((1-\Gamma)B_1^{\infty\prime} - \delta_1'\right)x_t. \tag{19}$$

In the absence of permanent shocks, when $\Lambda = -B_1^{\infty}$, this log bond risk premium equals half of the stochastic discount factor variance $E_t[rx_{t+1}^{(\infty)}] = \frac{1}{2}\Lambda' V(x_t)\Lambda$; it attains the upper bound on log risk premia. Consistent with the result in Equation (**??**), the expected log currency excess return is equal to:

$$E_t\left[rx_{t+1}^{FX}\right] = \frac{1}{2}\Lambda' V(x_t)\Lambda - \frac{1}{2}\Lambda^{*\prime} V(x_t)\Lambda^*. \tag{20}$$

Differences in the market prices of risk Λ imply non-zero currency risk premia. Adding the previous two expressions in Equations (19) and (20), we obtain the foreign bond risk premium in dollars. The foreign bond risk premium in dollars equals the domestic bond premium in the absence of permanent shocks: $E_t\left[rx_{t+1}^{(\infty),*}\right] + E_t\left[rx_{t+1}^{FX}\right] = \frac{1}{2}\Lambda' V(x_t)\Lambda$.

In general, there is a spread between dollar returns on domestic and foreign bonds. We describe a general condition for long-run uncovered return parity in the presence of permanent shocks. In a GDTSM with global factors, the long bond uncovered return parity condition holds only if the countries' SDFs share the parameters $\Lambda_h = \Lambda_h^*$ and $\delta_{1h} = \delta_{1h}^*$, which govern exposure to the permanent global shocks.

The log risk premia on the domestic and foreign infinite-maturity bonds (once expressed in the same currency) are identical provided that the entropies of the domestic and foreign permanent components are the same:

$$\begin{aligned}
(\Lambda_h' + B_{1h}^{\infty\prime})V(0)(\Lambda_h + B_{1h}^{\infty}) &= (\Lambda_h^{*\prime} + B_{1h}^{*\infty\prime})V(0)(\Lambda_h^* + B_{1h}^{\infty *}), \\
(\Lambda_h' + B_{1h}^{\infty\prime})V_x(\Lambda_h + B_{1h}^{\infty}) &= (\Lambda_h^{*\prime} + B_{1h}^{*\infty\prime})V_x(\Lambda_h^* + B_{1h}^{\infty *}).
\end{aligned}$$

These conditions are satisfied if that these countries share $\Lambda_h = \Lambda_h^*$ and $\delta_{1h} = \delta_{1h}^*$ which govern exposure to the global shocks. In this case, the expected log currency excess return is driven entirely by differences between the exposures to transitory shocks: Λ_{k-h} and Λ_{k-h}^*. If there are only permanent shocks ($h = k$), then the currency risk premium is zero.[14]

D.5 An Example: A Reduced-Form Factor Model

This section provides details on the properties of bond and currency premia in the Lustig, Roussanov, and Verdelhan (2014) model. We now turn to a flexible N-country, reduced-form model that can both replicate the deviations from U.I.P. and generate large currency carry trade returns on currency portfolios. To replicate the portfolio evidence, as Lustig, Roussanov, and Verdelhan (2011) show, no arbitrage models need to incorporate global shocks to the SDFs along with country heterogeneity in the exposure to those shocks. Following Lustig, Roussanov, and Verdelhan (2014), we consider a world with N countries and currencies in a setup inspired by classic term structure models.[15] In the model, the risk prices associated with country-specific shocks depend only on country-specific factors, but the risk prices of world shocks depend on world and country-specific factors. To describe these risk prices, the authors introduce a common state variable z_t^w, shared by all countries, and a country-specific state variable z_t^i. The

[14]To compare these conditions to the results obtained in the CIR model, recall that we have constrained the parameters in the CIR model such that: $\sigma^{CIR} = -\sqrt{\beta}$, and $\Lambda = -\frac{1}{\sigma^{CIR}}\sqrt{\gamma^{CIR}}$. Differences in Λ_h in the k-factor model are equivalent to differences in γ in the CIR model: in both cases, they correspond to different loadings of the log pricing kernel on the "permanent" shocks. Differences in term premia can also come form differences in the sensitivity of the risk-free rate to the permanent state variable (i.e., different δ_1 parameters). These correspond to differences in χ in the CIR model.

[15]In the Online Appendix, we cover a wide range of term structure models, from the seminal Vasicek (1977) model to the classic Cox, Ingersoll, and Ross (1985) model and to the most recent, multi-factor dynamic term structure models. To save space, we focus here on their most recent international finance version, illustrated in Lustig, Roussanov, and Verdelhan (2014).

country-specific and world state variables follow autoregressive square-root processes:

$$\begin{aligned} z^i_{t+1} &= (1-\phi)\theta + \phi z^i_t - \sigma\sqrt{z^i_t}u^i_{t+1}, \\ z^w_{t+1} &= (1-\phi^w)\theta^w + \phi^w z^w_t - \sigma^w\sqrt{z^w_t}u^w_{t+1}. \end{aligned}$$

Lustig, Roussanov, and Verdelhan (2014) assume that in each country i, the logarithm of the real SDF $\widetilde{m}^i$ follows a three-factor conditionally Gaussian process:

$$-\widetilde{m}^i_{t+1} = \alpha + \chi z^i_t + \sqrt{\gamma z^i_t}u^i_{t+1} + \tau z^w_t + \sqrt{\delta^i z^w_t}u^w_{t+1} + \sqrt{\kappa z^i_t}u^g_{t+1},$$

where u^i_{t+1} is a country-specific SDF shock, while u^w_{t+1} and u^g_{t+1} are common to all countries' SDFs. All three innovations are i.i.d. Gaussian, with zero mean and unit variance. To be parsimonious, Lustig, Roussanov, and Verdelhan (2014) limit the heterogeneity in the SDF parameters to the different loadings δ^i on the world shock u^w_{t+1}; all the other parameters are identical for all countries. Therefore, the model is a restricted version of the multi-factor dynamic term structure models, and there exist closed form solutions for bond yields and risk premia.

There are two types of common shocks. The first type, u^w_{t+1}, is priced proportionally to country exposure δ^i, and since δ^i is a fixed characteristic of country i, differences in such exposure are *permanent*. The second type, u^g_{t+1}, is priced proportionally to z^i_t, so heterogeneity with respect to this innovation is *transitory*: all countries are equally exposed to this shock *on average*, but conditional exposures vary over time and depend on country-specific economic conditions. Finally, the real risk-free rate is $\widetilde{r}^{f,i}_t = \alpha + \left(\chi - \frac{1}{2}(\gamma+\kappa)\right) z^i_t + \left(\tau - \frac{1}{2}\delta^i\right) z^w_t$.

Country i's inflation process is given by $\pi^i_{t+1} = \pi_0 + \eta^w z^w_t + \sigma_\pi \epsilon^i_{t+1}$, where the inflation innovations ϵ^i_{t+1} are i.i.d. Gaussian. It follows that the log nominal risk-free rate in country i is given by $r^{f,i}_t = \pi_0 + \alpha + \left(\chi - \frac{1}{2}(\gamma+\kappa)\right) z^i_t + \left(\tau + \eta^w - \frac{1}{2}\delta^i\right) z^w_t - \frac{1}{2}\sigma^2_\pi$. The nominal bond prices in logs are affine in the state variable z and z^w: $p^{(n),i}_t = -C^{n,\$,i}_0 - C^{n,\$}_1 z_t - C^{n,\$,i}_2 z^w_t$, where the loadings $(C^{n,\$,i}_0, C^{n,\$}_1, C^{n,\$,i}_0)$ are defined in the Appendix. Equation (7) implies that the foreign currency risk premium is given by:

$$E_t(rx^{FX,i}_{t+1}) = -\frac{1}{2}(\gamma+\kappa)(z^i_t - z_t) + \frac{1}{2}(\delta - \delta^i)z^w_t.$$

Investors obtain high foreign currency risk premia when investing in currencies with relative small exposure to the two global shocks. That is the source of short-term carry trade risk premia.

SDF Decomposition The log nominal bond prices are affine in the state variable z and z^w: $p^{i,(n)}_t = -C^{i,n}_0 - C^n_1 z_t - C^{i,n}_2 z^w_t$. To calculate the parameter set $(C^{i,n}_0, C^{i,n}_1, C^{i,n}_2)$, we follow the usual recursive process. In particular, the price of a one-period nominal bond is:

$$P^{i,(1)} = E_t(M^i_{t+1}) = E_t\left(e^{-\alpha-\chi z_t - \tau z^w_t - \sqrt{\gamma z^i_t}u^i_{t+1} - \sqrt{\delta^i z^w_t}u^w_{t+1} - \sqrt{\kappa z^i_t}u^g_{t+1} - \pi_0 - \eta^w z^w_t - \sigma_\pi \epsilon^i_{t+1}}\right).$$

As a result, $C^1_0 = \alpha + \pi_0 - \frac{1}{2}\sigma^2_\pi$, $C^1_1 = \chi - \frac{1}{2}(\gamma+\kappa)$, and $C^{i,1}_2 = \tau - \frac{1}{2}\delta^i + \eta^w$.

The rest of the bond prices are calculated recursively using the Euler equation: $P^{i,(n)}_t = E_t(M^{i,\$}_{t+1}P^{i,(n-1)}_{t+1})$. This leads to the following difference equations:

$$\begin{aligned} -C^{i,n}_0 - C^n_1 z_t - C^{i,n}_2 z^w_t &= -\alpha - \chi z_t - \tau z^w_t - C^{n-1}_0 - C^{n-1}_1\left[(1-\phi)\theta + \phi z_t\right] - C^{i,n-1}_2\left[(1-\phi^w)\theta^w + \phi^w z^w_t\right] \\ &+ \frac{1}{2}(\gamma+\kappa)z_t + \frac{1}{2}\left(C^{n-1}_1\right)^2\sigma^2 z_t - \sigma\sqrt{\gamma}C^{n-1}_1 z_t \\ &+ \frac{1}{2}\delta^i z^w_t + \frac{1}{2}\left(C^{i,n-1}_2\right)^2(\sigma^w)^2 z^w_t - \sigma^w\sqrt{\delta^i}C^{i,n-1}_2 z^w_t \\ &- \pi_0 - \eta^w z^w_t + \frac{1}{2}\sigma^2_\pi \end{aligned}$$

Solving the equations above, we recover the set of bond price parameters:

$$\begin{aligned} C^{i,n}_0 &= \alpha + \pi_0 - \frac{1}{2}\sigma^2_\pi + C^{n-1}_0 + C^{n-1}_1(1-\phi)\theta + C^{i,n-1}_2(1-\phi^w)\theta^w, \\ C^n_1 &= \chi - \frac{1}{2}(\gamma+\kappa) + C^{n-1}_1\phi - \frac{1}{2}\left(C^{n-1}_1\right)^2\sigma^2 + \sigma\sqrt{\gamma}C^{n-1}_1 \\ C^{i,n}_2 &= \tau - \frac{1}{2}\delta^i + \eta^w + C^{i,n-1}_2\phi^w - \frac{1}{2}\left(C^{i,n-1}_2\right)^2(\sigma^w)^2 + \sigma^w\sqrt{\delta^i}C^{i,n-1}_2. \end{aligned}$$

The temporary pricing component of the pricing kernel is:

$$\Lambda_t^{\mathbb{T}} = \lim_{n\to\infty} \frac{\beta^{t+n}}{P_t^n} = \lim_{n\to\infty} \beta^{t+n} e^{C_0^{i,n}+C_1^n z_t + C_2^{i,n} z_t^w},$$

where the constant β is chosen in order to satisfy Assumption 1 in Alvarez and Jermann (2005): $0 < \lim_{n\to\infty} \frac{P_t^n}{\beta^n} < \infty$. The temporary pricing component of the SDF is thus equal to:

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = \beta e^{C_1^\infty (z_{t+1}-z_t)+C_2^{i,\infty}(z_{t+1}^w - z_t^w)} = \beta e^{C_1^\infty \left[(\phi-1)(z_t^i-\theta)-\sigma\sqrt{z_t^i}u_{t+1}^i\right]+C_2^{i,\infty}\left[(\phi^w-1)(z_t^w-\theta^w)-\sigma\sqrt{z_t^w}u_{t+1}^w\right]}.$$

The martingale component of the SDF is then:

$$\begin{aligned}\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} &= \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = \beta^{-1} e^{-\alpha-\chi z_t^i - \sqrt{\gamma z_t^i}u_{t+1}^i - \tau z_t^w - \sqrt{\delta^i z_t^w}u_{t+1}^w - \sqrt{\kappa z_t^i}u_{t+1}^g} \\ &\quad e^{C_1^\infty \left[(\phi-1)(z_t^i-\theta)-\sigma\sqrt{z_t^i}u_{t+1}^i\right]+C_2^{i,\infty}\left[(\phi^w-1)(z_t^w-\theta^w)-\sigma\sqrt{z_t^w}u_{t+1}^w\right]}.\end{aligned}$$

As a result, we need $\chi = C_1^\infty(1-\phi)$ to make sure that the country-specific factor does not contribute a martingale component. This special case corresponds to the absence of permanent shocks to the SDF: when $C_1^\infty(1-\phi)=\chi$ and $\kappa=0$, the permanent component of the stochastic discount factor is constant. To see this result, let us go back to the implicit definition of B_1^∞ in Equation (16):

$$\begin{aligned} 0 &= -\frac{1}{2}(\gamma+\kappa) - \frac{1}{2}\left(C_1^\infty\right)^2\sigma^2 + \sigma\sqrt{\gamma}C_1^\infty \\ 0 &= \left(\sigma C_1^\infty - \sqrt{\gamma}\right)^2, \end{aligned}$$

where we have imposed $\kappa=0$. In this special case, $C_1^\infty = \sqrt{\gamma}/\sigma$. Using this result in Equation (16), the permanent component of the SDF reduces to:

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = \beta^{-1} e^{-\tau z_t^w - \sqrt{\delta^i z_t^w}u_{t+1}^w} e^{C_2^{i,\infty}\left[(\phi^w-1)(z_t^w-\theta^w)-\sigma\sqrt{z_t^w}u_{t+1}^w\right]}.$$

Bond Premia The expected log excess return on a zero coupon bond is thus given by:

$$E_t[rx_{t+1}^{(n)}] = [-\frac{1}{2}\left(C_1^{n-1}\right)^2\sigma^2 + \sigma\sqrt{\gamma}C_1^{n-1}]z_t + [-\frac{1}{2}\left(C_2^{i,n-1}\right)^2\sigma^2 + \sigma\sqrt{\delta}^i C_2^{i,n-1}]z_t^w.$$

The expected log excess return of an infinite maturity bond is then:

$$E_t[rx_{t+1}^{(\infty)}] = [-\frac{1}{2}\left(C_1^{\infty}\right)^2\sigma^2 + \sigma\sqrt{\gamma}C_1^{\infty}]z_t + [-\frac{1}{2}\left(C_2^{i,\infty}\right)^2\sigma^2 + \sigma\sqrt{\delta}^i C_2^{i,\infty}]z_t^w.$$

The $-\frac{1}{2}\left(C_1^\infty\right)^2\sigma^2$ is a Jensen term. The term premium is driven by $\sigma\sqrt{\gamma}C_1^\infty z_t$, where C_1^∞ is defined implicitly in the second order equation $B_1^\infty = \chi - \frac{1}{2}(\gamma+\kappa) + C_1^\infty\phi - \frac{1}{2}\left(C_1^\infty\right)^2\sigma^2 + \sigma\sqrt{\gamma}C_1^\infty$. Consider the special case of $C_1^\infty(1-\phi)=\chi$ and $\kappa=0$ and $C_2^{i,\infty}(1-\phi)=\tau$. In this case, the expected term premium is simply $E_t[rx_{t+1}^{(\infty)}] = \frac{1}{2}(\gamma z_t + \delta z_t^w)$, which is equal to one-half of the variance of the log stochastic discount factor.

Currency Premia The expected log excess return of the infinite maturity bond of country i is:

$$E_t[rx_{t+1}^{(\infty),i}] = \left[C_1^\infty(1-\phi) - \chi + \frac{1}{2}(\gamma+\kappa)\right]z_t^i + \left[C_2^{i,\infty}(1-\phi^w) - \tau + \frac{1}{2}\delta^i - \eta^w\right]z_t^w.$$

The foreign currency risk premium is given by:

$$E_t[rx_{t+1}^{FX,i}] = -\frac{1}{2}(\gamma+\kappa)(z_t^i - z_t) + \frac{1}{2}(\delta-\delta^i)(z_t^w).$$

Investors obtain high foreign currency risk premia when investing in currencies whose exposure to the global shocks is smaller. That is the source of short-term carry trade risk premia. The foreign bond risk premium in dollars is simply given by the sum of the two expressions above:

$$\begin{aligned} E_t[rx_{t+1}^{(\infty),i}] + E_t[rx_{t+1}^{FX,i}] &= \left[\frac{1}{2}(\gamma+\kappa)z_t + (C_1^{\infty}(1-\phi)-\chi)z_t^i\right] \\ &+ \left[\frac{1}{2}\delta + C_2^{i,\infty}(1-\phi^w) - \tau - \eta^w\right] z_t^w. \end{aligned}$$

Simulation Results We simulate the Lustig, Roussanov, and Verdelhan (2014) model, obtaining a panel of $T = 33,600$ monthly observations and $N = 30$ countries. The calibration parameters are reported in Table A20 and the simulation results in Table A21. Each month, the 30 countries are ranked by their interest rates (Section I) or by the slope of the yield curves (Section II) into six portfolios. Low interest rate currencies on average have higher exposure δ to the world factor. As a result, these currencies appreciate in case of an adverse world shock. Long positions in these currencies earn negative excess returns rx^{fx} of -4.09% on average per annum. On the other hand, high interest rate currencies typically have high δ. Long positions in these currencies earn positive excess returns (rx^{FX}) of 2.35% on average per annum. At the short end, the carry trade strategy, which goes long in the sixth portfolio and short in the first one, delivers an excess return of 6.45% and a Sharpe ratio of 0.54.

This spread is not offset by higher local currency bond risk premia in the low interest rate countries with higher δ. The log excess return on the 30-year zero coupon bond is 0.67% in the first portfolio compared to 0.97 % in the last portfolio. At the 30-year maturity, the high-minus-low carry trade strategy still delivers a profitable excess return of 6.75% and a Sharpe ratio of 0.50. This large currency risk premium at the long end of the curve stands in stark contrast to the data. Similar results obtain when sorting countries by the slopes of their yield curves. Countries with flat yield curves tend to be countries with high short-term interest rates, while countries with steep yield curves tend to be countries with low short-term interest rates. As a result, the currency carry trade is long the last portfolio in Section II and short the first portfolio. At the 30-year maturity, the carry trade strategy still delivers a profitable excess return of 6.18% and a Sharpe ratio of 0.46.

Our theoretical results help explain the shortcomings of this simulation. In the Lustig, Roussanov, and Verdelhan (2014) calibration, the conditions for long run bond parity are not satisfied. First, global shocks have permanent effects in all countries, because $C_2^{i,\infty}(1-\phi^w) < \tau + \eta^w$ for all $i = 1, \ldots, 30$. Second, the global shocks are not symmetric, because δ varies across countries. The heterogeneity in δ's across countries generates substantial dispersion in exposure to the permanent component. As a result, our long-run uncovered bond parity condition is violated.

Finally, the Lustig, Roussanov, and Verdelhan (2014) model has country-specific and common shocks and carry trade risk premia arise from asymmetric exposures to global shocks. If the entropy of the permanent SDF component cannot differ across countries, then all countries' pricing kernels need the same loadings on the permanent component of the global factors. In the Lustig, Roussanov, and Verdelhan (2014) model, the permanent component of the SDF is given by:

$$\begin{aligned} \log\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} &= \log\beta^{-1} - \alpha - \chi z_t^i - \sqrt{\gamma z_t^i}u_{t+1}^i - \tau z_t^w - \sqrt{\delta^i z_t^w}u_{t+1}^w - \sqrt{\kappa z_t^i}u_{t+1}^g \\ & C_1^{\infty,\$}\left[(\phi-1)(z_t^i-\theta) - \sigma\sqrt{z_t^i}u_{t+1}^i\right] + C_2^{\infty,\$,i}\left[(\phi^w-1)(z_t^w-\theta^w) - \sigma\sqrt{z_t^w}u_{t+1}^w\right]. \end{aligned}$$

The U.S. term premium is simply $E_t[rx_{t+1}^{(\infty)}] = \frac{1}{2}(\gamma z_t + \delta z_t^w)$, which is equal to one-half of the variance of the log stochastic discount factor. The foreign long bond risk premium in dollars is then simply:

$$E_t[rx_{t+1}^{(\infty),*}] + E_t[rx_{t+1}^{FX,*}] = \left[\frac{1}{2}(\gamma+\kappa)z_t + (C_1^{\infty,\$}(1-\phi)-\chi)z_t^*\right] + \left[\frac{1}{2}\delta + C_2^{\infty,\$,*}(1-\phi^w) - \tau - \eta^w\right] z_t^w,$$

where $C_1^{\infty,\$}, C_2^{\infty,\$}$ represent the loadings of the nominal long rates on the two factors. Condition 1 thus holds if $C_1^{\infty,\$}(1-\phi) = \chi$, $\kappa = 0$, and $C_2^{\infty,\$,*}(1-\phi^w) = \tau + \eta^w$. The first two restrictions rule out permanent effects of country-specific shocks, while the last restriction rules out permanent effects of global shocks (u^w). When these restrictions are satisfied, the pricing kernel is not subject to permanent shocks, and the expected foreign log holding period return on a foreign long-term bond converted into U.S. dollars is equal to the U.S. term premium: $E_t[rx_{t+1}^{(\infty),*}] + E_t[rx_{t+1}^{FX,*}] = \frac{1}{2}(\gamma z_t + \delta z_t^w)$. The higher foreign currency risk premium for investing in high δ countries is exactly offset by the lower bond risk premium. As all these models show, Proposition 1 and Condition 1 offer a simple diagnostic to assess the term structure of currency carry trade risk premia in no-arbitrage models.

The restrictions $C_1^{\infty,\$}(1-\phi) = \chi$, $\kappa = 0$, and $C_2^{\infty,\$,*,}(1-\phi^w) = \tau + \eta^w$ have a natural interpretation as restrictions on the long-run loadings of the exchange rate on the risk factors: $\sum_{i=1}^{\infty} E_t[\Delta s_{t+i}] = \sum_{i=1}^{\infty} E_t[m_{t+i} - m_{t+i}^*] = \sum_{i=1}^{\infty} \phi^{i-1}\chi(z_t^* - z_t)$. As can easily be verified, these two restrictions imply that the long-run loading of the exchange rate on the factors equals the loading of

Table A20: Parameter Estimates

Stochastic discount factor					
α (%)	χ	τ	γ	κ	δ
0.76	0.89	0.06	0.04	2.78	0.36
State variable dynamics					
ϕ	θ (%)	σ (%)	ϕ^w	θ^w (%)	σ^w (%)
0.91	0.77	0.68	0.99	2.09	0.28
Inflation dynamics				Heterogeneity	
η^w	π_0 (%)	σ^π (%)		δ_h	δ_l
0.25	−0.31	0.37		0.22	0.49
Implied SDF dynamics					
$E(Std_t(\widetilde{m}))$	$Std(Std_t(\widetilde{m}))$ (%)	$E(Corr(\widetilde{m}_{t+1}, \widetilde{m}^i_{t+1}))$		$Std(z)$ (%)	$Std(z^w)$ (%)
0.59	4.21	0.98		0.50	1.32

Notes: This table reports the parameter values for the estimated version of the model. The model is defined by the following set of equations:

$$
\begin{aligned}
-\widetilde{m}^i_{t+1} &= \alpha + \chi z^i_t + \sqrt{\gamma z^i_t} u^i_{t+1} + \tau z^w_t + \sqrt{\delta^i z^w_t} u^w_{t+1} + \sqrt{\kappa z^i_t} u^g_{t+1}, \\
z^i_{t+1} &= (1-\phi)\theta + \phi z^i_t - \sigma \sqrt{z^i_t} u^i_{t+1}, \\
z^w_{t+1} &= (1-\phi^w)\theta^w + \phi^w z^w_t - \sigma^w \sqrt{z^w_t} u^w_{t+1}, \\
\pi^i_{t+1} &= \pi_0 + \eta^w z^w_t + \sigma_\pi \epsilon^i_{t+1}.
\end{aligned}
$$

All countries share the same parameter values except for δ^i, which is distributed uniformly on $[\delta_h, \delta_l]$. The home country exhibits the average δ, which is equal to 0.36.

Table A21: Simulated Excess Returns on Carry Strategies in the Lustig, Roussanov, and Verdelhan (2014) Model

	Low	2	3	4	5	High
Section I: Sorting by Interest Rate Levels						
Panel A: Exchange Rates, Interest Rates, and Bond Returns						
Δs	1.93	0.79	0.44	0.06	−0.16	−0.85
$\sigma_{\Delta s}$	11.04	9.55	9.06	8.98	9.02	9.54
$r^{f,*} - r^f$	−2.16	−1.21	−0.63	−0.10	0.43	1.50
$rx^{(30),*}$	0.67	0.75	0.79	0.89	0.93	0.97
Panel B: Carry Returns with Short-Term Bills						
rx^{FX}	−4.09	−2.00	−1.06	−0.16	0.59	2.35
Panel C: Carry Returns with Long-Term Bonds						
$rx^{(30),\$}$	−3.42	−1.25	−0.27	0.72	1.52	3.33
Section II: Sorting by Interest Rate Slopes						
Panel A: Exchange Rates, Interest Rate Slopes, and Bond Returns						
Δs	−2.06	−1.12	−0.49	−0.03	0.50	1.92
$\sigma_{\Delta s}$	11.35	9.60	8.97	8.84	8.95	9.93
$y^{10} - y^{1/4}$	−0.87	−0.42	−0.13	0.12	0.38	1.03
$rx^{(30),*}$	0.87	0.87	0.86	0.87	0.86	0.84
Panel B: Carry Returns with Short-Term Bills						
rx^{FX}	3.23	1.78	0.83	0.08	−0.76	−2.92
Panel C: Carry Returns with Long-Term Bonds						
$rx^{(30),\$}$	4.09	2.65	1.69	0.94	0.11	−2.09

Notes: The table reports summary statistics on simulated data from the Lustig, Roussanov, and Verdelhan (2014) model. Data are obtained from a simulated panel with $T = 33,600$ monthly observations and $N = 30$ countries. In Section I, countries are sorted by interest rates into six portfolios. In Section II, they are sorted by the slope of their yield curves (defined as the difference between the 10-year yield and the three-month yield). In each section, Panel A reports the average change in exchange rate (Δs), the average interest rate difference ($r^{f,*} - r^f$) (or the average slope, $y^{10} - y^{1/4}$), the average foreign bond excess returns for bonds of 30-year maturities in local currency ($rx^{(30),*}$). Panel B reports the average log currency excess returns (rx^{FX}). Panel C reports the average foreign bond excess returns for bonds of 30-year maturities in home currency ($rx^{(30),\$}$). The moments are annualized.

long-term interest rates:

$$\lim_{k\to\infty} E_t[\Delta s_{t\to t+k}] = \frac{\chi}{(1-\phi)}(z_t^* - z_t) = C_1^{\infty,\$}(z_t^* - z_t) = \lim_{k\to\infty} k\left(y_t^{(k),*} - y_t^{(k)}\right),$$

where we have used $C_2^{\infty,\$} = C_2^{\infty,\$,*} = \tau + \eta^w$. Hence, in the context of this model, our restrictions enforce long-run U.I.P.[16] In this special case, $\frac{\chi}{(1-\phi)} = C_1^{\infty} = \sqrt{\gamma}/\sigma > 0$. An increase in risk abroad causes the long rates to go up abroad and the foreign exchange rate to depreciate in the long run, but given these long-run restrictions, the initial expected exchange rate impact has to have the same sign ($\chi > 0$), thus violating the empirical evidence, as we explain below.

Our preference-free conditions constrains the sum of slope regression coefficients in a regression of future exchange rate changes Δs_{t+i} on the current interest rate spread $r_t^{f,\$,*} - r_t^{f,\$}$ to be equal to the response of long-term interest rates. Engel (2016), Valchev (2016), and Dahlquist and Penasse (2016) study these slope coefficients and find that they switch signs with the horizon i: an increase in the short-term interest rate initially cause exchange rates to appreciate, but they subsequently depreciate on average. In the factor model with a single country-specific factor, these slope coefficients in a regression of Δs_{t+i} on the $r_t^{f,\$,*} - r_t^{f,\$}$, given by

$$E_t \Delta s_{t+i} = \frac{\phi^{i-1}\chi}{\chi - \frac{1}{2}\gamma}\left(r_t^{f,\$,*} - r_t^{f,\$}\right),$$

decline geometrically as i increases, and their infinite sum equals $\frac{C_1^{\infty}}{\chi - \frac{1}{2}\gamma}$. When $(\chi - \frac{1}{2}\gamma) < 0$, the model can match the short-run forward premium puzzle: when the foreign short rate increases, the currency subsequently appreciates, but it continues to appreciate as long rates decline abroad. As a result, this model cannot match the sign switch in these regression coefficients. A richer version of the factor model with multiple country-specific risk factors can generate richer dynamics. Consider the same model with two country-specific risk factors. The long-run impulse responses of the exchange rate to short-term interest rate shocks is driven by:

$$\sum_{i=1}^{\infty} E_t[\Delta s_{t+i}] = \sum_{i=1}^{\infty} E_t[m_{t+i} - m_{t+i}^*] = \sum_{i=1}^{\infty}\left[\phi_1^{i-1}\chi_1(z_t^{1,*} - z_t^1) + \phi_2^{i-1}\chi_2(z_t^{2,*} - z_t^2)\right].$$

The slope coefficients in a regression of future exchange rate changes on the current interest rate spread $r_t^{f,\$,*} - r_t^{f,\$}$ are given by

$$E_t \Delta s_{t+i} = \frac{\phi_1^{i-1}\chi_1(\chi_1 - \frac{1}{2}\gamma_1) + \phi_2^{i-1}\chi_2(\chi_2 - \frac{1}{2}\gamma_2)}{(\chi_1 - \frac{1}{2}\gamma_1)^2 + (\chi_2 - \frac{1}{2}\gamma_2)^2}\left(r_t^{f,\$,*} - r_t^{f,\$}\right).$$

These coefficients can switch signs as we increase the maturity i if the risk factors have sufficiently heterogeneous persistence (ϕ_1, ϕ_2), and provided that $(\chi_1 - \frac{1}{2}\gamma_1)$ and $(\chi_2 - \frac{1}{2}\gamma_2)$ have opposite signs.

D.6 Sketching a Model with Temporary and Permanent Shocks

The Lustig, Roussanov, and Verdelhan (2014) calibration fails to replicate the term structure of carry trade risk premia. We turn to a model that explicitly features global permanent and transitory shocks. We show that the heterogeneity in the SDFs' loadings on the *permanent* global shocks needs to be ruled out in order to match the empirical evidence on the term structure of carry risk.

Model We assume that in each country i, the logarithm of the real SDF m^i follows a three-factor conditionally Gaussian process:

$$-m_{t+1}^i = \alpha + \chi z_t^i + \sqrt{\gamma z_t^i}u_{t+1}^i + \tau^i z_t^w + \sqrt{\delta^i z_t^w}u_{t+1}^w + \tau^{\mathbb{P},i}z_t^{\mathbb{P},w} + \sqrt{\delta^{\mathbb{P}} z_t^{\mathbb{P},w}}u_{t+1}^w + \sqrt{\kappa z_t^i}u_{t+1}^g.$$

The state variables follow similar square root processes as in the previous model:

$$\begin{aligned}
z_{t+1}^i &= (1-\phi)\theta + \phi z_t^i - \sigma\sqrt{z_t^i}u_{t+1}^i, \\
z_{t+1}^w &= (1-\phi^w)\theta^w + \phi^w z_t^w - \sigma^w\sqrt{z_t^w}u_{t+1}^w. \\
z_{t+1}^{\mathbb{P},w} &= (1-\phi^{\mathbb{P},w})\theta^{\mathbb{P},w} + \phi^{\mathbb{P},w} z_t^w - \sigma^{\mathbb{P},w}\sqrt{z_t^{\mathbb{P},w}}u_{t+1}^{\mathbb{P},w}.
\end{aligned}$$

But one of the common factors, z_t^w, is rendered transitory by imposing that $C_2^{i,\infty}(1-\phi^w) = \tau^i$. To make sure that the global shocks have no permanent effect for each value of δ^i, we need to introduce another source of heterogeneity across countries. Countries

[16]When all innovations have an impact on risk, as is the case in this model, Condition 1 rules out permanent shocks.

must differ in τ, ϕ^w, σ^w, or η^w (or a combination of those). Without this additional source of heterogeneity, there are at most two values of δ^i that are possible (for each set of parameters). [17] Here we simply choose to let the parameters τ differ across countries.

Bond Prices Our model only allows for heterogeneity in the exposure to the transitory common shocks (δ^i), but not in the exposure to the permanent common shock ($\delta^{\mathbb{P}}$). The nominal log zero-coupon n-month yield of maturity in local currency is given by the standard affine expression $y_t^{(n)} = \frac{1}{n}\left(C_0^n + C_1^n z_t + C_2^n z_t^w + C_3^n z_t^{\mathbb{P},w}\right)$, where the coefficients satisfy second-order difference equations. Given this restriction, the bond risk premium is equal to:

$$\begin{aligned} E_t[rx_{t+1}^{(i,\infty)}] &= \left[C_1^\infty(1-\phi) - \chi + \frac{1}{2}(\gamma+\kappa)\right] z_t + \frac{1}{2}\delta^i z_t^w \\ &+ \left[C_3^\infty(1-\phi^{\mathbb{P},w}) - \tau^{\mathbb{P}} + \frac{1}{2}\delta^{\mathbb{P}} - \eta^w\right] z_t^{\mathbb{P},w}. \end{aligned}$$

The log currency risk premium is equal to $E_t[rx_{t+1}^{FX,i}] = (\gamma+\kappa)(z_t - z_t^i)/2 + (\delta - \delta^i)z_t^w/2$. The permanent factor $z_t^{w,\mathbb{P}}$ drops out. This also implies that the expected foreign log holding period return on a foreign long bond converted into U.S. dollars is equal to:

$$\begin{aligned} E_t[rx_{t+1}^{(i,\infty)}] + E_t[rx_{t+1}^{FX,i}] &= \left[(C_1^\infty(1-\phi) - \chi)z_t^i + \frac{1}{2}(\gamma+\kappa)z_t\right] + \frac{1}{2}\delta z_t^w \\ &+ \left[C_3^\infty(1-\phi^{\mathbb{P},w}) - \tau^{\mathbb{P}} + \frac{1}{2}\delta^{\mathbb{P}} - \eta^w\right] z_t^{\mathbb{P},w}. \end{aligned}$$

Given the symmetry that we have imposed, the difference between the foreign term premium in dollars and the domestic term premium is then given by: $[C_1^\infty(1-\phi) - \chi]\,(z_t^i - z_t)$. There is no difference in long bond returns that can be traced back to the common factor; only the idiosyncratic factor. The spread due to the common factor is the only part that matters for the long-term carry trade, which approximately produces zero returns here.

Term Structure of Carry Trade Risk Premia To match short-term carry trade returns, we need asymmetric exposure to the transitory shocks, governed by (δ), but not to permanent shocks, governed by ($\delta^{\mathbb{P}}$). If the foreign kernel is less exposed to the transitory shocks than the domestic kernel ($\delta > \delta^i$), there is a large positive foreign currency risk premium (equal here to $(\delta - \delta^i)z_t^w/2$), but that premium is exactly offset by a smaller foreign term premium and hence does not affect the foreign bond risk premium in dollars. The countries with higher exposure will also tend to have lower interest rates when the transitory volatility z_t increases, provided that $\left(\tau - \frac{1}{2}\delta\right) < 0$. Hence, in this model, the high δ^i funding currencies in the lowest interest rate portfolios will tend to earn negative currency risk premia, but positive term premia. The reverse would be true for the low δ^i investment currencies in the high interest rate portfolios. This model thus illustrates our main theoretical findings: chasing high interest rates does not necessarily work at the long end of the maturity spectrum. If there is no heterogeneity in the loadings on the permanent global component of the SDF, then the foreign term premium on the longest bonds, once converted to U.S. dollars is identical to the U.S. term premium.

[17] This result appears when plugging the no-permanent-component condition in the differential equation that governs the loading of the bond price on the global state variable.

E Structural Dynamic Asset Pricing Models

This section of the Appendix presents the details of pricing kernel decomposition for three classes of structural dynamic asset pricing models equilibrium models: models with external habit formation, models with long run risks, and models with rare disasters. In a nutshell, among the reduced-form term structure models we consider, Condition 1 implies novel parameter restrictions for all models (and in some cases, it rules out all permanent shocks or the time-variation in the price of risk). In the habit model with common shocks, the carry trade risk premia and Condition 1 imply that countries exhibit the same risk-aversion and the same volatility of consumption growth shocks but differ in the persistences of their habit levels. The long-run risk models satisfy Condition 1 only with common shocks and for knife-edge parameter values. For the disaster risk models, common shocks are also necessary for Condition 1 to hold and the downward term structure of carry trade risk premia implies heterogeneity in the rate of time preference, the rate of depreciation, or the country-specific growth rate, but no heterogeneity in the coefficient of risk aversion, the common and country-specific consumption drops in case of a disaster, and the probability of a disaster. In order to save space, we summarize the implications of Condition 1 in Table A22.

Table A22: Condition 1: Dynamic Asset Pricing Model Scorecard

	Symmetric Model with Country-specific Shocks	*Asymmetric Model with Common Shocks*
External Habit Model	✓	✓Only Heterogeneity in ϕ No Heterogeneity in (γ, σ^2)
Long Run Risks Model	No	Only Knife-edge cases
Disaster Model	No	✓Only Heterogeneity in (R, λ, g_w) No Heterogeneity in (γ, B, F, p_t)

This table summarizes whether each class of models can satisfy Condition 1. The left section of the table focuses on models with only country-specific shocks in which all countries have the same parameters. The right section focuses on models models with only common shocks and heterogeneity in the parameters. In the external habit model (Campbell and Cochrane, 1999; Wachter, 2006; Verdelhan, 2010; Stathopoulos, 2017), the parameters ϕ and B govern the dynamics of the surplus consumption ratio process. In the long run risk model (Bansal and Yaron, 2004; Colacito and Croce, 2011; Bansal and Shaliastovich, 2013; Engel, 2016), Condition 1 is always violated except in knife-edge cases. In the disaster model (Farhi and Gabaix, 2016; Gabaix, 2012; Wachter, 2013) , the parameters R, λ, and g_w govern the rate of time preference, the rate of depreciation and the country-specific growth rate, while the parameters γ, B, F, p_t are the coefficient of risk aversion, the common consumption drop in case of a disaster, the country-specific consumption drop in case of a disaster, and the probability of a disaster. The details are in section E of the Online Appendix.

E.1 External Habit Model

In the Campbell and Cochrane (1999) habit model used by Wachter (2006), Verdelhan (2010), and Stathopoulos (2017) to study the properties of interest rates and exchange rates, the log pricing kernel has law of motion

$$\log \frac{\Lambda_{t+1}}{\Lambda_t} = \log \delta - \gamma g - \gamma(1-\phi)(\bar{su} - s_t) - \gamma(1+\lambda(s_t))\varepsilon_{t+1},$$

with the aggregate consumption growth rate satisfying

$$\Delta c_{t+1} = g + \varepsilon_{t+1},$$

with $\varepsilon_{t+1} \sim N(0, \sigma^2)$, and the log surplus consumption ratio evolving as follows:

$$s_{t+1} = (1-\phi)\bar{su} + \phi s_t + \lambda(s_t)\varepsilon_{t+1}.$$

Finally, the sensitivity function λ is

$$\lambda(s_t) = \begin{cases} \frac{1}{\bar{S}}\sqrt{1-2(s_t-\bar{su})}-1, \text{if } s < s_{max} \\ 0, \text{if } s \geq s_{max} \end{cases},$$

where $\bar{S} = \sigma\sqrt{\frac{\gamma}{1-\phi-B/\gamma}}$ is the steady-state value of the surplus consumption ratio and $s_{max} = \bar{su} + \frac{1}{2}(1-\bar{S}^2)$ is the upper bound of the log surplus consumption ratio. The parameter B is important, as its sign determines the cyclicality of the real interest rate.

The equilibrium log risk-free rate is

$$r_t^f = -E_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = -\log\delta + \gamma g + \gamma(1-\phi)(\bar{su}-su_t) - \frac{1}{2}\gamma^2\sigma^2(1+\lambda(su_t))^2,$$

which can be also written as

$$r_t^f = -\log\delta + \gamma g - \frac{1}{2}\frac{\gamma^2\sigma^2}{\bar{Su}^2} + B(\bar{su}-su_t).$$

The parameter $B = \gamma(1-\phi) - \frac{\gamma^2\sigma^2}{\bar{S}^2}$. Therefore, if $B = 0$, the log risk-free rate is constant: the intertemporal smoothing effect is exactly offset by the precautionary savings effect. If, on the other hand, $B \neq 0$, then the log risk-free rate is perfectly correlated with the surplus consumption ratio s: it is negatively correlated with su (and hence countercyclical) if $B > 0$, and positively correlated with s (and hence procyclical) if $B < 0$. This is because, if $B > 0$, the intertemporal smoothing effect dominates the precautionary savings effect: when su is above its steady-state level, mean-reversion implies that marginal utility is expected to increase in the future, incentivizing agents to save and decreasing interest rates. On the other hand, if $B < 0$, the precautionary savings motive dominates, so agents save more when s is low and marginal utility is more volatile.

To decompose the pricing kernel, we use the guess and verify method. In particular, guess an eigenfunction ϕ of the form

$$\phi(s) = e^{cs},$$

where c is a constant. Then, the (one-period) eigenfunction problem can be written as

$$E_t\left[\exp\left(\log\delta - \gamma\left[g + (\phi-1)(su_t-\bar{su}) + (1+\lambda(su_t))\varepsilon_{t+1}\right] + csu_{t+1}\right)\right] = \exp(\beta + csu_t)$$

which, after some algebra, yields

$$\log\delta - \gamma g - \gamma(\phi-1)(su_t-\bar{su}) + c(1-\phi)\bar{su} + c\phi su_t + \frac{\sigma^2}{2}\left((c-\gamma)(1+\lambda(su_t)) - c\right)^2 = \beta + csu_t.$$

Setting $c = \gamma$, the expression above becomes

$$\log\delta - \gamma g - \gamma(\phi-1)(su_t-\bar{su}) + \gamma(1-\phi)\bar{su} + \gamma\phi su_t + \frac{\gamma^2\sigma^2}{2} = \beta + \gamma su_t,$$

and, matching the constant terms, we get $\beta = \log\delta - \gamma g + \frac{\gamma^2\sigma^2}{2}$. Therefore, the transitory component of the pricing kernel is $\Lambda_t^{\mathbb{T}} = e^{\beta t - csu_t}$, so the transitory pricing kernel component is

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = e^{\beta - c(su_{t+1}-su_t)} = e^{\log\delta - \gamma g + \frac{\gamma^2\sigma^2}{2} - \gamma\left((1-\phi)(\bar{su}-su_t) + \lambda(su_t)\varepsilon_{t+1}\right)},$$

and the permanent pricing kernel component is

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = e^{-\frac{\gamma^2\sigma^2}{2} - \gamma\varepsilon_{t+1}}.$$

In the Campbell and Cochrane (1999) model, the permanent pricing kernel component reflects innovations in consumption growth, which permanently affect the level of consumption, whereas the transitory pricing kernel component is driven by innovations in the surplus consumption ratio, which is a stationary variable. However, the two types of innovations are perfectly correlated by assumption, so the two pricing kernel components exhibit positive comovement: a negative consumption growth innovation not only permanently reduces the level of consumption, but also transitorily decreases the surplus consumption ratio of the agent, increasing the local curvature of her utility function. As a result, a negative consumption growth shock implies a positive shock for both pricing kernel components.

Finally, we consider the properties of the pricing kernel and its components. In each country, the conditional entropy of the pricing kernel is

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{\gamma^2\sigma^2}{2}(1+\lambda(su_t))^2 = \frac{\gamma^2\sigma^2}{2}\frac{1}{\bar{S}^2}(1-2(su_t-\bar{su})),$$

the conditional entropy of the permanent pricing kernel component is

$$L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \frac{\gamma^2\sigma^2}{2},$$

and the conditional entropy of the transitory pricing kernel component is

$$L_t\left(\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\right) = \frac{\gamma^2\sigma^2}{2}\lambda(su_t)^2.$$

Notably, the permanent pricing kernel component has constant conditional entropy, whereas the conditional entropy of both the pricing kernel and the transitory pricing kernel component are time varying, as they are functions of the log surplus consumption ratio s. It follows that the conditional term premium, in local currency terms, is

$$E_t\left[rx^{(\infty)}_{t+1}\right] = L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \frac{\gamma^2\sigma^2}{2}\frac{1}{\bar{S}^2}(1-2(su_t-\bar{su})) - \frac{\gamma^2\sigma^2}{2} = \frac{\gamma^2\sigma^2}{2}\left[\frac{1}{\bar{Su}^2}(1-2(su_t-\bar{su}))-1\right].$$

The conditional SDF entropy is

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{\gamma^2\sigma^2}{2}(1+\lambda(su_t))^2 = \frac{\gamma^2\sigma^2}{2}\frac{1}{\bar{S}^2}(1-2(su_t-\bar{su})).$$

The currency risk premium is given by $E_t[rx^{FX}_{t+1}] = \frac{\gamma^2\sigma^2}{2\bar{S}^2}(su^*_t - su_t)$.

If $B > 0$, then the log risk-free rate is countercyclical: when s is above its steady-state level, mean-reversion implies that marginal utility is expected to increase in the future, incentivizing agents to save and decreasing interest rates. Wachter (2006) shows that this condition is necessary for an upward sloping real term structure of interest rates. However, as pointed out by Verdelhan (2010), the model requires procyclical interest rates ($B < 0$) in order to generate the empirically observed relationship between interest rate differentials and currency risk premia at the short end: as equation (7) implies, there must be more priced risk in low interest rate countries than in high interest countries. Hence, this model cannot match currency risk premia and term premia. The price of the long-term bond converges to $\lim_{k\to\infty} P^{(k)}_t = \exp(\gamma(su_t - \bar{su}))$. Finally, the permanent component of exchange rate changes is given by $\log\left(\frac{S^{\mathbb{P}}_{t+1}}{S^{\mathbb{P}}_t}\right) = -\gamma\left(\Delta c_{t+1} - \Delta c^*_{t+1}\right)$, so it is not affected by the surplus consumption ratio. The conditional entropy of the permanent SDF component is constant (Borovička, Hansen, and Scheinkman, 2016):

$$L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right) = \frac{\gamma^2\sigma^2}{2},$$

whereas the conditional entropies of both the SDF and the transitory SDF component are time varying, as they are functions of the log surplus consumption ratio su.

Result 4. *To satisfy Condition 1 in the external habit model, the following restriction needs to hold* $\gamma\sigma^2 = \gamma^*\sigma^{*,2}$.

In a symmetric habit model (i.e., a model in which all countries share the same parameters) with country-specific shocks, Condition 1 is automatically satisfied: variation in the price of risk, governed by su, does not affect marginal utility and exchange rates in the long run. We can generalize this model to N countries. In order for Condition 1 to hold, countries can only differ in their surplus consumption ratio persistence parameters (ϕ), as differences in the other parameters (γ, σ^2) would imply differences in the conditional entropy of the permanent component of the pricing kernels, and thus differences in long-maturity bond returns expressed in the same units. Thus, Condition 1 limits the source of heterogeneity to choices leading to different real interest rate persistence across countries.

Finally, consider a symmetric version of the two-country external habit model in which the 2 countries share all of the parameters. The long-run loading of the exchange rate on the surplus consumption ratio is given by: $\sum_{i=1}^{\infty} E_t[\Delta s_{t+i}] = \sum_{i=1}^{\infty} E_t[\log\frac{\Lambda_{t+i}}{\Lambda_t} - \log\frac{\Lambda^*_{t+i}}{\Lambda^*_t}] = -\sum_{i=1}^{\infty}\phi^{i-1}\gamma(1-\phi)(su^*_t - su_t) = -\gamma(su^*_t - su_t)$. Thus, long-run U.I.P holds,

$$\lim_{k\to\infty} E_t[\Delta s_{t\to t+k}] = -\gamma(su^*_t - su_t) = \lim_{k\to\infty} k\left(y^{(k),*}_t - y^{(k)}_t\right),$$

even though exchange rates are non-stationary in levels, because the innovations to risk premia, driven by the surplus consumption ratio, are transitory. A decrease in the foreign surplus consumption ratio causes foreign long-term rates to increase and the foreign currency to depreciate in the long run.

Result 5. *In the symmetric external habit model, the slope coefficients in regressions of Δs_{t+i} on the interest rate spread $r_t^{f,*} - r_t^f$, given by $-\frac{\phi^{i-1}\gamma(1-\phi)}{B}$, decline geometrically in absolute value as i increases, and their infinite sum equals $-\frac{\gamma}{B}$.*

When $B < 0$, all these slope coefficients are positive: a decrease in the foreign short rate causes the foreign currency to depreciate on average next period and all periods after that, in line with the increase in the foreign long rate. As pointed out by Engel (2016), these slope coefficients cannot switch signs to match the empirical evidence.

E.2 Long-Run Risks Model

We now consider the long-run risks model, proposed by Bansal and Yaron (2004) and further explored by Colacito and Croce (2011), Bansal and Shaliastovich (2013) and Engel (2016) in the context of exchange rates. In this class of models, the representative agent has utility over consumption given by:

$$\log U_t = (1 - \frac{1}{\psi}) \log \left((1-\delta) C^{1-\frac{1}{\psi}} + \delta E_t \left[U_{t+1}^{1-\gamma} \right]^{\frac{1-\frac{1}{\psi}}{1-\gamma}} \right),$$

where ψ represents the intertemporal elasticity of substitution in an environment without risk. Aggregate consumption growth Δc_{t+1} has a persistent component x_t, and both consumption growth shocks and shocks in x_t exhibit conditional heteroskedasticity:

$$\begin{aligned} \Delta c_{t+1} &= \mu + x_t + \sqrt{u_t}\varepsilon_{t+1}^c, \\ x_{t+1} &= \phi^x x_t + \sqrt{w_t}\varepsilon_{t+1}^x, \\ u_{t+1} &= (1-\phi^u)\theta^u + \phi^u u_t + \sigma^u \varepsilon_{t+1}^u, \\ w_{t+1} &= (1-\phi^w)\theta^w + \phi^w w_t + \sigma^w \varepsilon_{t+1}^w. \end{aligned}$$

All innovations are i.i.d. standard normal. The log SDF evolves as:

$$\log \frac{\Lambda_{t+1}}{\Lambda_t} = A_0 + A_1 x_t + A_2 u_t + A_3 w_t + B_1 \sqrt{u_t}\varepsilon_{t+1}^c + B_2 \sqrt{w_t}\varepsilon_{t+1}^x + B_3 \varepsilon_{t+1}^u + B_4 \varepsilon_{t+1}^w,$$

where $\{A_0, A_1, A_2, A_3, B_1, B_2, B_3, B_4\}$ are constants.[18] For convenience, we assume that the agent has preferences for early resolution of uncertainty ($\gamma > \frac{1}{\psi}$), so $B_2 < 0$. It immediately follows that conditional SDF entropy and the equilibrium log risk-free rate are given by:

$$\begin{aligned} L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) &= \frac{1}{2} var_t \left(\log \frac{\Lambda_{t+1}}{\Lambda_t} \right) = \frac{1}{2}\left({B_1}^2 u_t + {B_2}^2 w_t + {B_3}^2 + {B_4}^2 \right), \\ r_t^f &= -A_0 - \frac{1}{2}\left({B_3}^2 + {B_4}^2 \right) + \frac{1}{\psi} x_t - \frac{1}{2}\left(\frac{\gamma - 1}{\psi} + \gamma \right) u_t - \frac{1}{2}\left(\frac{1}{\psi} - \gamma \right)\left(\frac{1}{\psi} - 1 \right)\left(\frac{\kappa}{1 - \kappa\phi^x} \right)^2 w_t. \end{aligned}$$

The necessary condition (7) highlights how this model can replicate the U.I.P. puzzle: for procyclical interest rates (with respect to u_t and w_t), high interest rates correspond to low volatility SDFs.

The real bond prices in logs are affine in the state variables: $p_t^{i,(n)} = -C_0^{i,n} - C_1^n x_t - C_2^{i,n} u_t - C_3^{i,n} w_t$. In the long-run risk model, the conditional entropy of the permanent SDF component is given by:

$$L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \frac{1}{2} var_t \left(\log \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} \right) = \frac{1}{2}\left({B_1}^2 u_t + (B_2 - C_1^\infty)^2 w_t + (B_3 - C_2^\infty \sigma^u)^2 + (B_4 - C_3^\infty \sigma^w)^2 \right).$$

where $C_1^\infty = \frac{1}{\psi(1-\phi^x)}$, $C_2^\infty = -\frac{A_2 + \frac{1}{2}B_1^2}{1-\phi^u}$, and $C_3^\infty = -\frac{A_3 + \frac{1}{2}(B_2 - C_1^\infty)^2}{1-\phi^w}$.

[18] More precisely, the constants are $A_1 = -\frac{1}{\psi}$, $A_2 = \left(\frac{1}{\psi} - \gamma\right)\frac{\gamma - 1}{2}$, $A_3 = \left(\frac{1}{\psi} - \gamma\right)\frac{\gamma-1}{2}\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2$, $B_1 = -\gamma$, $B_2 = \left(\frac{1}{\psi} - \gamma\right)\frac{\kappa}{1-\kappa\phi^x}$, $B_3 = \left(\frac{1}{\psi} - \gamma\right)\frac{1-\gamma}{2}\frac{\kappa}{1-\kappa\phi^u}\sigma^u$, and $B_4 = \left(\frac{1}{\psi} - \gamma\right)\frac{1-\gamma}{2}\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2\frac{\kappa}{1-\kappa\phi^w}\sigma^w$, where $\kappa \equiv \frac{\delta e^{\left(1-\frac{1}{\psi}\right)\bar{m}}}{1-\delta+\delta e^{\left(1-\frac{1}{\psi}\right)\bar{m}}}$ and $\bar{m}$ is the point around which a log-linear approximation is taken (see Engel (2015) for details); if $\bar{m} = 0$, then $\kappa = \delta$.

Table A23: Pricing Kernel Loadings in the Long Run Risks Model

Loadings	Parameters	Loadings	Parameters
Panel A: Homoskedastic Model			
$\log\frac{\Lambda_{t+1}}{\Lambda_t} = A_0 + A_1 x_t + B_1\sqrt{\theta^u}\varepsilon^c_{t+1} + B_2\sqrt{\theta^w}\varepsilon^x_{t+1}.$			
A_1	$-\frac{1}{\psi}$	B_1	$-\gamma$
		B_2	$\left(\frac{1}{\psi}-\gamma\right)\frac{\kappa}{1-\kappa\phi^x}$
Transitory Component $\log\Lambda^{\mathbb{T}}_t = \beta t - c x_t$			
c_1	$\frac{A_1}{1-\phi^x}$	β	$A_0 + \frac{1}{2}{B_1}^2\theta^u + \frac{1}{2}(B_2+c)^2\theta^w,$
Panel B: Heteroskedastic Model			
$\log\frac{\Lambda_{t+1}}{\Lambda_t} = A_0 + A_1 x_t + A_2 u_t + A_3 w_t + B_1\sqrt{u_t}\varepsilon^c_{t+1} + B_2\sqrt{w_t}\varepsilon^x_{t+1} + B_3\varepsilon^u_{t+1} + B_4\varepsilon^w_{t+1}.$			
A_1	$-\frac{1}{\psi}$	B_1	$-\gamma$
A_2	$\left(\frac{1}{\psi}-\gamma\right)\frac{\gamma-1}{2}$	B_2	$\left(\frac{1}{\psi}-\gamma\right)\frac{\kappa}{1-\kappa\phi^x}$
A_3	$\left(\frac{1}{\psi}-\gamma\right)\frac{\gamma-1}{2}\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2$	B_3	$\left(\frac{1}{\psi}-\gamma\right)\frac{1-\gamma}{2}\frac{\kappa}{1-\kappa\phi^u}\sigma^u$
		B_4	$\left(\frac{1}{\psi}-\gamma\right)\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2\frac{\kappa}{1-\kappa\phi^w}\sigma^w.$
Transitory Component $\log\Lambda^{\mathbb{T}}_t = \beta t - c_1 x_t - c_2 u_t - c_3 w_t$			
c_1	$\frac{A_1}{1-\phi^x}$	c_2	$\frac{A_2+\frac{1}{2}B_1^2}{1-\phi^u}$
c_3	$\frac{A_3+\frac{1}{2}(B_2+\frac{A_1}{1-\phi^x})^2}{1-\phi^w}$	β	$A_0 + c_2(1-\phi^u)\theta^u + c_3(1-\phi^w)\theta^w + \frac{1}{2}(B_3+c_2\sigma^u)^2 + \frac{1}{2}(B_4+c_3\sigma^w)^2.$

Notes: Pricing kernel loading parameters in the long run risks model. Parameter κ is defined as $\kappa \equiv \frac{\delta e^{\left(1-\frac{1}{\psi}\right)\bar{m}}}{1-\delta+\delta e^{\left(1-\frac{1}{\psi}\right)\bar{m}}}$, where $\bar{m}$ is the point around which a log-linear approximation is taken (see Engel (2016) for details); if $\bar{m}=0$, then $\kappa=\delta$.

In this full version of the model, the log SDF follows the law of motion:

$$\log \frac{\Lambda_{t+1}}{\Lambda_t} = A_0 + A_1 x_t + A_2 u_t + A_3 w_t + B_1 \sqrt{u_t} \varepsilon^c_{t+1} + B_2 \sqrt{w_t} \varepsilon^x_{t+1} + B_3 \varepsilon^u_{t+1} + B_4 \varepsilon^w_{t+1},$$

where $\{A_0, A_1, A_2, A_3, B_1, B_2, B_3, B_4\}$ are constants, the values of which are reported in Panel B of Table A23. As usual, we assume that the agent has preferences for early resolution of uncertainty ($\gamma > \frac{1}{\psi}$), so $B_2 < 0$.

The equilibrium log risk-free rate is

$$r^f_t = -E_t \left(\log \frac{\Lambda_{t+1}}{\Lambda_t} \right) - L_t \left(\frac{\Lambda_{t+1}}{\Lambda_t} \right) = -A_0 - A_1 x_t - A_2 u_t - A_3 w_t - \frac{1}{2} \left({B_1}^2 u_t + {B_2}^2 w_t + {B_3}^2 + {B_4}^2 \right),$$

or

$$r^f_t = -A_0 - \frac{1}{2} \left({B_3}^2 + {B_4}^2 \right) - A_1 x_t - \left(A_2 + \frac{{B_1}^2}{2} \right) u_t - \left(A_3 + \frac{{B_2}^2}{2} \right) w_t,$$

or

$$r^f_t = -A_0 - \frac{1}{2} \left({B_3}^2 + {B_4}^2 \right) + \frac{1}{\psi} x_t - \frac{1}{2} \left(\frac{\gamma - 1}{\psi} + \gamma \right) u_t - \frac{1}{2} \left(\frac{1}{\psi} - \gamma \right) \left(\frac{1}{\psi} - 1 \right) \left(\frac{\kappa}{1 - \kappa \phi^x} \right)^2 w_t.$$

Thus, the risk-free rate is positively associated with x, the predictable component of consumption growth, due to the intertemporal smoothing effect, and negatively associated with u, the conditional variance of the consumption growth shock, as the intertemporal smoothing effect is dominated by the precautionary savings effect. Finally, the sign of the relationship between the risk-free rate and w, the conditional variance of the consumption drift shock, depends on the value of the IES parameter: if $\psi > 1$, then the relationship is negative, as the precautionary savings effect dominates, whereas if $\psi < 1$, then the relationship is positive, as the intertemporal smoothing effect dominates.

To decompose the pricing kernel, we use the guess and verify method. In particular, guess an eigenfunction ϕ of the form

$$\phi(x, u, w) = e^{c_1 x + c_2 u + c_3 w}$$

where $\{c_1, c_2, c_3\}$ are constants. Then, the (one-period) eigenfunction problem can be written as

$$E_t \left[\exp(\log \frac{\Lambda_{t+1}}{\Lambda_t} + c_1 x_{t+1} + c_2 u_{t+1} + c_3 w_{t+1}) \right] = \exp(\beta + c_1 x_t + c_2 u_t + c_3 w_t)$$

which, exploiting the log-normality of the term inside the expectation, implies

$$E_t \left(\log \frac{\Lambda_{t+1}}{\Lambda_t} + c_1 x_{t+1} + c_2 u_{t+1} + c_3 w_{t+1} \right) + \frac{1}{2} var_t \left(\log \frac{\Lambda_{t+1}}{\Lambda_t} + c_1 x_{t+1} + c_2 u_{t+1} + c_3 w_{t+1} \right) = \beta + c_1 x_t + c_2 u_t + c_3 w_t.$$

After some algebra, matching terms yields

$$\beta = A_0 + c_2 (1 - \phi^u) \theta^u + c_3 (1 - \phi^w) \theta^w + \frac{1}{2} (B_3 + c_2 \sigma^u)^2 + \frac{1}{2} (B_4 + c_3 \sigma^w)^2$$

$$c_1 = A_1 + c_1 \phi^x$$

$$c_2 = A_2 + c_2 \phi^u + \frac{1}{2} B_1^2$$

$$c_3 = A_3 + c_3 \phi^w + \frac{1}{2} (B_2 + c_1)^2$$

so

$$c_1 = \frac{A_1}{1 - \phi^x} = -\frac{1}{\psi} \frac{1}{1 - \phi^x} < 0,$$

$$c_2 = \frac{A_2 + \frac{1}{2} B_1^2}{1 - \phi^u} = \frac{1}{2} \left(\frac{\gamma - 1}{\psi} + \gamma \right) \frac{1}{1 - \phi^u} > 0,$$

$$c_3 = \frac{A_3 + \frac{1}{2} (B_2 + c_1)^2}{1 - \phi^w} = \frac{\left(\frac{1}{\psi} - \gamma \right) \frac{\gamma - 1}{2} \left(\frac{\kappa}{1 - \kappa \phi^x} \right)^2 + \frac{1}{2} \left(\left(\frac{1}{\psi} - \gamma \right) \frac{\kappa}{1 - \kappa \phi^x} - \frac{1}{\psi} \frac{1}{1 - \phi^x} \right)^2}{1 - \phi^w} > 0,$$

where the sign for c_2 and c_3 is determined under the assumption that $\gamma > \frac{1}{\psi}$. The transitory component of the pricing kernel is

$$\Lambda^{\mathbb{T}}_t = e^{\beta t - c_1 x_t - c_2 u_t - c_3 w_t},$$

so the transitory SDF component is

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = e^{\beta + c_1(1-\phi^x)x_t - c_2(1-\phi^u)(\theta^u - u_t) - c_3(1-\phi^w)(\theta^w - w_t) - c_1\sqrt{w_t}\varepsilon_{t+1}^x - c_2\sigma^u\varepsilon_{t+1}^u - c_3\sigma^w\varepsilon_{t+1}^w},$$

and the permanent SDF component is

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = e^{A_0 + A_1x_t + A_2u_t + A_3w_t + B_1\sqrt{u_t}\varepsilon_{t+1}^c + B_2\sqrt{w_t}\varepsilon_{t+1}^x + B_3\varepsilon_{t+1}^u + B_4\varepsilon_{t+1}^w} \times$$
$$e^{-\beta - c_1(1-\phi^x)x_t + c_2(1-\phi^u)(\theta^u - u_t) + c_3(1-\phi^w)(\theta^w - w_t) + c_1\sqrt{w_t}\varepsilon_{t+1}^x + c_2\sigma^u\varepsilon_{t+1}^u + c_3\sigma^w\varepsilon_{t+1}^w},$$

or, after some algebra,

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = e^{-(1/2)(B_3 + c_2\sigma^u)^2 - (1/2)(B_4 + c_3\sigma_w)^2 - (1/2)B_1^2u_t - (1/2)(B_2 + c_1)^2w_t} \times$$
$$e^{B_1\sqrt{u_t}\varepsilon_{t+1}^c + (B_2 + c_1)\sqrt{w_t}\varepsilon_{t+1}^x + (B_3 + c_2\sigma^u)\varepsilon_{t+1}^u + (B_4 + c_3\sigma^w)\varepsilon_{t+1}^w}.$$

In summary, both SDF components are exposed to the consumption drift innovation ε^x, the consumption growth variance innovation ε^u, and the consumption drift variance innovation ε^w, but only the permanent SDF component is exposed to the consumption growth innovation ε^c. As a result, overall SDF and the permanent SDF component have identical loadings on the consumption growth shock. However, the dependence on the rest of the innnovations depends on the agent's preferences regarding the resolutions of uncertainty. If the agent prefers early resolution ($\gamma > \frac{1}{\psi}$), we have $c_1 < 0$, $c_2 > 0$ and $c_3 > 0$.

We can start with exposure to consumption drift shocks. Since $B_2 < 0$, $c_1 < 0$ implies that the permanent SDF component is more sensitive to consumption drift shocks than the total SDF, while the transitory SDF component has the opposite sign. For example, a negative consumption drift shock ($\varepsilon^x < 0$) is associated with an increase of the agent's overall SDF and its permanent component and a decline of its transitory component. This is because the long-run effect of a consumption drift innovation in the pricing kernel (captured by the permanent SDF component) is higher than its short-run effect (captured by the overall SDF). Intuitively, a negative consumption drift shock lowers marginal utility in the long run both through an immediate decline in the continuation utility (reflected in the overall SDF) and through the cumulative effect of a persistent reduction in x, which is equal to $-\frac{1}{\psi}\sum_{j=0}^{\infty}(\phi^x)^j\sqrt{\theta^w} = -\frac{1}{\psi}\frac{1}{1-\phi^x}\sqrt{\theta^w} = c_1\sqrt{\theta^w}$.

As regards the two variance shocks, whether long-run marginal utility reacts more or less than short-run marginal utility depends on the sign of B_3 and B_4. If $\gamma > 1$, i.e. the agent is more risk-averse than a log utility investor, then $B_3 > 0$ and $B_4 > 0$, so short-run marginal utility increases upon realization of any positive variance shock. Thus, $c_2 > 0$ and $c_3 > 0$ imply that long-run marginal utility reacts more than short-run marginal utility: when either $\varepsilon^u > 0$ or $\varepsilon^w > 0$, the permanent SDF component increases more than total SDF, with the transitory SDF component declining. On the other hand, if $\gamma < 1$, then $B_3 < 0$ and $B_4 < 0$, in which case short-run marginal utility declines upon realization of any positive variance shock. As a result, $c_2 > 0$ and $c_3 > 0$ imply that long-run marginal utility reacts less than short-run marginal utility: when either $\varepsilon^u > 0$ or $\varepsilon^w > 0$, the permanent SDF component declines less than total SDF, as the transitory SDF component also falls.

Conditional SDF entropy is

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \frac{1}{2}\left({B_1}^2u_t + {B_2}^2w_t + {B_3}^2 + {B_4}^2\right),$$

whereas the conditional entropy of the permanent SDF component is

$$L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \frac{1}{2}\left({B_1}^2u_t + (B_2 + c_1)^2w_t + (B_3 + c_2\sigma^u)^2 + (B_4 + c_3\sigma^w)^2\right).$$

For conditional SDF entropy to be identical across countries, it is sufficient that the conditional variances u and w are identical across countries, that $B_1 = B_1^*$ and $B_2 = B_2^*$ (i.e. that $\gamma = \gamma^*$ and $\left(\frac{1}{\psi} - \gamma\right)\frac{\delta}{1-\delta\phi^x} = \left(\frac{1}{\psi^*} - \gamma^*\right)\frac{\delta^*}{1-\delta^*\phi^{x,*}}$), and that $(B_3 + c_2\sigma^u)^2 + (B_4 + c_3\sigma^w)^2 = (B_3^* + c_2^*\sigma^{u,*})^2 + (B_4^* + c_3^*\sigma^{w,*})^2$. For the conditional entropy of the permanent SDF component to be identical across countries, we need $B_2 + c_1 = B_2^* + c_1^*$ instead of $B_2 = B_2^*$. Therefore, we will have non-identical SDF entropy and identical entropy of the permanent SDF component across countries if $B_2 + c_1 = B_2^* + c_1^*$ and $c_1 - c_1^* = B_2^* - B_2 \neq 0$. For example, those conditions are satisfied if $\gamma = \gamma^*$, $\delta = \delta^*$ and $\psi = \psi^*$, but $\phi^x \neq \phi^{x,*}$ such that $(1 - \delta\phi^x)(1 - \delta\phi^{x,*}) = \delta^2(1 - \gamma\psi)(1 - \phi^x)(1 - \phi^{x,*})$.

Finally, the term premium, in local currency terms, is

$$E_t\left[rx_{t+1}^{(\infty)}\right] = \frac{1}{2}\left({B_2}^2 - (B_2 + c_1)^2\right)w_t + \frac{1}{2}\left({B_3}^2 - (B_3 + c_2\sigma^u)^2\right) + \frac{1}{2}\left({B_4}^2 - (B_4 + c_3\sigma^w)^2\right).$$

Following the discussion above, if $\gamma > \frac{1}{\psi}$ (in which case $B_2 < 0$), then the conditional term premium is negatively associated with w, the variance of the consumption growth drift. This is because negative consumption drift shocks increase long-run marginal utility more than they increase short-run marginal utility, so long-term bonds hedge long-run risk, as their price increases upon realization of negative consumption drift shocks. Therefore, the higher the conditional volatility of those shocks, the more attractive long-term bonds are as a hedging asset, and the lower risk premium they earn.

Consider a symmetric Model with country-specific shocks. The quantity of risk is governed by u_t, the volatility of consumption growth, and w_t, the volatility of expected consumption growth. Both of these forces feed into the quantity of permanent risk unless $B_1 = B_2 - C_1^\infty = 0$. Thus, in a symmetric LRR model (i.e., when countries share the same parameters) with country-specific shocks and heteroskedasticity, Condition 1 holds only if the model parameters satisfy the following restriction: $\gamma = 0 = \frac{1}{\psi}$, implying that the pricing kernel is constant and the investor is risk-neutral. In this case, the model counterfactually replicates the U.I.P. condition in the short-run.

In the long-run, U.I.P. is violated for risk-related innovations because the long-run loadings of the level of the exchange rate on (u_t, w_t) do not line up with the loadings of the long rates:

$$\sum_{i=1}^{\infty} E_t[\Delta s_{t+i}] = \frac{A_1}{1-\phi_x}(x_t - x_t^*) + \frac{A_2}{1-\phi_u}(u_t - u_t^*) + \frac{A_3}{1-\phi_w}(w_t - w_t^*) \neq C_1^\infty(x_t^* - x_t) + C_2^\infty(u_t^* - u_t) + C_3^\infty(w_t^* - w_t),$$

because $C_2^\infty \neq -\frac{A_2}{1-\phi_u}$ and $C_3^\infty \neq -\frac{A_3}{1-\phi_w}$.

Next, consider an asymmetric model with common shocks. Thus, a natural extension to the model would feature common volatility processes, such that $u_t = u_t^*$ and $w_t = w_t^*$, relieving the strong parameter restriction above (see Colacito, Croce, Gavazzoni, and Ready, 2017, for a multi-country LRR model with common shocks). Condition 1 again tells us where to introduce heterogeneity in a future version of this model. For the conditional entropy of the permanent SDF component to be identical across countries, we need the following parameter restriction: $B_1 = B_1^*$ and $B_2 - C_1^\infty = B_2^* - C_1^{*,\infty}$. In this case, we have different SDF entropy to generate carry risk premia at the short end of the curve, but the same entropy of the permanent SDF component across countries if $B_2 - C_1^\infty = B_2^* - C_1^{*,\infty}$ and $C_1^\infty - C_1^{*,\infty} = B_2^* - B_2 \neq 0$.

These restrictions have bite. Consider an example with only heterogeneity in the persistence of the shocks. Our conditions are satisfied if $\gamma = \gamma^*$, $\delta = \delta^*$ and $\psi = \psi^*$, but $\phi^x \neq \phi^{x,*}$ such that $(1-\delta\phi^x)(1-\delta\phi^{x,*}) = \delta^2(1-\gamma\psi)(1-\phi^x)(1-\phi^{x,*})$. That restriction cannot be satisfied when agents have a preference for early resolution of uncertainty ($\gamma\psi > 1$), as is invariably assumed in LRR models. The constant component of the entropy above adds even more parameter restrictions.

E.3 Disasters Model

In the Farhi and Gabaix (2016) version of the Gabaix (2012) and Wachter (2013) rare disasters model with time-varying disaster intensity, the SDF has the following law of motion:

$$\frac{\Lambda_{t+1}}{\Lambda_t} = \frac{\Lambda_{t+1}^*}{\Lambda_t^*}\frac{\omega_{t+1}}{\omega_t}\frac{1+Ax_{t+1}}{1+Ax_t},$$

where Λ^* denotes the global component of marginal utility:

$$\frac{\Lambda_{t+1}^*}{\Lambda_t^*} = e^{-R} \times \begin{cases} 1, \text{if there is no disaster at } t+1 \\ B_{t+1}^{-\gamma}, \text{if there is a disaster at } t+1, \end{cases}$$

ω_{t+1} denotes the country-specific productivity:

$$\frac{\omega_{t+1}}{\omega_t} = e^{g_\omega} \times \begin{cases} 1, \text{if there is no disaster at } t+1 \\ F_{t+1}, \text{if there is a disaster at } t+1, \end{cases}$$

x_t is the (scaled by e^{-h_*}) time-varying component of the resilience of the country (with persistence ϕ_H) and $A > 0$ depends on the model parameters, one of which is the investment depreciation rate λ. After some algebra, we obtain the following expression for conditional entropy:

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \log(1+H_t) - p_t E_t^D\left[\log(B_{t+1}^{\gamma}F_{t+1})\right] + L_t\left(\frac{1+Ax_{t+1}}{1+Ax_t}\right),$$

where p_t is the conditional probability of a disaster occurring next period and E_t^D is the period t expectation conditional on a disaster occurring next period. The equilibrium log risk-free rate is

$$r_t^f = (R - g_\omega - h_*) + \log\left(\frac{1 + Ax_t}{1 + (Ae^{-\phi_H} + 1)x_t}\right).$$

Thus, the risk-free rate is decreasing in x and, thus, in the resilience of the country. Again, high interest rate countries correspond to low volatility SDFs.

Finally, the conditional entropy of the permanent SDF component is

$$L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \log(1 + H_t) - p_t E_t^D\left[\log(B_{t+1}^{\gamma} F_{t+1})\right] + L_t\left(\frac{c + x_{t+1}}{c + x_t}\right).$$

In the Farhi and Gabaix (2016) rare disasters model, the SDF has law of motion

$$\frac{\Lambda_{t+1}}{\Lambda_t} = \frac{\Lambda_{t+1}^*}{\Lambda_t^*}\frac{\omega_{t+1}}{\omega_t}\frac{1 + Ax_{t+1}}{1 + Ax_t},$$

where

$$\frac{\Lambda_{t+1}^*}{\Lambda_t^*} = e^{-R} \times \begin{cases} 1, \text{if there no disaster at } t+1 \\ B_{t+1}^{-\gamma}, \text{if there is a disaster at } t+1 \end{cases}$$

is the global numeraire SDF,

$$\frac{\omega_{t+1}}{\omega_t} = e^{g_\omega} \times \begin{cases} 1, \text{if there is no disaster at } t+1 \\ F_{t+1}, \text{if there is a disaster at } t+1 \end{cases}$$

is the productivity growth of the country, and x is defined as $x_t \equiv e^{-h_*}\hat{H}_t$, where $\hat{H}$ is the time-varying component of the resilience of the country, to be discussed below. Finally, $A \equiv \frac{e^{-R-\lambda+g_\omega+h_*}}{1-e^{-R-\lambda+g_\omega+h_*-\phi_H}}$, where λ is the investment depreciation rate, and $h_* \equiv \log(1 + H_*)$. Finally, we assume that $R + \lambda - g_\omega - h_* > 0$, so $A > 0$.

Resilience is defined as

$$H_t = H_* + \hat{H}_t = p_t E_t^D\left[B_{t+1}^{\gamma} F_{t+1} - 1\right],$$

where p_t is the conditional probability of a disaster occurring next period and E_t^D is the period t expectation conditional on a disaster occurring next period. The time-varying component of resilience has law of motion

$$\hat{H}_{t+1} = \frac{1 + H_*}{1 + H_t} e^{-\phi_H}\hat{H}_t + \varepsilon_{t+1}^H,$$

with the conditional expectation of ε^H being zero independently of the realization of a disaster. As a result, the conditional expectation of x is

$$E_t(x_{t+1}) = e^{-\phi_H}\frac{x_t}{1 + x_t}.$$

The equilibrium log risk-free rate is

$$r_t^f = -\log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = (R - g_\omega - h_*) + \log\left(\frac{1 + Ax_t}{1 + (Ae^{-\phi_H} + 1)x_t}\right),$$

so it is decreasing in x.

To decompose the pricing kernel, we use the guess and verify method. In particular, guess an eigenfunction ϕ of the form

$$\phi(x) = \frac{c + x}{1 + Ax},$$

where c is a constant.[19] Then, the (one-period) eigenfunction problem can be written as

$$E_t\left[\frac{\Lambda_{t+1}^*}{\Lambda_t^*}\frac{\omega_{t+1}}{\omega_t}\frac{1 + Ax_{t+1}}{1 + Ax_t}\frac{c + x_{t+1}}{1 + Ax_{t+1}}\right] = e^{\beta}\frac{c + x_t}{1 + Ax_t}$$

[19]It can be shown that conjecturing the more general eigenfunction $\phi(x) = \frac{c_1 + c_2 x}{1 + Ax}$, where c_1 and c_2 are constants, leads to same SDF decomposition as the one derived below.

which yields

$$E_t\left[\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\frac{\omega_{t+1}}{\omega_t}\right]E_t\left[\frac{1+Ax_{t+1}}{1+Ax_t}\frac{c+x_{t+1}}{1+Ax_{t+1}}\right]=e^{\beta}\frac{c+x_t}{1+Ax_t}$$

so

$$e^{-R+g_\omega+h_*}(1+x_t)E_t\left[\frac{1+Ax_{t+1}}{1+Ax_t}\frac{c+x_{t+1}}{1+Ax_{t+1}}\right]=e^{\beta}\frac{c+x_t}{1+Ax_t}.$$

The expression above becomes:

$$e^{-R+g_\omega+h_*}(1+x_t)E_t\left[c+x_{t+1}\right]=e^{\beta}(c+x_t),$$

so, plugging in the expression for the conditional expectation of x, we get

$$e^{-R+g_\omega+h_*}(1+x_t)\left(c+e^{-\phi_H}\frac{x_t}{1+x_t}\right)=e^{\beta}(c+x_t),$$

which yields

$$\beta=-R+g_\omega+h_*$$

and

$$c=1-e^{-\phi_H}.$$

The lower bound of x is $e^{-\phi_H}-1$, so $c+x_t>0$ for all t; thus, the conjectured eigenfunction is strictly positive, as required. The transitory component of the pricing kernel is

$$\Lambda^{\mathbb{T}}_t=e^{\beta t}\frac{1+Ax_t}{c+x_t}$$

so the transitory SDF component is

$$\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}=e^{\beta}\frac{1+Ax_{t+1}}{c+x_{t+1}}\frac{c+x_t}{1+Ax_t}=e^{-R+g_\omega+h_*}\frac{1+Ax_{t+1}}{1+Ax_t}\frac{c+x_t}{c+x_{t+1}}$$

and the permanent SDF component is

$$\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}=\frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_t}\right)^{-1}\quad=\quad e^{R-g_\omega-h_*}\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\frac{\omega_{t+1}}{\omega_t}\frac{c+x_{t+1}}{c+x_t}.$$

The transitory SDF component is only exposed to resilience shocks (ε^H), but not to disaster risk; the entirety of the disaster risk for marginal utility is reflected in the permanent SDF component, as disasters permanently affect the future level of marginal utility.

We can now calculate the conditional entropy of the SDF and its components. It holds that

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)=\log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)-E_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right),$$

so we can write

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)=L_t\left(\frac{\Lambda^*_{t+1}}{\Lambda^*_t}\frac{\omega_{t+1}}{\omega_t}\right)+L_t\left(\frac{1+Ax_{t+1}}{1+Ax_t}\right).$$

After some algebra, we get

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)=\log(1+H_t)-p_tE^D_t\left[\log(B^{\gamma}_{t+1}F_{t+1})\right]+L_t\left(\frac{1+Ax_{t+1}}{1+Ax_t}\right).$$

Similarly, the conditional entropy of the permanent SDF component is

$$L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right)=\log(1+H_t)-p_tE^D_t\left[\log(B^{\gamma}_{t+1}F_{t+1})\right]+L_t\left(\frac{c+x_{t+1}}{c+x_t}\right).$$

Therefore, the conditional term premium, in local currency terms, is

$$E_t\left[rx^{(\infty)}_{t+1}\right]=L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right)-L_t\left(\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_t}\right)=L_t\left(\frac{1+Ax_{t+1}}{1+Ax_t}\right)-L_t\left(\frac{c+x_{t+1}}{c+x_t}\right).$$

First, we consider a version of the model in which the parameters are the same in each country, but the shocks are country-specific. The time-varying disaster risk directly, driven partly by some country-specific shocks, affects the total quantity of permanent risk, thus violating Condition 1, unless the disaster intensity is constant (p_t, x_t, H_t are constant), and hence all risk premia are constant. Second, we consider a version of the disaster model with common shocks, but asymmetric exposures. It is possible to introduce differences across countries that produce differences in carry trade portfolio returns at the short but not at the long end of the curve, but the heterogeneity is clearly restricted by Condition 1 to the parameters R, λ or g_w.

Table A22 in the main text summarizes the results for all three models. First, among the models we consider with only country-specific shocks in which all countries have the same parameters, only the external habit model can satisfy Condition 1. In the habit model, Condition 1 is trivially satisfied because the quantity of permanent risk is constant. In this external habit model, the time-variation in the price of risk driven by the surplus-consumption ratio is purely transitory in nature. Country-specific consumption growth shocks do enter the permanent component of the pricing kernel, but the quantity of permanent consumption risk is constant. In the other models, the variation in the price of risk invariably has permanent effects and, hence, Condition 1 cannot be satisfied. Second, in models we consider with only common shocks and heterogeneity in the parameters, Condition 1 imposes tight parametric restrictions on the types of heterogeneity that can be allowed. For example, in the external habit and disaster model, we cannot have heterogeneity in the coefficient of relative risk aversion: $\gamma\sigma^2$ needs to be constant across countries in the habits model.

F Theoretical Background and Proofs of Preference-Free Results

This section starts with a review of the Hansen and Scheinkman (2009) results and their link to the Alvarez and Jermann (2005) decomposition used in the main text. Then, we report our theoretical results on bond and currency returns in two special cases: the case of a Gaussian economy and the case of an economy with no permanent pricing kernel shocks. The section concludes with the proofs of all the theoretical results in the main body of the paper. To make the paper self-contained, we reproduce here some proofs of intermediary results already in the literature, notably in Alvarez and Jermann (2005).

F.1 Existence and Uniqueness of Multiplicative Decomposition of the pricing kernel

Consider a continuous-time, right continuous with left limits, strong Markov process X and the filtration $\mathcal{F}$ generated by the past values of X, completed by the null sets. In the case of infinite-state spaces, X is restricted to be a semimartingale, so it can be represented as the sum of a continuous process X^c and a pure jump process X^j. The pricing kernel process Λ is a strictly positive process, adapted to $\mathcal{F}$, for which it holds that the time t price of any payoff Π_s realized at time s ($s \geq t$) is given by

$$P_t(\Pi_s) = E\left[\frac{\Lambda_s}{\Lambda_t}\Pi_s|\mathcal{F}_t\right].$$

The pricing kernel process also satisfies $\Lambda_0 = 1$. Hansen and Scheinkman (2009) show that Λ is a multiplicative functional and establish the connection between the multiplicative property of the pricing kernel process and the semigroup property of pricing operators $\mathbb{M}$.[20] In particular, consider the family of operators $\mathbb{M}$ described by

$$\mathbb{M}_t\psi(x) = E\left[\Lambda_t\psi(X_t)|X_0 = x\right]$$

where $\psi(X_t)$ is a random payoff at t that depends solely on the Markov state at t. The family of linear pricing operators $\mathbb{M}$ satisfies $\mathbb{M}_0 = \mathbb{I}$ and $\mathbb{M}_{t+u}\psi(x) = \mathbb{M}_t\psi(x)\mathbb{M}_u\psi(x)$ and, thus, defines a semigroup, called pricing semigroup.

Further, Hansen and Scheinkman (2009) show that Λ can be decomposed as

$$\Lambda_t = e^{\beta t}\frac{\phi(X_0)}{\phi(X_t)}\Lambda_t^{\mathbb{P}}$$

where $\Lambda^{\mathbb{P}}$ is a multiplicative functional and a local martingale, ϕ is a principal (i.e. strictly positive) eigenfunction of the extended generator of $\mathbb{M}$ and β is the corresponding eigenvalue (typically negative).[21] If, furthermore, $\Lambda^{\mathbb{P}}$ is martingale, then the eigenpair (β, ϕ) also solves the principal eigenvalue problem:[22]

$$\mathbb{M}_t\phi(x) = E\left[\Lambda_t\phi(X_t)|X_0 = x\right] = e^{\beta t}\phi(x).$$

Conversely, if the expression above holds for a strictly positive ϕ and $\mathbb{M}_t\phi$ is well-defined for $t \geq 0$, then $\Lambda^{\mathbb{P}}$ is a martingale. Thus, a strictly positive solution to the eigenvalue problem above implies a decomposition

$$\Lambda_t = e^{\beta t}\frac{\phi(X_0)}{\phi(X_t)}\Lambda_t^{\mathbb{P}}$$

where $\Lambda^{\mathbb{P}}$ is guaranteed to be a martingale. The decomposition above implies that the one-period SDF is given by

$$M_{t+1} = \frac{\Lambda_{t+1}}{\Lambda_t} = e^{\beta}\frac{\phi(X_t)}{\phi(X_{t+1})}\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}$$

and satisfies

$$E\left[M_{t+1}\phi(X_{t+1})|X_t = x\right] = e^{\beta t}\phi(x).$$

Hansen and Scheinkman (2009) provide sufficient conditions for the existence of a solution to the principal eigenfunction problem and, thus, for the existence of the aforementioned pricing kernel decomposition. Notably, multiple solutions may exist,

[20] A functional Λ is multiplicative if it satisfies $\Lambda_0 = 1$ and $\Lambda_{t+u} = \Lambda_t\Lambda_u(\theta_t)$, where θ_t is a shift operator that moves the time subscript of the relevant Markov process forward by t periods. Products of multiplicative functionals are multiplicative functionals. The multiplicative property of the pricing kernel arises from the requirement for consistency of pricing across different time horizons.

[21] The extended generator of a multiplicative functional Λ is formally defined in Hansen and Scheinkman (2009) and, intuitively, assigns to a Borel function ψ a Borel function ξ such that $\Lambda_t\xi(X_t)$ is the expected time derivative of $\Lambda_t\psi(X_t)$.

[22] Since $\Lambda^{\mathbb{P}}$ is a local martingale bounded from below, it is a supermartingale. For $\Lambda^{\mathbb{P}}$ to be a martingale, additional conditions need to hold, as discussed in Appendix C of Hansen and Scheinkman (2009).

so the pricing kernel decomposition above is generally not unique. However, if the state space is finite and the Markov chain is irreducible, then Perron-Frobenious theory implies that there is a unique principal eigenvector (up to scaling), and thus a unique pricing kernel decomposition. Although multiple solutions typically exist, Hansen and Scheinkman (2009) show that the only (up to scaling) principal eigenfunction of interest for long-run pricing is the one associated with the smallest eigenvalue, as the multiplicity of solutions is eliminated by the requirement for stochastic stability of the Markov process X. In particular, only this solution ensures that the process X remains stationary and Harris recurrent under the probability measure implied by the martingale $\Lambda^{\mathbb{P}}$.

Finally, Hansen and Scheinkman (2009) show that the aforementioned pricing kernel decomposition can be useful in approximating the prices of long-maturity zero-coupon bonds. In particular, the time t price of a bond with maturity $t+k$ is given by

$$P_t^{(k)} = E\left[\frac{\Lambda_{t+k}}{\Lambda_t}|X_t = x\right] = e^{\beta k} E^{\mathbb{P}}\left[\frac{1}{\phi(X_{t+k})}|X_t = x\right]\phi(x) \approx e^{\beta k} E^{\mathbb{P}}\left[\frac{1}{\phi(X_{t+k})}\right]\phi(x)$$

where $E^{\mathbb{P}}$ is the expectation under the probability measure implied by the martingale $\Lambda^{\mathbb{P}}$ and the right-hand-side approximation becomes arbitrarily accurate as $k \to \infty$. Thus, in the limit of arbitrarily large maturity, the price of the zero-coupon bond depends on the current state solely through $\phi(x)$ and not through the expectation of the transitory component. Notably, this implies that the relevant ϕ is the one that ensures that X remains stationary under the probability measure implied by $\Lambda^{\mathbb{P}}$, i.e. the unique principal eigenfunction that implies stochastic stability for X, and β is the corresponding eigenvalue.

Indeed, Alvarez and Jermann (2005) *construct* a pricing kernel decomposition by considering a constant $\hat{\beta}$ that satisfies

$$0 < \lim_{k\to\infty} \frac{P_t^{(k)}}{\hat{\beta}^k} < \infty$$

and *defining* the transitory pricing kernel component as

$$\Lambda_t^{\mathbb{T}} = \lim_{k\to\infty} \frac{\hat{\beta}^{t+k}}{P_t^{(k)}} < \infty.$$

In contrast to Hansen and Scheinkman (2009), the decomposition of Alvarez and Jermann (2005) is constructive and not unique, as their Assumptions 1 and 2 do not preclude the existence of alternative pricing kernel decompositions to a martingale and a transitory component. Note that the Alvarez and Jermann (2005) decomposition implies that $\hat{\beta} = e^{\beta}$, where β is the smallest eigenvalue associated with a principal eigenfunction in the Hansen and Scheinkman (2009) eigenfunction problem.

F.2 Long-Horizon U.I.P. in Gaussian Economy

The long-horizon U.I.P. condition states that the expected return over k periods on a foreign bond, once converted into domestic currency, is equal to the expected return on a domestic bond over the same investment horizon.[23] The per period log risk premium on a long position in foreign currency over k periods consists of the yield spread minus the per period expected rate of depreciation over those k periods:

$$E_t[rx^{FX}_{t\to t+k}] = y_t^{(k),*} - y_t^{(k)} - \frac{1}{k}E_t[\Delta s_{t\to t+k}]. \tag{21}$$

The long-horizon U.I.P condition states that this risk premium is zero. As is well-known, this risk premium is the sum of a term premium and future currency risk premia. To see that, start from the definition of the one-period currency risk premium: $E_t\left[\Delta s_{t\to t+1}\right] = r_t^{f,*} - r_t^f - E_t\left[rx_{t+1}^{FX}\right]$. Summing up over k periods leads to:

$$E_t[\Delta s_{t\to t+k}] \quad = \quad E_t\left[\sum_{j=1}^{k}\left(r_{t+j-1}^{f,*} - r_{t+j-1}^{f}\right)\right] - E_t\left[\sum_{j=1}^{k} rx_{t+j}^{FX}\right]. \tag{22}$$

From Equations (21) and (22), it follows that the log currency risk premium over k periods is given by:

$$E_t[rx^{FX}_{t\to t+k}] \quad = \quad (y_t^{(k),*} - y_t^{(k)}) + \frac{1}{k}\sum_{j=1}^{k} E_t\left(r_{t+j-1}^{f} - r_{t+j-1}^{f,*}\right) + \frac{1}{k}\sum_{j=1}^{k} E_t(rx_{t+j}^{FX}). \tag{23}$$

The first two terms measure the deviations from the expectations hypothesis over the holding period k, whereas the last term measures the deviations from short-run U.I.P. over the k periods. We can use a multi-horizon version of Equation (7) to show that

[23]Chinn and Meredith (2004) document some time-series evidence that supports a conditional version of UIP at longer holding periods, while Boudoukh, Richardson, and Whitelaw (2016) show that past forward rate differences predict future changes in exchange rates.

the currency risk premium over k periods depends on conditional SDF entropy:

$$E_t[rx^{FX}_{t\to t+k}] = \frac{1}{k}\left[L_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda^*_{t+k}}{\Lambda^*_t}\right)\right]. \tag{24}$$

The expression above states that only differences in k-period conditional SDF entropy give rise to long-run deviations from U.I.P. Therefore, the risk premium on a multi-period long position in foreign currency depends on how quickly SDF entropy builds up domestically and abroad over the holding period.[24] If the pricing kernel is conditionally Gaussian over horizon k, the k-horizon foreign currency risk premium satisfies:

$$E_t[rx^{FX}_{t\to t+k}] = \frac{1}{2k}\left[var_t\left(\log\frac{\Lambda_{t+k}}{\Lambda_t}\right) - var_t\left(\log\frac{\Lambda^*_{t+k}}{\Lambda^*_t}\right)\right].$$

Let us assume that the variance of the one-period SDF is constant. The annualized variance of the increase in the log SDF can be expressed as follows:

$$\frac{var\left(\log\Lambda_{t+k}/\Lambda_t\right)}{kvar\left(\Lambda_{t+1}/\Lambda_t\right)} = 1 + 2\sum_{j=1}^{k-1}\left(1-\frac{j}{k}\right)\rho_j,$$

where ρ_j denotes the j-th autocorrelation (Cochrane, 1988).[25] In the special case where the domestic and foreign countries share the same one-period volatility of the innovations, this expression for the long-run currency risk premium becomes:

$$E[rx^{FX}_{t\to t+k}] = var\left(\Delta\log\Lambda_{t+1}\right)\left[\sum_{j=1}^{k-1}\left(1-\frac{j}{k}\right)(\rho_j - \rho^*_j)\right].$$

This is the Bartlett kernel estimate with window k of the spread in the spectral density of the log SDF at zero, which measures the size of the permanent component of the SDF. More positive autocorrelation in the domestic than in the foreign pricing kernel tends to create long-term yields that are lower at home than abroad, once expressed in the same currency. The difference in yields, converted in the same units, is governed by a horse race between the speed of mean reversion in the pricing kernel at home and abroad.

To develop some intuition for the long run, we consider the limit behavior of the foreign currency risk premium when $k\to\infty$. In the long run, the currency risk premium over many periods converges to the difference in the size of the random walk components:

$$\begin{aligned}
\lim_{k\to\infty} E[rx^{FX}_{t\to t+k}] &= \frac{1}{2}var\left(\Delta\log\Lambda_{t+1}\right)\lim_{k\to\infty}\left[1+2\sum_{j=1}^{\infty}\rho_j\right] - \frac{1}{2}var\left(\Delta\log\Lambda_{t+1}\right)\lim_{k\to\infty}\left[1+2\sum_{j=1}^{\infty}\rho^*_j\right] \\
&= \frac{1}{2}\left[S_{\Delta\log\Lambda_{t+1}} - S_{\Delta\log\Lambda^*_{t+1}}\right],
\end{aligned}$$

where S denotes the spectral density. The last step follows from the definition of the spectral density (see Cochrane, 1988). If the log of the exchange rate ($\log S_t$) is stationary, then the log of the foreign ($\log\Lambda^*_t$) and domestic pricing kernels ($\log\Lambda_t$) are cointegrated with co-integrating vector $(1,-1)$ and hence share the same stochastic trend component. This in turn implies that they have the same spectral density evaluated at zero. As a result, exchange rate stationarity implies that the long-run currency risk premium goes to zero.

[24]To develop some intuition, we consider a Gaussian example in the Appendix. In the special case where the domestic and foreign countries share the same one-period volatility of the innovations, this expression for the long-run currency risk premium becomes:

$$E[rx^{FX}_{t\to t+k}] = var\left(\Delta\log\Lambda_{t+1}\right)\left[\sum_{j=1}^{k-1}\left(1-\frac{j}{k}\right)(\rho_j - \rho^*_j)\right].$$

This is the Bartlett kernel estimate with window k of the spread in the spectral density of the log SDF at zero, which measures the size of the permanent component of the SDF.

[25]Cochrane (1988) uses these per period variances of the log changes in GDP to measure the size of the random walk component in GDP.

F.3 Economy without Permanent Innovations

Consider the special case in which the pricing kernel is not subject to permanent innovations, i.e., $\lim_{k\to\infty} \frac{E_{t+1}[\Lambda_{t+k}]}{E_t[\Lambda_{t+k}]} = 1$. For example, the Markovian environment considered by Ross (2015) to derive his recovery theorem satisfies this condition. Building on this work, Martin and Ross (2013) derive closed-form expressions for bond returns in a similar environment. Alvarez and Jermann (2005) show that this case has clear implications for domestic returns: if the pricing kernel has no permanent innovations, then the term premium on an infinite maturity bond is the largest risk premium in the economy.[26]

The absence of permanent innovations also has a strong implication for the term structure of the carry trade risk premia. When the pricing kernels do not have permanent innovations, the foreign term premium in dollars equals the domestic term premium:

$$E_t\left[rx_{t+1}^{(\infty),*}\right] + (f_t - s_t) - E_t[\Delta s_{t+1}] = E_t\left[rx_{t+1}^{(\infty)}\right].$$

The proof here is straightforward. In general, the foreign currency risk premium is equal to the difference in entropy. In the absence of permanent innovations, the term premium is equal to the entropy of the pricing kernel, so the result follows. More interestingly, a much stronger result holds in this case. Not only are the risk premia identical, but the returns on the foreign bond position are the same as those on the domestic bond position state-by-state, because the foreign bond position automatically hedges the currency risk exposure. As already noted, if the domestic and foreign pricing kernels have no permanent innovations, then the one-period returns on the longest maturity foreign bonds in domestic currency are identical to the domestic ones:

$$\lim_{k\to\infty} \frac{S_t}{S_{t+1}} \frac{R_{t+1}^{(k),*}}{R_{t+1}^{(k)}} = 1.$$

In this class of economies, the returns on long-term bonds expressed in domestic currency are equalized:

$$\lim_{k\to\infty} rx_{t+1}^{(k),*} + (f_t - s_t) - \Delta s_{t+1} = rx_{t+1}^{(k)}.$$

In countries that experience higher marginal utility growth, the domestic currency appreciates but is exactly offset by the capital loss on the bond. For example, in a representative agent economy, when the log of aggregate consumption drops more below trend at home than abroad, the domestic currency appreciates, but the real interest rate increases, because the representative agent is eager to smooth consumption. The foreign bond position automatically hedges the currency exposure.

Alvarez and Jermann (2005) propose the following example of an economy without permanent shocks: a representative agent economy with power utility investors in which the log of aggregate consumption is a trend-stationary process with normal innovations. In particular, consider the following pricing kernel (Alvarez and Jermann, 2005):

$$\log \Lambda_t = \sum_{i=0}^{\infty} \alpha_i \epsilon_{t-i} + \beta \log t,$$

with $\epsilon \sim N(0, \sigma^2), \alpha_0 = 1$. If $\lim_{k\to\infty} \alpha_k^2 = 0$, then the pricing kernel has no permanent component. The foreign pricing kernel is defined similarly.

In the model, the term premium equals one half of the SDF variance: $E_t\left(rx_{t+1}^{(\infty)}\right) = \sigma^2/2$, the highest possible risk premium in this economy, as the returns on the long bond are perfectly negatively correlated with the stochastic discount factor. When marginal utility is temporarily high, the representative agent would like to borrow, driving up interest rates and lowering the price of the long-term bond.

In this economy, the foreign term premium in dollars is identical to the domestic term premium:

$$E_t\left[rx_{t+1}^{(\infty),*}\right] + (f_t - s_t) - E_t[\Delta s_{t+1}] = \frac{1}{2}\sigma^2 = E_t\left[rx_{t+1}^{(\infty)}\right].$$

This result is straightforward to establish: recall that the currency risk premium is equal to the half of the difference between the domestic and the foreign SDF variance. Currencies with a high local currency term premium (high σ^2) also have an offsetting negative currency risk premium, while those with a small term premium have a large currency risk premium. Hence, U.S. investors receive the same dollar premium on foreign as on domestic bonds. There is no point in chasing high term premia around the world, at least not in economies with only temporary innovations to the pricing kernel. Currencies with the highest local term premia also have the lowest (i.e., most negative) currency risk premia.

[26] If there are no permanent innovations to the pricing kernel, then the return on the bond with the longest maturity equals the inverse of the SDF: $\lim_{k\to\infty} R_{t+1}^{(k)} = \Lambda_t/\Lambda_{t+1}$. High marginal utility growth translates into higher yields on long maturity bonds and low long bond returns, and vice-versa.

G Additional Implications

We end this Appendix with two additional implications of our main results that can further help build the next generation of international finance models and guide future empirical work.

G.1 A Lower Bound on Cross-Country Correlations of the Permanent SDF Components

Brandt, Cochrane, and Santa-Clara (2006) show that the combination of relatively smooth exchange rates and much more volatile SDFs implies that SDFs are very highly correlated across countries. A 10% volatility in exchange rate changes and a volatility of marginal utility growth rates of 50% imply a correlation of at least 0.98. We do not interpret the correlation of SDFs or their components in terms of cross-country risk-sharing, because doing so requires additional assumptions. The nature and magnitude of international risk sharing is an important and open question in macroeconomics (see, for example, Cole and Obstfeld (1991); Wincoop (1994); Lewis (2000); Gourinchas and Jeanne (2006); Lewis and Liu (2015); Coeurdacier, Rey, and Winant (2013); Didier, Rigobon, and Schmukler (2013); as well as Colacito and Croce (2011) and Stathopoulos (2017) on the high international correlation of state prices). A necessary but not sufficient condition to interpret the SDF correlation is for example that the domestic and foreign agents consume the same baskets of goods and participate in complete financial markets. Even in this case, the interpretation is subject to additional assumptions. In a multi-good world, variation in the relative prices of the goods drives a wedge between the pricing kernels, even in the case of perfect risk sharing (Cole and Obstfeld (1991)). Likewise, when markets are segmented, as in Alvarez, Atkeson, and Kehoe (2002) and Alvarez, Atkeson, and Kehoe (2009), the correlation of SDFs does not imply risk-sharing of the non-participating agents. Using our framework, we can derive a specific bound on the covariance of the permanent SDF component across different countries.

Proposition 4. *If the permanent SDF component is unconditionally lognormal, the cross-country covariance of the SDF' permanent components is bounded below by:*

$$cov\left(\log \frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}, \log \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) \quad \geq \quad E\left(\log \frac{R_{t+1}^*}{R_{t+1}^{(\infty),*}}\right) + E\left(\log \frac{R_{t+1}}{R_{t+1}^{(\infty)}}\right) - \frac{1}{2} var\left(\log \frac{S_{t+1}^{\mathbb{P}}}{S_t^{\mathbb{P}}}\right), \tag{25}$$

for any positive returns R_{t+1} and R_{t+1}^*. A conditional version of the expression holds for conditionally lognormal permanent pricing kernel components.

Therefore, this result extends the insights of Brandt, Cochrane, and Santa-Clara (2006) to the permanent components of the SDFs. Chabi-Yo and Colacito (2015) extend this lower bound to non-Gaussian pricing kernels and different horizons.

Since exchange rate changes and their transitory components are observable (due to the observability of the bonds' holding period returns), one can compute the variance of the permanent component of exchange rates, $var\left(\log \frac{S_{t+1}^{\mathbb{P}}}{S_t^{\mathbb{P}}}\right)$, which is the last term in the expression above. In the data, the contribution of that term is on the order of 1% or less. Given the large size of the equity premium compared to the term premium (a 7.5% difference according to Alvarez and Jermann, 2005), and the relatively small variance of the permanent component of exchange rates, the lower bound in Proposition 4 implies a large correlation of permanent SDF components across countries.

In Figure A12, we plot the implied correlation of the permanent SDF components against the volatility of the permanent SDF component in the symmetric two-country case, for two different scenarios: the dotted line is for $Std\left(\log S_t^{\mathbb{P}}/S_{t+1}^{\mathbb{P}}\right) = 10\%$, and the plain line is for $Std\left(\log S_t^{\mathbb{P}}/S_{t+1}^{\mathbb{P}}\right) = 16\%$. In both cases, the implied correlation of the permanent components of the domestic and foreign SDFs is clearly above 0.90.

While Brandt, Cochrane, and Santa-Clara (2006) show that the SDFs are highly correlated across countries, we find that the permanent components of the SDFs, which are the main sources of volatility for the SDFs, are highly correlated across countries.

G.2 A New Long-Term Bond Return Parity Condition

We end this paper with a potential new benchmark for exchange rates. While hundreds of papers have tested the U.I.P. condition, which assumes risk neutrality, we suggest a novel corner case, this time taking risk into account. When countries share permanent innovations to their SDFs, a simple long bond return parity condition emerges. The proposition below provides the result.

Proposition 5. *If the domestic and foreign pricing kernels have common permanent innovations, so $\Lambda_{t+1}^{\mathbb{P}}/\Lambda_t^{\mathbb{P}} = \Lambda_{t+1}^{\mathbb{P},*}/\Lambda_t^{\mathbb{P},*}$ for all states, then the one-period returns on the longest maturity foreign bonds in domestic currency terms are identical to the returns of the corresponding domestic bonds:*

$$R_{t+1}^{(\infty),*} \frac{S_t}{S_{t+1}} = R_{t+1}^{(\infty)}, \text{ for all states.} \tag{26}$$

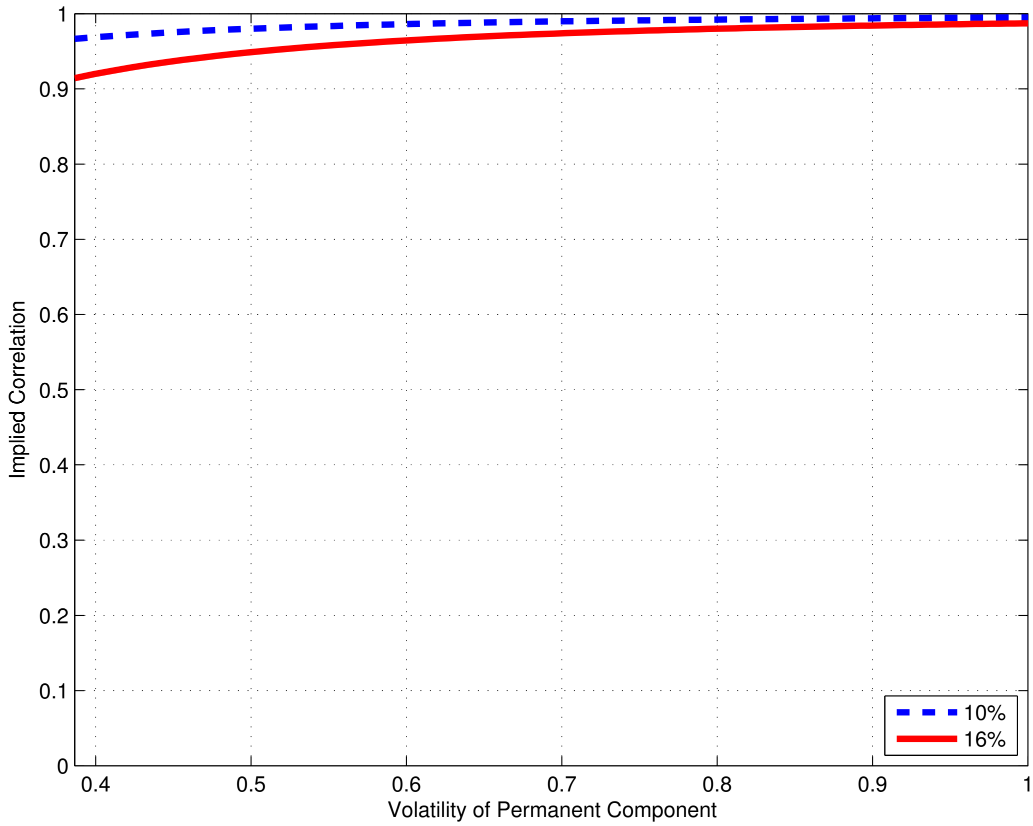

Figure A12: Cross-country Correlation of Permanent SDF Shocks — In this figure, we plot the implied correlation of the domestic and foreign permanent components of the SDF against the standard deviation of the permanent component of the SDF. The dotted line is for $Std\left(\log S_t^{\mathbb{P}}/S_{t+1}^{\mathbb{P}}\right) = 10\%$. The straight line is for $Std\left(\log S_t^{\mathbb{P}}/S_{t+1}^{\mathbb{P}}\right) = 16\%$. Following Alvarez and Jermann (2005), we assume that the equity minus bond risk premia are 7.5% in the domestic and foreign economies.

While Proposition 1 is about expected returns, Proposition 5 focuses on realized returns. In this polar case, even if most of the innovations to the pricing kernel are highly persistent, the shocks that drive exchange rates are not, because the persistent shocks are the same across countries. In that case, the exchange rate is a stationary process. In the absence of arbitrage opportunities, the currency exposure of a foreign long-term bond position to the stationary components of the pricing kernels is fully hedged by its interest rate risk exposure and does not affect the return differential with domestic bonds, which then measures the wedge between the non-stationary components of the domestic and foreign pricing kernels. When nominal exchange rates are stationary, this wedge is zero and long bond return parity obtains: bonds denominated in different currencies earn the same dollar returns, date by date.

H Finite vs. Infinite Maturity Bond Returns

Our empirical results pertain to 10- and 15-year bond returns while our theoretical results pertain to infinite-maturity bonds. This discrepancy raises the question of the theoretical validity of our empirical analysis. To address this question, we use the state-of-the-art Joslin, Singleton, and Zhu (2011) term structure model to study empirically the difference between the 10-year and infinite-maturity bonds. In particular, we estimate a version of the Joslin, Singleton, and Zhu (2011) term structure model with three factors, the three first principal components of the yield covariance matrix.[27] This Gaussian dynamic term structure model is estimated on zero-coupon rates over the period from April 1985 to December 2015, the same period used in our empirical work, for each country in our benchmark sample. Each country-specific model is estimated independently, without using any exchange rate data. The maturities considered are 6 months, and 1, 2, 3, 5, 7, and 10 years. Using the parameter estimates, we derive the implied bond returns for different maturities. We report simulated data for Australia, Canada, Germany, Japan, Norway, Switzerland, U.K., and U.S. and ignore the simulated data for New Zealand and Sweden as the parameter estimates imply that bond yields turn sharply negative on long maturities for those two countries. We study both unconditional and conditional returns, forming portfolios of countries sorted by the level or slope of their yield curves, as we did in the data. Table A24 reports the simulated moments.

We first consider the unconditional holding period bond returns across countries. The average (annualized) log return on the 10-year bond is lower than the log return on the infinite-maturity bond for all countries except Australia, the U.K., and the U.S., but the differences are not statistically significant, except for Japan. The unconditional correlation between the two log returns ranges from 0.88 to 0.96 across countries; for example, it is 0.89 for the U.S. Furthermore, the estimations imply very volatile log SDFs that exhibit little correlation across countries. As a result, the implied exchange rate changes are much more volatile than in the data. We then turn to conditional bond returns, obtained by sorting countries into two portfolios, either by the level of their short-term interest rate or by the slope of their yield curve. The portfolio sorts recover the results highlighted in the previous section: low (high) short-term interest rates correspond to high (low) average local bond returns. Likewise, low (high) slopes correspond to low (high) average local bond returns. The infinite maturity bonds tend to offer larger conditional returns than the 10-year bonds, but the differences are not significant. The correlation between the conditional returns of the 10-year and infinite maturity bond portfolios ranges from 0.86 to 0.93 across portfolios.

A clear limit of this experiment is that term structure models are not built to match infinite-maturity bonds, as these are unobservable. We thus learn from the term structure models by continuity. In theory, it is certainly possible to write a model where the 10-year bond returns, once expressed in the same currency, offer similar average returns across countries (as we find in the data), while the infinite maturity bonds do not. In that case, there would be a gap between our theory and the data. In such a model, however, exchange rates would have unit root components driven by common shocks and the cross-sectional distribution of exchange rates would fan out over time. For developing countries with strong trade links and similar inflation rates, this seems hard to defend. Moreover, although we cannot rule out its existence, we do not know of such a model. In the state-of-the-art of the term structure modeling, our inference about infinite-maturity bonds from 10-year bonds is reasonable.

[27] We thank the authors for making their code available on their web pages.

Table A24: Simulated Bond Returns

	Panel A: Country Returns							
	US	Australia	Canada	Germany	Japan	Norway	Switzerland	UK
$y^{(10)}$ (data)	5.58	6.97	5.81	4.97	2.77	4.26	3.17	6.10
$y^{(10)}$	5.58	6.97	5.81	4.97	2.77	4.26	3.18	6.09
$rx^{(10)}$	5.60	4.50	4.53	4.33	4.05	3.14	2.95	3.50
s.e.	[1.43]	[1.71]	[1.45]	[1.17]	[1.13]	[1.71]	[1.12]	[1.52]
$rx^{(\infty)}$	−0.44	2.17	6.69	6.33	7.38	5.96	6.42	2.74
s.e.	[10.87]	[10.38]	[8.47]	[2.33]	[2.46]	[3.89]	[3.23]	[4.52]
Corr ($rx^{(10)}$, $rx^{(\infty)}$)	0.89	0.92	0.89	0.92	0.93	0.96	0.93	0.88
$rx^{(\infty)} - rx^{(10)}$	−6.04	−2.33	2.16	2.00	3.33	2.83	3.47	−0.76
s.e.	[9.63]	[8.77]	[6.98]	[1.34]	[1.46]	[2.30]	[2.22]	[3.33]
$\sigma_{m^\star}$	239.17	241.92	127.14	118.45	211.76	132.76	227.59	153.22
$corr(m, m^\star)$	1.00	0.01	0.33	0.20	0.03	0.05	0.14	0.03
$\sigma_{\Delta s}$		310.81	202.65	244.63	314.14	190.44	271.17	279.99
	Panel B: Portfolio Returns							
		Sorted by Level			Sorted by Slope			
Sorting variable (level/slope)		2.57	5.60		0.15	1.81		
$rx^{(10)}$		4.10	4.52		3.00	5.48		
s.e.		[1.05]	[1.24]		[1.13]	[1.23]		
$rx^{(\infty)}$		4.48	6.00		0.79	9.61		
s.e.		[4.17]	[5.62]		[4.33]	[5.90]		
Corr ($rx^{(10)}$, $rx^{(\infty)}$)		0.86	0.93		0.89	0.90		
$rx^{(\infty)} - rx^{(10)}$		0.38	1.48		−2.21	4.13		
s.e.		[3.28]	[4.58]		[3.40]	[4.84]		

Notes: Panel A reports moments on simulated data at the country level. For each country, the table first compares the 10-year yield in the data and in the model, and then reports the annualized average simulated log excess return (in percentage terms) of bonds with maturities of 10 years and infinity, as well as the correlation between the two bond returns. The table also reports the annualized volatility of the log SDF, the correlation between the foreign log SDF and the U.S. log SDF, and the annualized volatility of the implied exchange rate changes. Panel B reports conditional moments obtained by sorting countries by either the level of their short-term interest rates or the slope of their yield curves into two portfolios. The table reports the average value of the sorting variable, and then the average returns on the 10-year and infinite-maturity bonds, along with their correlation. The simulated data come from the benchmark 3-factor model (denoted RPC) in Joslin, Singleton, and Zhu (2011) that sets the first 3 principal components of bond yields as the pricing factors. The model is estimated on zero-coupon rates for Germany, Japan, Norway, Switzerland, U.K., and U.S. The sample estimation period is 4/1985–12/2015. The standard errors (denoted s.e. and reported between brackets) were generated by block-bootstrapping 10,000 samples of 369 monthly observations.