# Online Appendix

# "The Term Structure of Currency Carry Trade Risk Premia"


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

- Section I presents robustness checks on the main time-series results reported in the paper:
  - subsection A reports the details of the time-series predictability results discussed in the main text;
  - subsection B reports time-series predictability results using standard asymptotic inference;
  - subsection C reports time-series predictability results using inflation and sovereign credit as additional controls;
  - subsection D 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 E reports time-series predictability results with GBP as base currency;
  - subsection F 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 II presents additional robustness checks for the cross-sectional portfolio results reported in the paper.
  - subsection A reports portfolio statistics for different time-windows (10/1983–12/2007, 1/1975–12/2007, 10/1983—12/2015);
  - subsection B 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 C 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 D 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 III reports additional results obtained with zero-coupon bonds for our benchmark sample of G10 countries and a larger sample of developed countries.

- Section IV 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 V 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 VI reports additional proofs of preference-free results.

- Section VII 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 VIII compares finite to infinite maturity bond returns in the benchmark Joslin, Singleton, and Zhu (2011) term structure model.

# I Robustness Checks on Time-Series Results

## A Time-Series Predictability

Table A1 reports additional information regarding the regressions presented in Table 1 of the main text. Similarly, Table A2 reports additional information regarding the properties of the long-short bond portfolio returns presented in Table 2 of the main text.

## B Time-Series Predictability using Standard Asymptotic Inference

Table A3 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.

## C Time-Series Predictability with Additional Controls

Table A4 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.

## D Time Series Regressions: Exchange Rate Changes and Local Bond Returns (Instead of Excess Returns)

Table A5 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

## Table A1: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | p-val | $\beta$ | s.e. | p-val | $R^2(\%)$ | $\alpha$ | s.e. | p-val | $\beta$ | s.e. | p-val | $R^2(\%)$ | $\alpha$ | s.e. | p-val | $\beta$ | 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 | | | | 0.15 | | | 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 (2011, 2012).

Table A2: 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.

## Table A3: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | 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 A4: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | 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.

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 A5 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.

## E Time Series Predictability with GBP as Base Currency

Table A6 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 A7 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

Table A5: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | 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 A6: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | 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.

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 A7: 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.

## F Individual Country Time-Series Predictability Results

Tables A8 and A9 report the time-series regression results when we end the sample in 2007: Table A8 considers the shorter 10/1983-12/2007 sample period, whereas Table A9 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 $R^2$s 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 A10 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 A11, A12 and A13 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 A11), 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 A12), 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

Table A8: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | 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 A9: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | 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 ($rx^{(10),\$}_{t+1} - rx^{(10)}_{t+1}$, left panel) or the currency excess return ($rx^{FX}_{t+1}$, middle panel) or the bond local currency return difference ($rx^{(10),*}_{t+1} - rx^{(10)}_{t+1}$, right panel) on the difference between the foreign nominal interest rate and the U.S. nominal interest rate ($r^{f,*}_t - r^f_t$, Panel A) or difference between the foreign nominal yield curve slope and the U.S. nominal yield curve slope ($[y^{(10,*)}_t - y^{(1,*)}_t] - [y^{(10)}_t - y^{(1)}_t]$, 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 A10: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2$(%) | $\alpha$ | s.e. | $\beta$ | 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.

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 A13), 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 A14 and A15 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 A16 and A17 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 A11: 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 A12: 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 A13: 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 A14: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | 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 A15: 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. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | s.e. | $R^2(\%)$ | $\alpha$ | s.e. | $\beta$ | 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 A16: 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 A17: 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.

# II 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.

## A Portfolio Cross-Sectional Evidence: Different Sample Periods

We start by considering different sample periods. Table A18 reports the results for the pre-crisis 10/1983-12/2007 sample, Table A19 for the pre-crisis 1/1975-12/2007 sample and Table A20 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.

## B 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.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 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

Table A18: 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 ($\Delta s$), 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 A19: 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 ($\Delta s$), 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 A20: 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 ($\Delta s$), 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.

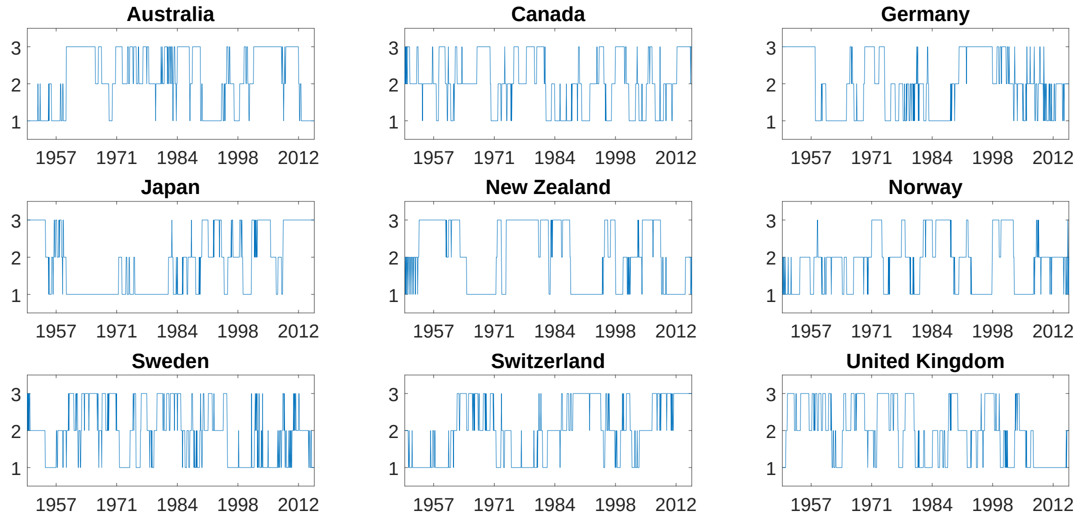


Figure A1: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 9 currencies sorted by the deviation of their short-term interest rates from the corresponding 10-year rolling mean. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to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

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 of the main text, 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 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

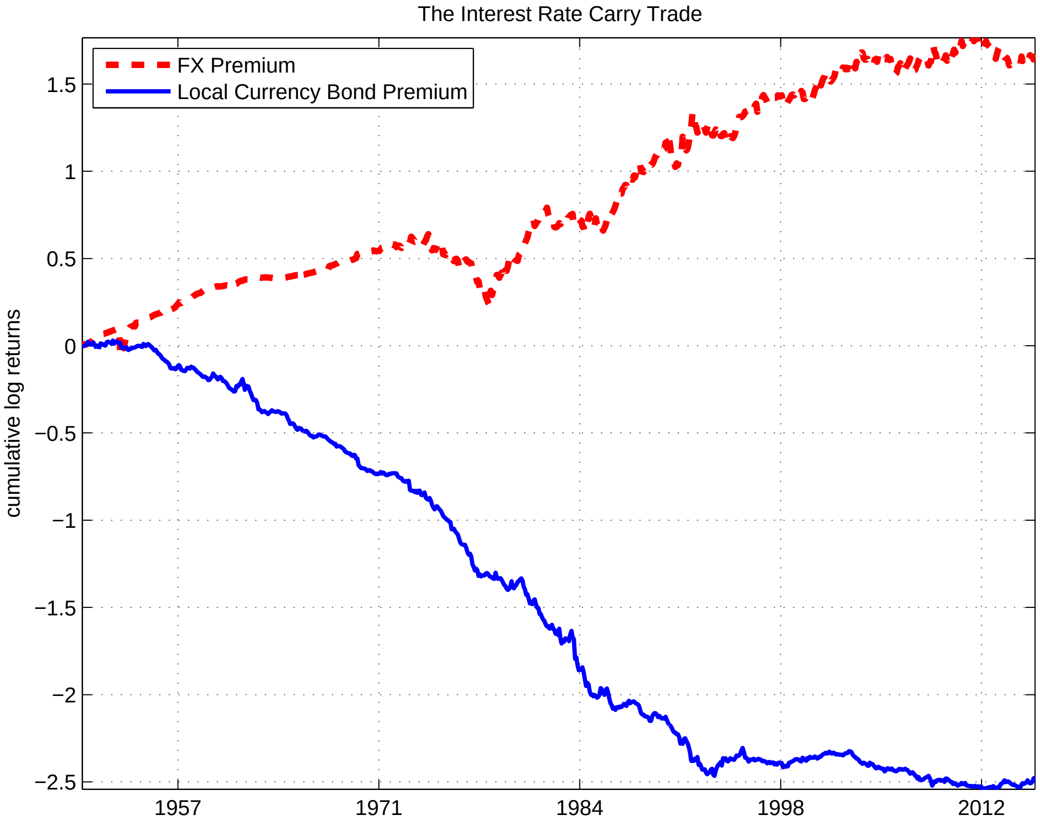


Figure A2: The Carry Trade and Term Premia – The figure presents the cumulative one-month log excess returns on investments in foreign Treasury bills and foreign 10-year bonds. The benchmark panel of countries includes Australia, Canada, Japan, Germany, Norway, New Zealand, Sweden, Switzerland, and the 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.

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 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.

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

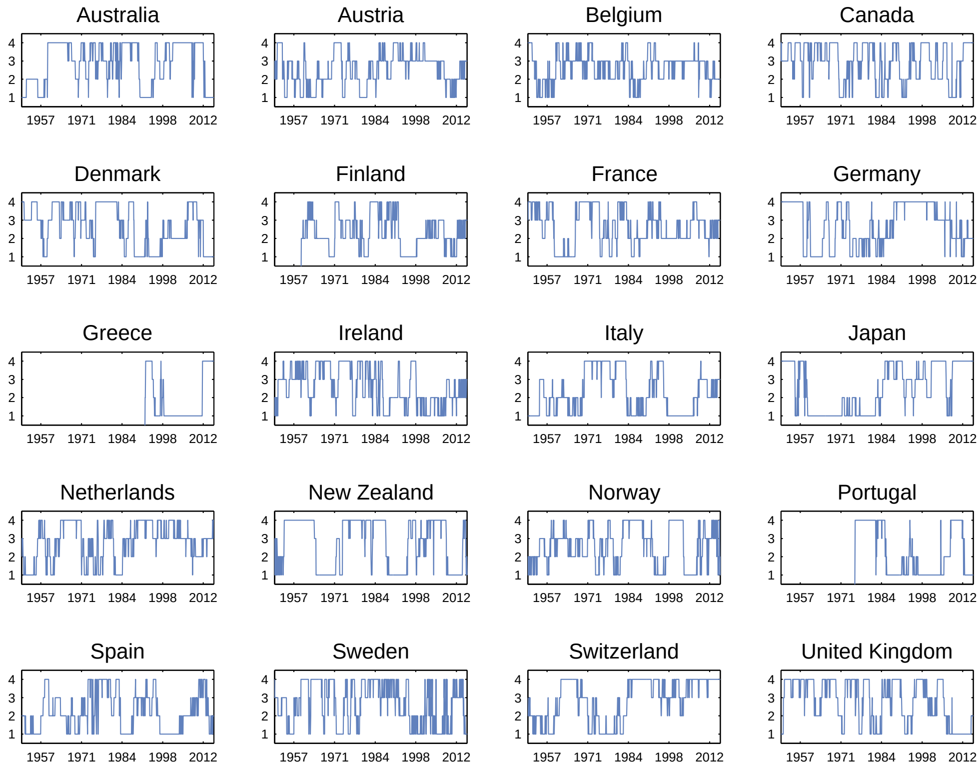


Figure A3: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

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.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).

## C 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.

### C.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.

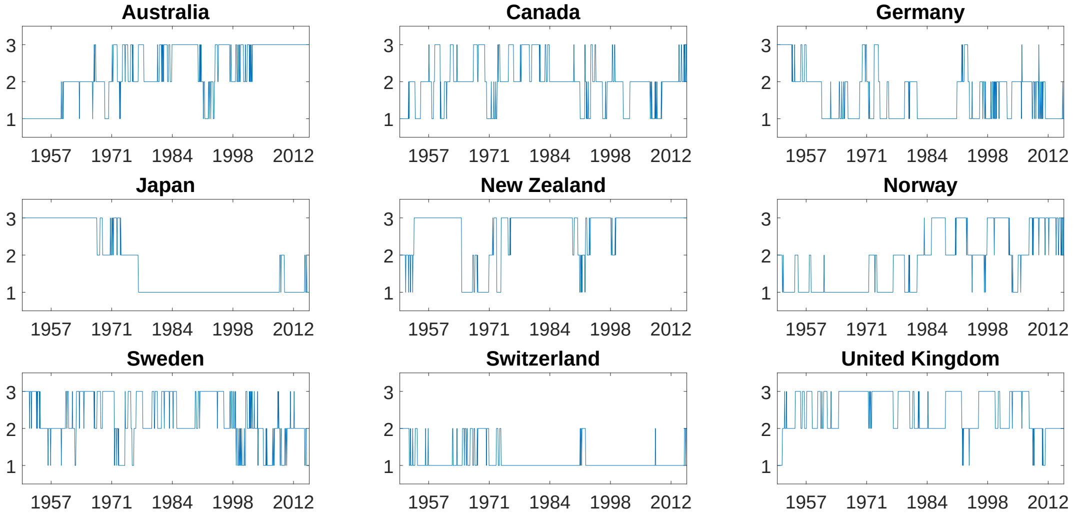


Figure A4: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 9 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

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 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).

### C.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

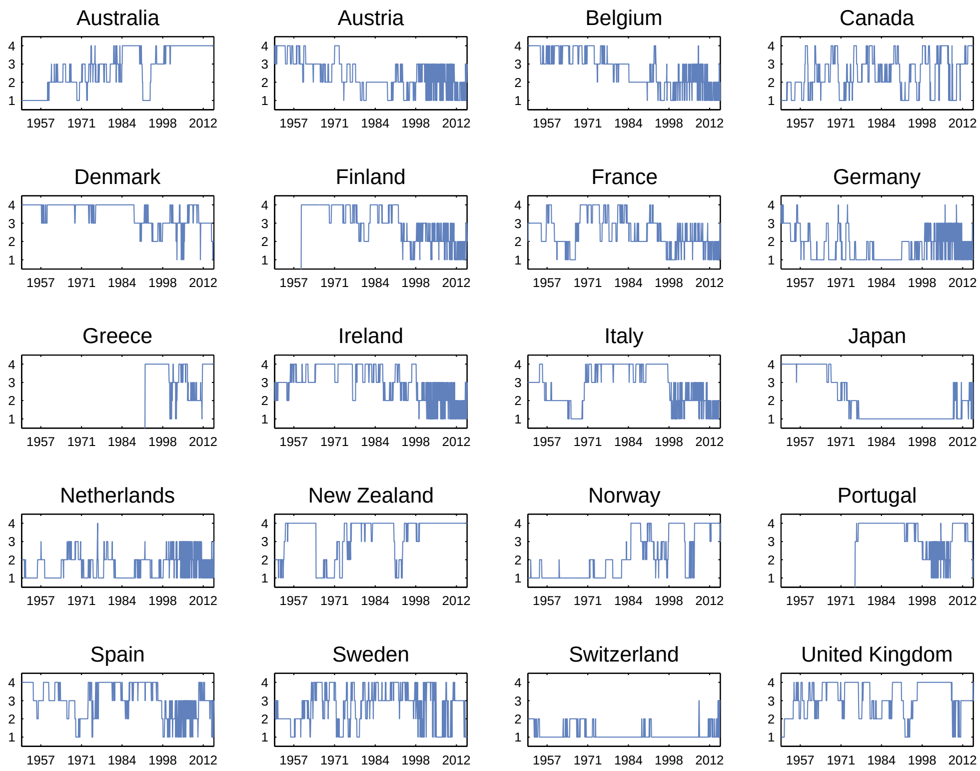


Figure A5: Composition of Interest Rate-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their short-term interest rates. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

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, 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.

### C.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).

# D 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.

## D.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.

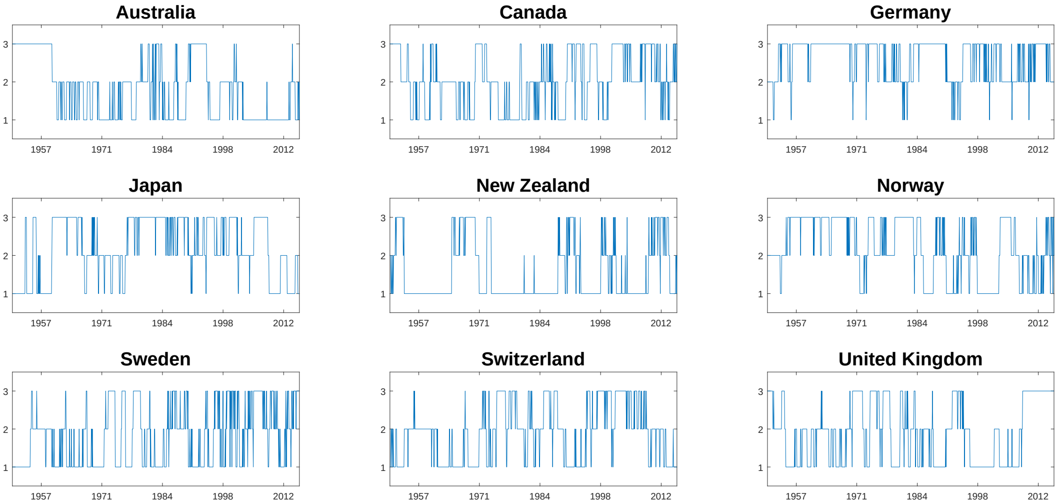


Figure A6: Composition of Slope-Sorted Portfolios — The figure presents the composition of portfolios of the currencies in the benchmark sample sorted by the slope of their yield curves. The portfolios are rebalanced monthly. The slope of the yield curve is measured by the spread between the 10-year 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 of the main text, 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).

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.

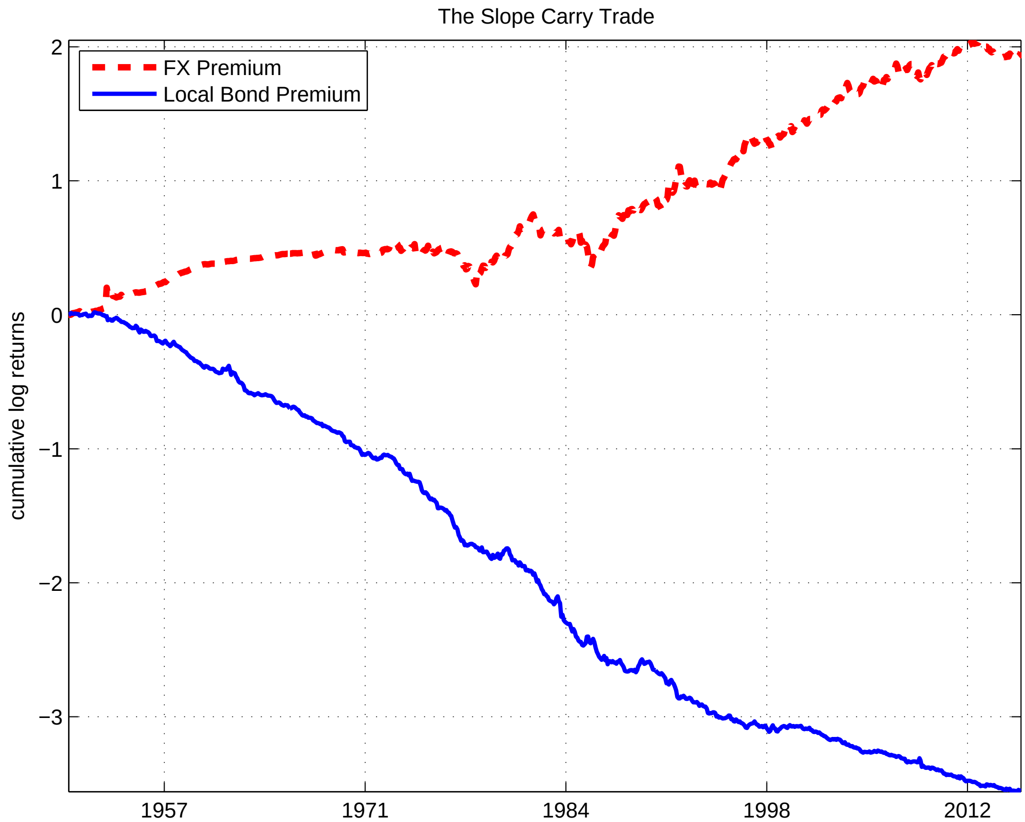


Figure A7: The Carry Trade and Term Premia: Conditional on the Slope of the Yield Curve – The figure presents the cumulative one-month log returns on investments in foreign Treasury bills and foreign 10-year bonds. The benchmark panel of countries includes Australia, Canada, Japan, Germany, Norway, New 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.

### D.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.

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

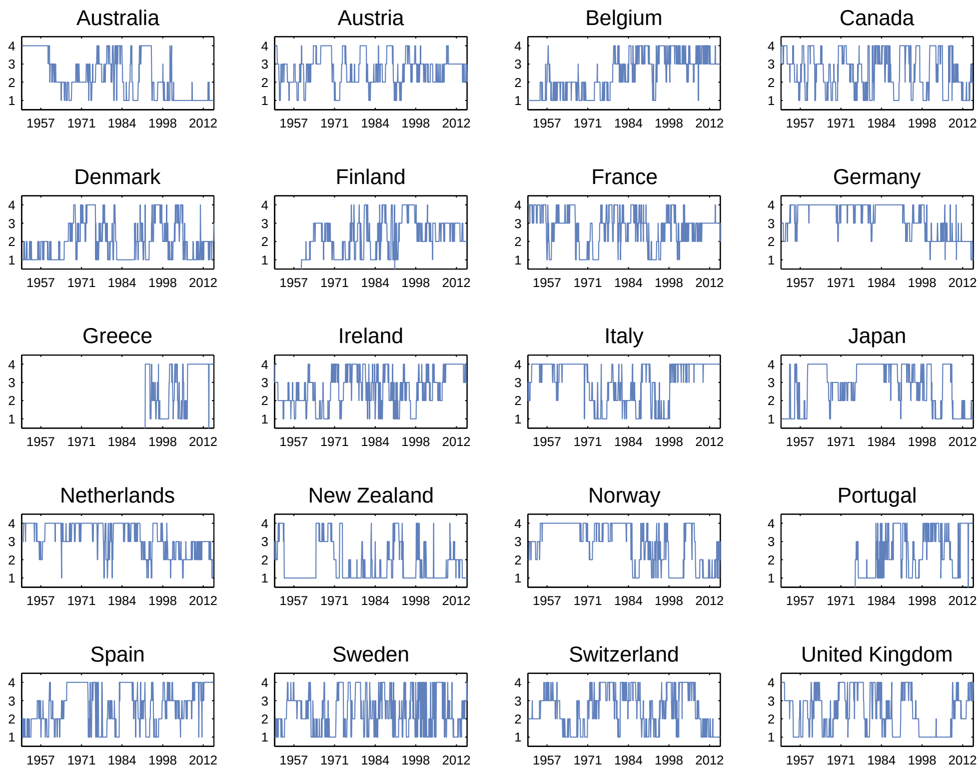


Figure A8: Composition of Slope-Sorted Portfolios — The figure presents the composition of portfolios of 20 currencies sorted by their yield curve slopes. The portfolios are rebalanced monthly. Data are monthly, from 1/1951 to 12/2015.

(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%).

### D.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%).

# III 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.

## A 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 Figure 1 in the main text).

## B 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.

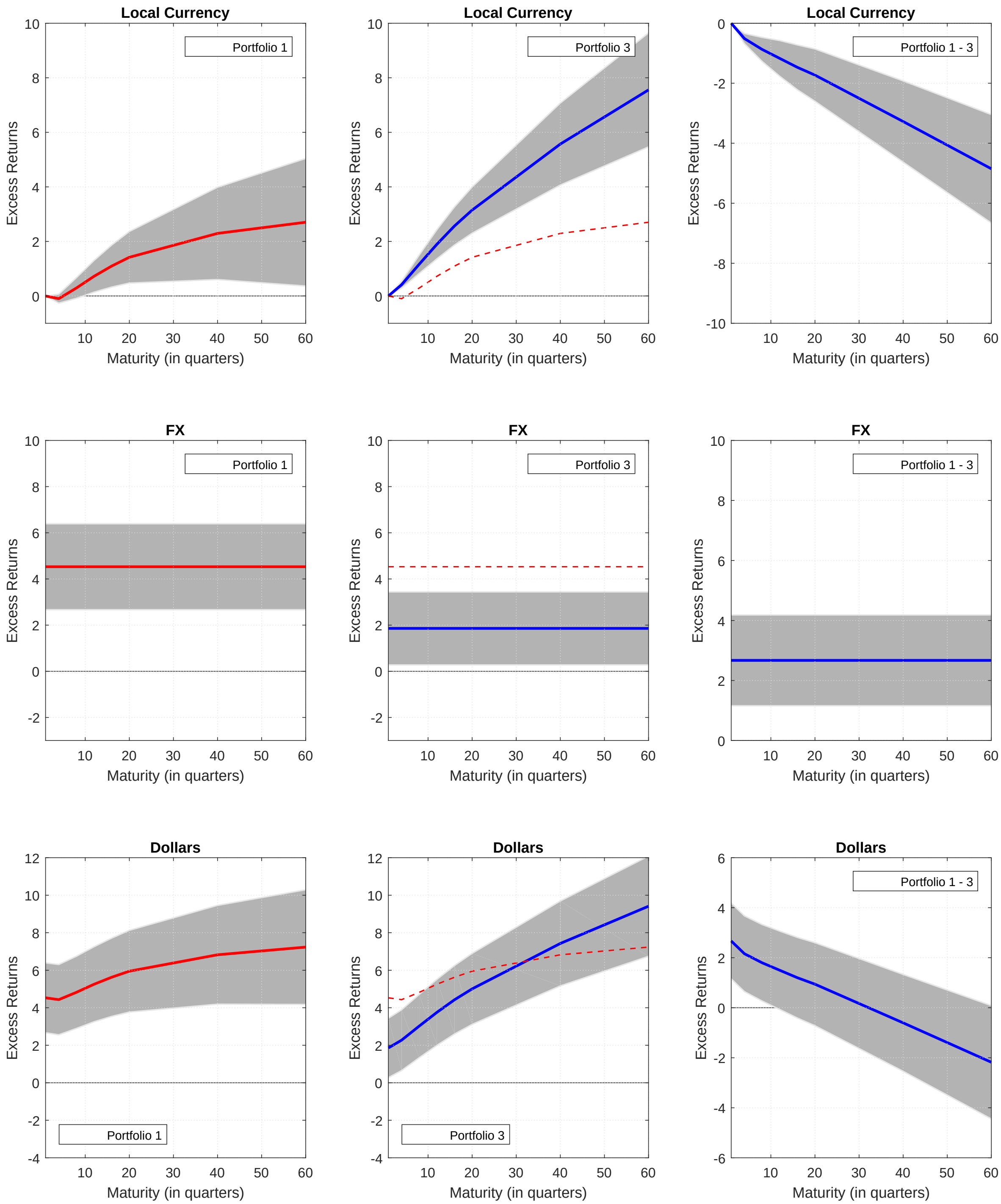


Figure A9: Dollar Bond Risk Premia Across Maturities— The figure shows the log excess returns on foreign bonds in local currency in the top panel, the currency excess return in the middle panel, and the log excess returns on foreign bonds in U.S. dollars in the bottom panel as a function of the bond 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.

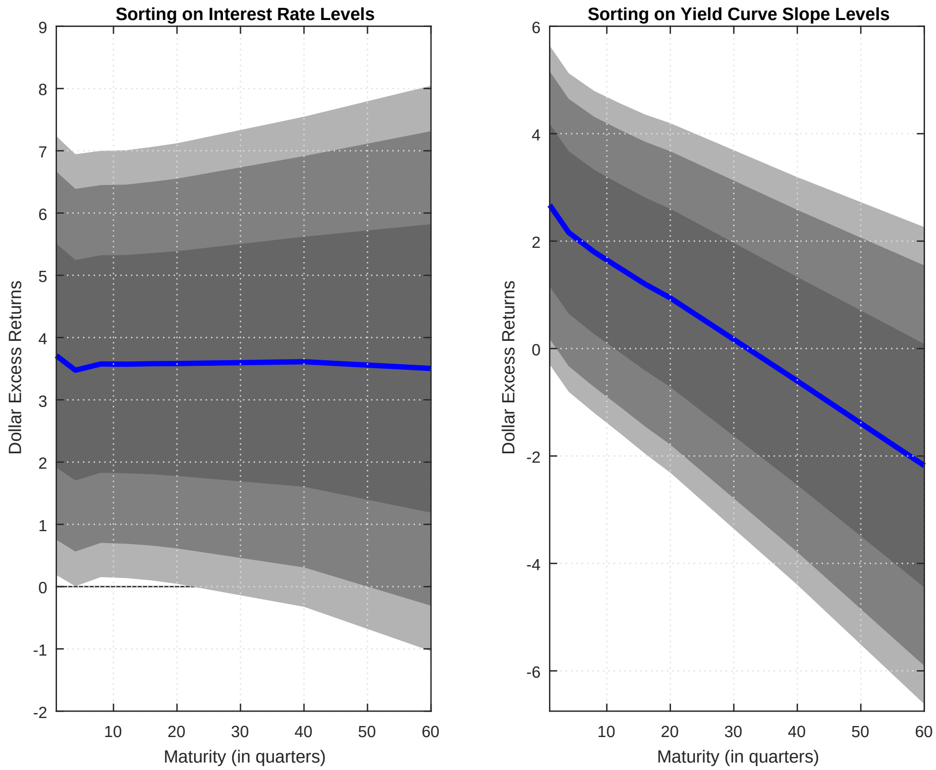


Figure A10: Long-Minus-Short Foreign Bond Risk Premia in U.S. Dollars— The figure shows the dollar log excess returns as a function of the bond maturities. Dollar excess returns correspond to the holding period returns expressed in U.S. dollars of investment strategies that go long and short foreign 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.

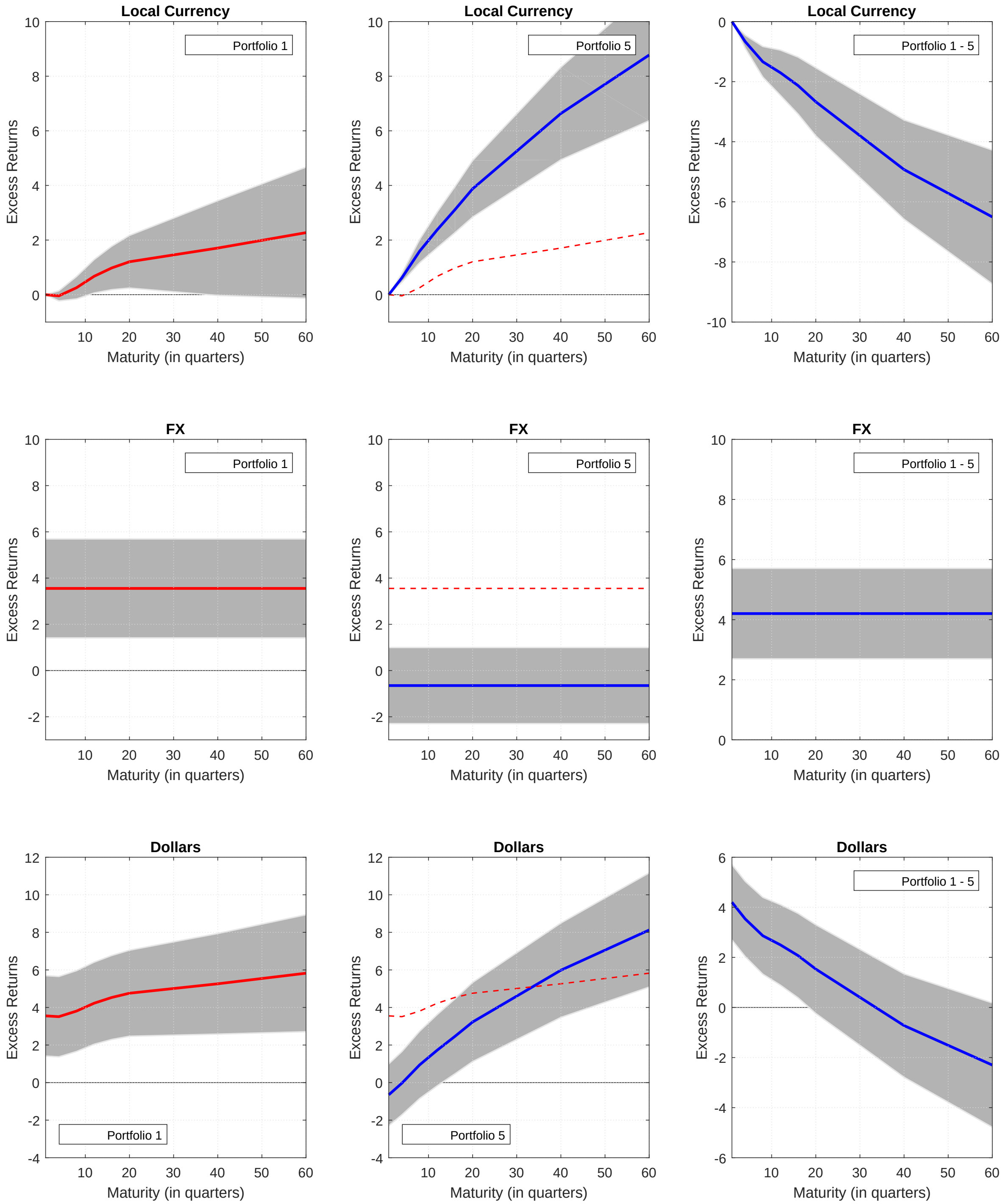


Figure A11: Dollar Bond Risk Premia Across Maturities: Extended Sample — The figure shows the local currency log excess returns in the top panel, and the dollar log excess returns in the bottom panel as a function of the bond maturities. The left panel focuses on Portfolio 1 (flat yield curve currencies) 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.

# IV 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 A21.

## A 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 $\varepsilon_{t+1}$ 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

Table A21: Condition 1: Dynamic Term Structure Model Scorecard

| | | |
|---|---|---|
| | *Vasicek (1977)* | *Cox, Ingersoll, and Ross (1985)* |
| *Model* | $-m_{t+1} = y_{1,t} + \frac{1}{2}\lambda^2\sigma^2 + \lambda\varepsilon_{t+1}$ | $-m_{t+1} = \alpha + \chi z_t + \sqrt{\gamma z_t} u_{t+1}$, |
| | $x_{t+1} = \rho x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}\left(0, \sigma^2\right)$ | $z_{t+1} = (1-\phi)\theta + \phi z_t - \sigma\sqrt{z_t} u_{t+1}$, |
| | $y_{1,t} = \delta + x_t$ | |
| *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* | $B_1^{\infty} =$ -$\lambda = \frac{1}{1-\rho}$ | $B_1^{\infty} = \frac{\chi}{1-\phi} = \sqrt{\gamma}/\sigma$ |
| | *Multi-factor Vasicek* | *Gaussian DTSM* |
| *Model* | $-m_{t+1} = y_{1,t} + \frac{1}{2}\Lambda_t'\Sigma\Lambda_t + \Lambda_t'\varepsilon_{t+1}$ | $-m_{t+1} = y_{1,t} + \frac{1}{2}\Lambda' V(x_t)\Lambda + \Lambda' V(x_t)^{1/2}\varepsilon_{t+1}$ |
| | $x_{t+1} = \Gamma x_t + \varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}(0, \Sigma)$ | $x_{t+1} = \Gamma x_t + V(x_t)^{1/2}\varepsilon_{t+1}, \quad \varepsilon_{t+1} \sim \mathcal{N}(0, I)$ |
| | $\Lambda_t = \Lambda_0 + \Lambda_1 x_t$ | $V_{ii}(x_t) = \alpha_i + \beta_i' x_t$ |
| | $y_{1,t} = \delta_0 + \delta_1' x_t$ | $y_{1,t} = \delta_0 + \delta_1' x_t$ |
| *Condition 1* | $(\Lambda_t' + B_1^{\infty\prime})\Sigma(\Lambda_t + B_1^{\infty}) = (\Lambda_t^{*\prime} + B_1^{*\infty\prime})\Sigma(\Lambda_t^* + B_1^{\infty *})$ | $(\Lambda' + B_1^{\infty\prime})V(0)(\Lambda + B_1^{\infty}) = (\Lambda^{*\prime} + B_1^{*\infty\prime})V(0)(\Lambda^* + B_1^{\infty *})$ |
| | | $(\Lambda' + B_1^{\infty\prime})V_x(\Lambda + B_1^{\infty}) = (\Lambda^{*\prime} + B_1^{*\infty\prime})V_x(\Lambda^* + B_1^{\infty *})$ |
| *Restriction* | $B_1^{\infty} = -\Lambda_0; \Lambda_1 = 0$ | $B_1^{\infty} = -\Lambda$ |
| | $B_1^{\infty\prime} = (I-\Gamma)^{-1}\delta_1' = -\Lambda_0'$ | $B_1^{\infty} = \left(\delta_1' + \frac{1}{2}\Lambda' V_x \Lambda\right)(1-\Gamma)^{-1}$ |

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 IV 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.

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 $\beta$ 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 $B_1^\infty$ is $1/(1-\rho)$. As a result, $B_0^n$ grows at a linear rate in the limit. We choose the constant $\beta$ 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 $\lambda$ is negative. The SDF is homoskedastic. The 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 Alvarez and Jermann (2005), 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:

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

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.[1]

## B 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:

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

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

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:

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

$x_t$ is a $k \times 1$ vector, and so are $\varepsilon_{t+1}$, $\delta_1$, $\Lambda_t$, and $\Lambda_0$, while $\Gamma$, $\Lambda_1$, and $\Sigma$ are $k \times k$ matrices.[2]

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.[3] The parameters of the temporary shocks satisfy $B^{\infty\prime}_{1k-h} = (I_{k-h} - \Gamma_{k-h})^{-1}\delta'_{1k-h} = -\Lambda'_{0k-h}$. This ensures that these shocks do not affect the permanent component of the pricing kernel.

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 $\Lambda_h$ and $\delta_{1h}$, which govern exposure to the permanent, global shocks. This condition eliminates any differences in permanent risk exposure across countries.[4] The nominal exchange

[2]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 $\Gamma = \rho$.

[3]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 $\Gamma$ is block-diagonal and $(I - \Gamma)$ is invertible, we can decompose the shocks as described

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

rate has no permanent component $\left(\frac{S_t^{\mathbb{P}}}{S_{t+1}^{\mathbb{P}}} = 1\right)$. From Backus, Foresi, and Telmer (2001), the expected log currency excess return is equal to:

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

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

## C Cox, Ingersoll, and Ross (1985) Model

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

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

where $M$ denotes the stochastic discount factor. In this model, log bond prices are affine in the state variable $z$: $p_t^{(n)} = -B_0^n - B_1^n z_t$. 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: $P_t^{(n)} = E_t(M_{t+1}P_{t+1}^{(n-1)})$. Thus the bond price coefficients evolve according to the following second-order difference equations:

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

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

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

where the constant $\beta$ 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: $\lim_{n\to\infty} B_0^n - B_0^{n-1} = \alpha + B_1^\infty(1-\phi)\theta$, where $B_1^\infty$ is defined implicitly in a second-order equation above. As a result, $B_0^n$ grows at a linear rate in the limit. We choose the constant $\beta$ to offset the growth in $B_0^n$ as $n$ becomes very large. Setting $\beta = e^{-\alpha - B_1^\infty(1-\phi)\theta}$ guarantees that Assumption 1 in Alvarez and Jermann (2005) is satisfied. The temporary pricing component of the pricing kernel is thus equal to:

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = \beta e^{B_1^\infty(z_{t+1}-z_t)} = \beta e^{B_1^\infty[(\phi-1)(z_t-\theta)-\sigma\sqrt{z_t}u_{t+1}]}.$$

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

$$\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}} = \frac{\Lambda_{t+1}}{\Lambda_t}\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right)^{-1} = \beta^{-1} e^{-\alpha-\chi z_t-\sqrt{\gamma z_t}u_{t+1}} e^{-B_1^\infty[(\phi-1)(z_t-\theta)-\sigma\sqrt{z_t}u_{t+1}]}. \tag{3}$$

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 $B_1^\infty$ 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 (1) 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 $B_1^\infty$ in Equation (3):

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

In this special case, $B_1^\infty = \sqrt{\gamma}/\sigma$. Using this result in Equation (3), the permanent component of the pricing kernel reduces to:

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

which is a constant.[5]

The expected bond excess return is:

$$E_t(r_{t+1}^{(n)}) - r_t^f + \frac{1}{2}var_t(r_{t+1}^{(n)}) = -cov_t(p_{t+1}^{(n-1)}, m_{t+1}) = B_1^{n-1}\sigma\sqrt{\gamma}z_t. \tag{4}$$

Recall that the risk-free rate is:

$$r_t^f = -E_t(\log M_{t+1}) - \frac{1}{2}Var_t(\log M_{t+1}) = \alpha + (\chi - \frac{1}{2}\gamma)z_t. \tag{5}$$

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 (3) 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

[5] 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 $B_1^\infty$ depends on $\chi$ and $\gamma$, as well as on the global parameters $\phi$ and $\sigma$. The two countries have perfectly correlated pricing kernels.

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 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 (4)) 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: $E_t[rx_{t+1}^{(\infty),*}] + E_t[rx_{t+1}^{FX}] = \frac{1}{2}\gamma z_t$. 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 $B_1^{\infty}$ depends on $\chi$ and $\gamma$, as well as on the global parameters $\phi$ and $\sigma$. 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 1.** *In the two-country CIR model, the transitory component of the exchange rate is given by:*

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

*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* $s_t = s_t^{\mathbb{T}}$*:*

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

*The expected rate of depreciation is*

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

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 (2018), 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 $\Delta s_{t+i}$ 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 (2018), 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 2.** *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* $\Delta s_{t+i}$ *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 $(\phi_1, \phi_2)$, and provided that $(\chi_1 - \frac{1}{2}\gamma_1)$ and $(\chi_2 - \frac{1}{2}\gamma_2)$ have opposite signs.

## D 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 $\varepsilon_{t+1}$, $\Lambda$, $\delta_1$, and $\beta_i$. But $\Gamma$ and $V$ are $k \times k$ matrices. A restricted version of the model would impose that $\beta_i$ 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, $P_t^{(n)} = \exp(-B_0^n - B_1^{n\prime} x_t)$. 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 $\beta_i$ 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 $\beta$ 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 $\Lambda$). 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 $\Lambda_{k-h} = -B^{\infty}_{1,k-h}$ (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:

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

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 $\Lambda_h$ in the $k$-factor model are equivalent to differences in $\gamma$ 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 ($B^{\infty\prime}_{1h}$), which can be traced back to differences in the sensitivity of the risk-free rate to the "permanent" state variable (i.e., different $\delta_1$ 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: $\Lambda' = -B^{\infty\prime}_1$ and $B^{\infty}_1(\Gamma - 1) + \delta'_1 + \frac{1}{2}\Lambda' V_x \Lambda = 0$. 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 $B^{\infty\prime}_1$ implied by the law of motion of $B_1$. As a result, the martingale component is constant as soon as $\Lambda = -B^{\infty}_1$.

Using Alvarez and Jermann (2005), the term premium on an infinite-maturity zero coupon bond is:

$$
E_t[rx^{(\infty)}_{t+1}] = -\delta_0 + \left((1-\Gamma)B^{\infty\prime}_1 - \delta'_1\right)x_t. \tag{6}
$$

In the absence of permanent shocks, when $\Lambda = -B^{\infty}_1$, this log bond risk premium equals half of the stochastic discount factor variance $E_t[rx^{(\infty)}_{t+1}] = \frac{1}{2}\Lambda' V(x_t)\Lambda$; it attains the upper bound on log risk premia. Consistent with the result in Equation (7) in the main text, the expected log currency excess return is equal to:

$$E_t\left[rx_{t+1}^{FX}\right] = \frac{1}{2}\Lambda' V(x_t)\Lambda - \frac{1}{2}\Lambda^{*\prime} V(x_t)\Lambda^*. \tag{7}$$

Differences in the market prices of risk $\Lambda$ imply non-zero currency risk premia. Adding the previous two expressions in Equations (6) and (7), 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: $\Lambda_{k-h}$ and $\Lambda_{k-h}^*$. If there are only permanent shocks ($h = k$), then the currency risk premium is zero.[6]

## E 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

[6] 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 $\Lambda_h$ in the $k$-factor model are equivalent to differences in $\gamma$ 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 $\delta_1$ parameters). These correspond to differences in $\chi$ in the CIR model.

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.[7] 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 country-specific and world state variables follow autoregressive square-root processes:

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

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:

$$-\widetilde{m}_{t+1}^i = \alpha + \chi z_t^i + \sqrt{\gamma z_t^i}u_{t+1}^i + \tau z_t^w + \sqrt{\delta^i z_t^w}u_{t+1}^w + \sqrt{\kappa z_t^i}u_{t+1}^g,$$

where $u_{t+1}^i$ is a country-specific SDF shock, while $u_{t+1}^w$ and $u_{t+1}^g$ 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 $\delta^i$ on the world shock $u_{t+1}^w$; 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, $u_{t+1}^w$, is priced proportionally to country exposure $\delta^i$, and since $\delta^i$ is a fixed characteristic of country $i$, differences in such exposure are *permanent*. The second type, $u_{t+1}^g$, is priced proportionally to $z_t^i$, so heterogeneity with respect to this innovation is *transitory*: all countries

[7] 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).

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 $\widetilde{r}_t^{f,i} = \alpha + \left(\chi - \frac{1}{2}(\gamma+\kappa)\right) z_t^i + \left(\tau - \frac{1}{2}\delta^i\right) z_t^w$.

Country $i$'s inflation process is given by $\pi_{t+1}^i = \pi_0 + \eta^w z_t^w + \sigma_\pi \epsilon_{t+1}^i$, where the inflation innovations $\epsilon_{t+1}^i$ are i.i.d. Gaussian. It follows that the log nominal risk-free rate in country $i$ is given by $r_t^{f,i} = \pi_0 + \alpha + \left(\chi - \frac{1}{2}(\gamma+\kappa)\right) z_t^i + \left(\tau + \eta^w - \frac{1}{2}\delta^i\right) z_t^w - \frac{1}{2}\sigma_\pi^2$. The nominal bond prices in logs are affine in the state variable $z$ and $z^w$: $p_t^{(n),i} = -C_0^{n,\$,i} - C_1^{n,\$} z_t - C_2^{n,\$,i} z_t^w$, where the loadings $(C_0^{n,\$,i}, C_1^{n,\$}, C_0^{n,\$,i})$ are defined in the Appendix. Equation (7) in the main text implies that 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 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 $z^w$: $p_t^{i,(n)} = -C_0^{i,n} - C_1^n z_t - C_2^{i,n} z_t^w$. To calculate the parameter set $(C_0^{i,n}, C_1^{i,n}, C_2^{i,n})$, we follow the usual recursive process. In particular, the price of a one-period nominal bond is:

$$P^{i,(1)} = E_t(M_{t+1}^i) = E_t\left(e^{-\alpha - \chi z_t - \tau z_t^w - \sqrt{\gamma z_t^i} u_{t+1}^i - \sqrt{\delta^i z_t^w} u_{t+1}^w - \sqrt{\kappa z_t^i} u_{t+1}^g - \pi_0 - \eta^w z_t^w - \sigma_\pi \epsilon_{t+1}^i}\right).$$

As a result, $C_0^1 = \alpha + \pi_0 - \frac{1}{2}\sigma_\pi^2$, $C_1^1 = \chi - \frac{1}{2}(\gamma+\kappa)$, and $C_2^{i,1} = \tau - \frac{1}{2}\delta^i + \eta^w$.

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

$$\begin{aligned}
-C_0^{i,n} - C_1^n z_t - C_2^{i,n} z_t^w &= -\alpha - \chi z_t - \tau z_t^w - C_0^{n-1} - C_1^{n-1}\left[(1-\phi)\theta + \phi z_t\right] - C_2^{i,n-1}\left[(1-\phi^w)\theta^w + \phi^w z_t^w\right] \\
&+ \frac{1}{2}(\gamma+\kappa) z_t + \frac{1}{2}\left(C_1^{n-1}\right)^2 \sigma^2 z_t - \sigma\sqrt{\gamma} C_1^{n-1} z_t \\
&+ \frac{1}{2}\delta^i z_t^w + \frac{1}{2}\left(C_2^{i,n-1}\right)^2 (\sigma^w)^2 z_t^w - \sigma^w \sqrt{\delta^i} C_2^{i,n-1} z_t^w \\
&- \pi_0 - \eta^w z_t^w + \frac{1}{2}\sigma_\pi^2
\end{aligned}$$

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

$$\begin{aligned}
C_0^{i,n} &= \alpha + \pi_0 - \frac{1}{2}\sigma_\pi^2 + C_0^{n-1} + C_1^{n-1}(1-\phi)\theta + C_2^{i,n-1}(1-\phi^w)\theta^w, \\
C_1^n &= \chi - \frac{1}{2}(\gamma+\kappa) + C_1^{n-1}\phi - \frac{1}{2}\left(C_1^{n-1}\right)^2\sigma^2 + \sigma\sqrt{\gamma}C_1^{n-1} \\
C_2^{i,n} &= \tau - \frac{1}{2}\delta^i + \eta^w + C_2^{i,n-1}\phi^w - \frac{1}{2}\left(C_2^{i,n-1}\right)^2(\sigma^w)^2 + \sigma^w\sqrt{\delta^i}C_2^{i,n-1}.
\end{aligned}$$

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 $\beta$ 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 $B_1^\infty$ in Equation (3):

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

where we have imposed $\kappa = 0$. In this special case, $C_1^\infty = \sqrt{\gamma}/\sigma$. Using this result in Equation (3), the permanent component of the SDF reduces to:

$$\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_{t}} = \frac{\Lambda_{t+1}}{\Lambda_{t}}\left(\frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_{t}}\right)^{-1} = \beta^{-1}e^{-\tau z^w_t-\sqrt{\delta^i z^w_t}u^w_{t+1}}e^{C^{i,\infty}_2\left[(\phi^w-1)(z^w_t-\theta^w)-\sigma\sqrt{z^w_t}u^w_{t+1}\right]}.$$

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

$$E_t[rx^{(n)}_{t+1}] = [-\frac{1}{2}\left(C^{n-1}_1\right)^2\sigma^2+\sigma\sqrt{\gamma}C^{n-1}_1]z_t+[-\frac{1}{2}\left(C^{i,n-1}_2\right)^2\sigma^2+\sigma\sqrt{\delta}^i C^{i,n-1}_2]z^w_t.$$

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

$$E_t[rx^{(\infty)}_{t+1}] = [-\frac{1}{2}\left(C^{\infty}_1\right)^2\sigma^2+\sigma\sqrt{\gamma}C^{\infty}_1]z_t+[-\frac{1}{2}\left(C^{i,\infty}_2\right)^2\sigma^2+\sigma\sqrt{\delta}^i C^{i,\infty}_2]z^w_t.$$

The $-\frac{1}{2}\left(C^{\infty}_1\right)^2\sigma^2$ is a Jensen term. The term premium is driven by $\sigma\sqrt{\gamma}C^{\infty}_1 z_t$, where $C^{\infty}_1$ is defined implicitly in the second order equation $B^{\infty}_1=\chi-\frac{1}{2}(\gamma+\kappa)+C^{\infty}_1\phi-\frac{1}{2}\left(C^{\infty}_1\right)^2\sigma^2+\sigma\sqrt{\gamma}C^{\infty}_1$. Consider the special case of $C^{\infty}_1(1-\phi)=\chi$ and $\kappa=0$ and $C^{i,\infty}_2(1-\phi)=\tau$. In this case, the expected term premium is simply $E_t[rx^{(\infty)}_{t+1}]=\frac{1}{2}(\gamma z_t+\delta z^w_t)$, 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:

$$E_t[rx^{(\infty),i}_{t+1}] = \left[C^{\infty}_1(1-\phi)-\chi+\frac{1}{2}(\gamma+\kappa)\right]z^i_t+\left[C^{i,\infty}_2(1-\phi^w)-\tau+\frac{1}{2}\delta^i-\eta^w\right]z^w_t.$$

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 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 A22 and the simulation results in Table A23. 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 $\delta$ 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 $-3.66\%$ on average per annum. On the other hand, high interest rate currencies typically have high $\delta$. Long positions in these currencies earn positive excess returns ($rx^{FX}$) of 2.45% 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 6.12% and a Sharpe ratio of 0.51.

This spread is not offset by higher local currency bond risk premia in the low interest rate countries with higher $\delta$. The log excess return on the 30-year zero coupon bond is 0.62% in the first portfolio compared to 0.89 % in the last portfolio. At the 30-year maturity, the high-minus-low carry trade strategy still delivers a profitable excess return of 6.39% and a Sharpe ratio of 0.47. 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 5.93% and a Sharpe ratio of 0.43.

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 $\delta$ varies across countries. The heterogeneity in $\delta$'s across countries generates substantial dispersion in exposure to the permanent component. As a result, our long-run uncovered bond parity condition is violated.

Table A22: Parameter Estimates

| | | | | | |
|---|---|---|---|---|---|
| | | Stochastic discount factor | | | |
| $\alpha$ (%) | $\chi$ | $\tau$ | $\gamma$ | $\kappa$ | $\delta$ |
| 0.76 | 0.89 | 0.06 | 0.04 | 2.78 | 0.36 |
| | | State variable dynamics | | | |
| $\phi$ | $\theta$ (%) | $\sigma$ (%) | $\phi^w$ | $\theta^w$ (%) | $\sigma^w$ (%) |
| 0.91 | 0.77 | 0.68 | 0.99 | 2.09 | 0.28 |
| | Inflation dynamics | | | Heterogeneity | |
| $\eta^w$ | $\pi_0$ (%) | $\sigma^\pi$ (%) | | $\delta_h$ | $\delta_l$ |
| 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 $\delta^i$, which is distributed uniformly on $[\delta_h, \delta_l]$. The home country exhibits the average $\delta$, which is equal to 0.36.

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

Table A23: 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 | | | | | |
| $\Delta s$ | 1.43 | 0.33 | −0.19 | −0.26 | −0.43 | −0.98 |
| $\sigma_{\Delta s}$ | 11.18 | 9.57 | 9.11 | 8.90 | 8.94 | 9.40 |
| $r^{f,*} - r^f$ | −2.23 | −1.28 | −0.69 | −0.15 | 0.39 | 1.47 |
| $rx^{(30),*}$ | 0.62 | 0.71 | 0.78 | 0.82 | 0.86 | 0.89 |
| | Panel B: Carry Returns with Short-Term Bills | | | | | |
| $rx^{FX}$ | −3.66 | −1.62 | −0.50 | 0.11 | 0.82 | 2.45 |
| | Panel C: Carry Returns with Long-Term Bonds | | | | | |
| $rx^{(30),\$}$ | −3.04 | −0.91 | 0.27 | 0.93 | 1.68 | 3.35 |
| Section II: Sorting by Interest Rate Slopes | | | | | | |
| | Panel A: Exchange Rates, Interest Rate Slopes, and Bond Returns | | | | | |
| $\Delta s$ | −2.27 | −1.47 | −0.90 | −0.13 | 0.20 | 1.44 |
| $\sigma_{\Delta s}$ | 11.13 | 9.42 | 8.89 | 8.80 | 8.96 | 10.12 |
| $y^{10} - y^{1/4}$ | −0.89 | −0.44 | −0.15 | 0.10 | 0.36 | 1.02 |
| $rx^{(30),*}$ | 0.82 | 0.83 | 0.82 | 0.76 | 0.78 | 0.79 |
| | Panel B: Carry Returns with Short-Term Bills | | | | | |
| $rx^{FX}$ | 3.40 | 2.09 | 1.19 | 0.13 | −0.51 | −2.50 |
| | Panel C: Carry Returns with Long-Term Bonds | | | | | |
| $rx^{(30),\$}$ | 4.22 | 2.92 | 2.01 | 0.89 | 0.28 | −1.71 |

*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 ($\Delta s$), 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.

permanent component of the SDF is given by:

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

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 $\delta$ countries is exactly offset by the lower bond risk premium. As all these models show, Proposition 1 and Condition 1 in the main text 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 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.[8] In this special case, $\frac{\chi}{(1-\phi)} = C_1^{\infty} = \sqrt{\gamma}/\sigma > 0$. An increase in risk abroad causes the long

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

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 $\Delta s_{t+i}$ 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 (2018), 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 $\Delta s_{t+i}$ 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 $(\phi_1, \phi_2)$, and provided that $(\chi_1 - \frac{1}{2}\gamma_1)$ and $(\chi_2 - \frac{1}{2}\gamma_2)$ have opposite signs.

## F 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 $\delta^i$, we need to introduce another source of heterogeneity across countries. Countries must differ in $\tau$, $\phi^w$, $\sigma^w$, or $\eta^w$ (or a combination of those). Without this additional source of heterogeneity, there are at most two values of $\delta^i$ that are possible (for each set of parameters). [9] Here we simply choose to let the parameters $\tau$ differ across countries.

**Bond Prices** Our model only allows for heterogeneity in the exposure to the transitory common shocks ($\delta^i$), 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

[9] 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.

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: $\left[C_1^{\infty}(1-\phi) - \chi\right](z_t^i - z_t)$. 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 ($\delta$), 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 $\delta^i$ 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 $\delta^i$ 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.

# V 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 equilibrium models: models with external habit formation, models with long run risks, and models with rare disasters.

## A External Habit Model

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

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

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 $\lambda$ is

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

where $\bar{S} = \sigma\sqrt{\frac{\gamma}{1-\phi-B/\gamma}}$ is the steady-state value of the surplus consumption ratio and $s_{max} = \bar{s} + \frac{1}{2}(1-\bar{S}^2)$ 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

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

which can be also written as

$$r_t^f = -\log\delta + \gamma g - \frac{1}{2}\frac{\gamma^2\sigma^2}{\bar{S}^2} + B(\bar{s}-s_t).$$

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 $s$ (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 $s$ 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. 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 $\phi$ of the form

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

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

$$E_t\left[\exp\left(\log\delta - \gamma\left[g + (\phi-1)(s_t-\bar{s}) + (1+\lambda(s_t))\varepsilon_{t+1}\right] + cs_{t+1}\right)\right] = \exp(\beta + cs_t)$$

which, after some algebra, yields

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

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

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

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

$$\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}} = e^{\beta - c(s_{t+1} - s_t)} = e^{\log\delta - \gamma g + \frac{\gamma^2\sigma^2}{2} - \gamma((1-\phi)(\bar{s} - s_t) + \lambda(s_t)\varepsilon_{t+1})},$$

and the permanent SDF component is

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

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

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

the conditional entropy of the permanent pricing kernel component is

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

and the conditional entropy of the transitory pricing kernel component is

$$L_t\left(\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right) = \frac{1}{2}var_t\left(\log\frac{\Lambda_{t+1}^{\mathbb{T}}}{\Lambda_t^{\mathbb{T}}}\right) = \frac{\gamma^2\sigma^2}{2}\lambda(s_t)^2.$$

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

$$E_t\left[rx_{t+1}^{(\infty)}\right] = L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) = \frac{\gamma^2\sigma^2}{2}\frac{1}{\bar{S}^2}(1-2(s_t-\bar{s})) - \frac{\gamma^2\sigma^2}{2} = \frac{\gamma^2\sigma^2}{2}\left[\frac{1}{\bar{S}^2}(1-2(s_t-\bar{s}))-1\right].$$

Condition 1 implies that $\gamma^2\sigma^2 = \gamma^{*,2}\sigma^{*,2}$.

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 $s$, does not affect marginal utility and exchange rates in the long run. The long-run loading of the exchange rate on the surplus consumption ratio is given by: $\sum_{i=1}^{\infty} E_t[\Delta s_{t+i}] = \sum_{i=1}^{\infty} E_t[\log \frac{\Lambda_{t+i}}{\Lambda_t} - \log \frac{\Lambda_{t+i}^*}{\Lambda_t^*}] = -\sum_{i=1}^{\infty} \phi^{i-1}\gamma(1-\phi)(s_t^* - s_t) = -\gamma(s_t^* - s_t)$. Thus, long-run U.I.P holds,

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

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 3.** *In the symmetric external habit model, the slope coefficients in regressions of $\Delta s_{t+i}$ 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.

In an asymmetric model, in order for Condition 1 to hold, countries can only differ in their surplus consumption ratio parameters and in their consumption growth rate parameter ($g$), as differences in the other parameters $(\gamma, \sigma^2)$ 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 sources of heterogeneity.

## B Long-Run Risks Model

We now consider the long-run risk 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 $\psi$ represents the intertemporal elasticity of substitution in an environment without risk. Aggregate consumption growth $\Delta c_{t+1}$ 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, the values of which (except $A_0$) are reported in Panel A of Table A24. As usual, we assume that the agent has preferences for early resolution of uncertainty ($\gamma > \frac{1}{\psi}$), so $B_2 < 0$. In Panel B of the same table, we report the SDF of the homoskedastic version of the model, which is a special case of the full version. In the remainder of this section, we focus on the full, heteroskedastic version of the model.

Table A24: Pricing Kernel Loadings in the Long Run Risks Model

| Loadings | Parameters | Loadings | Parameters |
|---|---|---|---|
| Panel A: Heteroskedastic Model | | | |
| $\log\frac{\Lambda_{t+1}}{\Lambda_t}=A_0+A_1x_t+A_2u_t+A_3w_t+B_1\sqrt{u_t}\varepsilon^c_{t+1}+B_2\sqrt{w_t}\varepsilon^x_{t+1}+B_3\varepsilon^u_{t+1}+B_4\varepsilon^w_{t+1}.$ | | | |
| $A_1$ | $-\frac{1}{\psi}$ | $B_1$ | $-\gamma$ |
| $A_2$ | $\left(\frac{1}{\psi}-\gamma\right)\frac{\gamma-1}{2}$ | $B_2$ | $\left(\frac{1}{\psi}-\gamma\right)\frac{\kappa}{1-\kappa\phi^x}$ |
| $A_3$ | $\left(\frac{1}{\psi}-\gamma\right)\frac{\gamma-1}{2}\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2$ | $B_3$ | $\left(\frac{1}{\psi}-\gamma\right)\frac{1-\gamma}{2}\frac{\kappa}{1-\kappa\phi^u}\sigma^u$ |
| | | $B_4$ | $\left(\frac{1}{\psi}-\gamma\right)\left(\frac{\kappa}{1-\kappa\phi^x}\right)^2\frac{\kappa}{1-\kappa\phi^w}\sigma^w.$ |
| Panel B: Homoskedastic Model | | | |
| $\log\frac{\Lambda_{t+1}}{\Lambda_t}=A_0+A_1x_t+B_1\sqrt{\theta^u}\varepsilon^c_{t+1}+B_2\sqrt{\theta^w}\varepsilon^x_{t+1}.$ | | | |
| $A_1$ | $-\frac{1}{\psi}$ | $B_1$ | $-\gamma$ |
| | | $B_2$ | $\left(\frac{1}{\psi}-\gamma\right)\frac{\kappa}{1-\kappa\phi^x}$ |

*Notes:* Pricing kernel loading parameters in the long run risks model. Parameter $\kappa$ 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 $\kappa=\delta$.

The conditional SDF entropy and the equilibrium log risk-free rate are given by:

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

Thus, the risk-free rate is positively associated with $x_t$, the predictable component of consumption growth, due to the intertemporal smoothing effect, and negatively associated with $u_t$, 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 $w_t$, 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. The necessary condition (7) in the main text 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.

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

$\phi$ 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

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

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

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

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

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

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

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

$$\frac{\Lambda^{\mathbb{P}}_{t+1}}{\Lambda^{\mathbb{P}}_{t}} = \frac{\Lambda_{t+1}}{\Lambda_{t}} \left( \frac{\Lambda^{\mathbb{T}}_{t+1}}{\Lambda^{\mathbb{T}}_{t}} \right)^{-1} = e^{-(1/2)(B_3+c_2\sigma^u)^2-(1/2)(B_4+c_3\sigma_w)^2-(1/2)B_1^2 u_t-(1/2)(B_2+c_1)^2 w_t} \times$$
$$e^{B_1\sqrt{u_t}\varepsilon^c_{t+1}+(B_2+c_1)\sqrt{w_t}\varepsilon^x_{t+1}+(B_3+c_2\sigma^u)\varepsilon^u_{t+1}+(B_4+c_3\sigma^w)\varepsilon^w_{t+1}}.$$

In summary, both SDF components are exposed to the consumption drift innovation $\varepsilon^x_{t+1}$, the consumption growth variance innovation $\varepsilon^u_{t+1}$, and the consumption drift variance innovation $\varepsilon^w_{t+1}$, but only the permanent SDF component is exposed to the consumption growth innovation $\varepsilon^c_{t+1}$. 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 ($\varepsilon^x_{t+1} < 0$) 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

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

The permanent SDF component can also be obtained following the Alvarez-Jermann decomposition, where the transient component is defined on the basis of the infinite-maturity bond. In this model, real bond prices in logs are affine in the state variables: $p_t^{i,(n)} = -C_0^{i,n} - C_1^n x_t - C_2^{i,n}u_t - C_3^{i,n}w_t$. The conditional entropy of the permanent SDF component satisfies

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

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 $C_3^\infty = -\frac{A_3+\frac{1}{2}(B_2-C_1^\infty)^2}{1-\phi^w}$. This is naturally the same expression as above, but emphasizing the link to the infinite-maturity bond coefficients.

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 $w_t$, 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.

Condition 1 implies that ${B_1}^2 u_t + (B_2 + c_1)^2 w_t + (B_3 + c_2\sigma^u)^2 + (B_4 + c_3\sigma^w)^2 = {B_1}^{*2} u_t^* + (B_2^* + c_1^*)^2 w_t^* + (B_3^* + c_2^*\sigma^{u*})^2 + (B_4^* + c_3^*\sigma^{w*})^2$.

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 $B_1 = 0$ and $B_2 + c_1 = 0$. 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: 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 et al., 2018, for a multi-country LRR model with common shocks). In this case, Condition 1 implies that ${B_1}^2 u_t + (B_2 + c_1)^2 w_t + (B_3 + c_2\sigma^u)^2 + (B_4 + c_3\sigma^w)^2 = {B_1}^{*2} u_t + (B_2^* + c_1^*)^2 w_t + (B_3^* + c_2^*\sigma^{u*})^2 + (B_4^* + c_3^*\sigma^{w*})^2$. 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 $B_2 + c_1 = B_2^* + c_1^*$ (or equivalently, $B_2 - C_1^\infty = B_2^* - C_1^{*,\infty}$). 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: $(B_3 + c_2\sigma^u)^2 + (B_4 + c_3\sigma^w)^2 = (B_3^* + c_2^*\sigma^{u*})^2 + (B_4^* + c_3^*\sigma^{w*})^2$.

## C 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

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

is the global component of the 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 $\lambda$ 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

$$H_t = H_* + \hat{H}_t = p_t E^D_t \left[B^{-\gamma}_{t+1}F_{t+1} - 1\right],$$

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

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

with the conditional expectation of $\varepsilon^H$ 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 $\phi$ of the form

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

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

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

which yields

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

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

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

so the transitory SDF component is

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

and the permanent SDF component is

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

The transitory SDF component is only exposed to resilience shocks ($\varepsilon^H$), 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

$$L_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) = \log E_t\left(\frac{\Lambda_{t+1}}{\Lambda_t}\right) - E_t\left(\log\frac{\Lambda_{t+1}}{\Lambda_t}\right) = L_t\left(\frac{\Lambda_{t+1}^*}{\Lambda_t^*}\frac{\omega_{t+1}}{\omega_t}\right) + L_t\left(\frac{1+Ax_{t+1}}{1+Ax_t}\right).$$

After some algebra, we get

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

Similarly, the conditional entropy of the permanent SDF component is

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

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

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

Condition 1 implies that $\log(1+H_t) - p_t E_t^D\left[\log(B_{t+1}^{-\gamma} F_{t+1})\right] + L_t\left(\frac{c+x_{t+1}}{c+x_t}\right) = \log(1+H_t^*) - p_t E_t^D\left[\log(B_{t+1}^{-\gamma} F_{t+1}^*)\right] + L_t\left(\frac{c^*+x_{t+1}^*}{c^*+x_t^*}\right)$, where the probability and intensity of a disaster ($p_t$ and $B_{t+1}^{-\gamma}$) are common across countries, but the processes for resilience and productivity growth (and, thus, the parameters $h_*$, $\phi_H$ and $g_\omega$ and the variables $H_t$ and $F_{t+1}$) are country-specific.

First, we consider a version of the model in which the parameters are the same in each country, but the shocks are country-specific. In this case, Condition 1 is not satisfied because country-specific shocks (for example, on resilience) affect the entropy of the permanent components of the SDFs.

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. Let us compare the two conditional entropies:

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

The first two terms are the same and thus any heterogeneity there would imply differences in long-term bond risk premia. But the last term is different. One could consider a model where $A \neq A^*$ (implying differences in carry trade risk premia at the short end of the yield curve), but $c = c^*$ (implying no differences in risk premia at the long end of the yield curve). Recall that $A \equiv \frac{e^{-R-\lambda+g_\omega+h_*}}{1-e^{-R-\lambda+g_\omega+h_*-\phi_H}}$. To obtain cross-country differences in the parameter $A$, while satisfying Condition 1, one could consider cross-country differences in the parameter $g_\omega$ (since parameters $R$ and $\lambda$ are common across countries in the model, and heterogeneity in $h_*$ would generate heterogeneity in $H_t$).

# VI 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).

## A 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 $\Lambda$ is a strictly positive process, adapted to $\mathcal{F}$, for which it holds that the time $t$ price of any payoff $\Pi_s$ 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 $\Lambda$ is a multiplicative functional and establish the connection between the multiplicative property of the pricing kernel process and the semigroup property of pricing operators $\mathbb{M}$.[10] 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 $\Lambda$ 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, $\phi$ is a principal (i.e. strictly positive) eigenfunction of the extended generator of $\mathbb{M}$ and $\beta$ is the corresponding eigenvalue (typically negative).[11] If, furthermore,

[10] A functional $\Lambda$ is multiplicative if it satisfies $\Lambda_0 = 1$ and $\Lambda_{t+u} = \Lambda_t\Lambda_u(\theta_t)$, where $\theta_t$ 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.

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

$\Lambda^{\mathbb{P}}$ is martingale, then the eigenpair $(\beta, \phi)$ also solves the principal eigenvalue problem:[12]

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

Conversely, if the expression above holds for a strictly positive $\phi$ 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

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

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, 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

[12] 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).

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 $\phi$ 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 $\beta$ 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 $\beta$ is the smallest eigenvalue associated with a principal eigenfunction in the Hansen and Scheinkman (2009) eigenfunction problem.

## B 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.[13] The per period log risk premium on a long position in foreign currency over $k$ periods consists of

[13] 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 yield spread minus the per period expected rate of depreciation over those $k$ periods:

$$E_t[rx^{FX}_{t\to t+k}] = y^{(k),*}_t - y^{(k)}_t - \frac{1}{k}E_t[\Delta s_{t\to t+k}]. \tag{8}$$

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: $E_t\left[\Delta s_{t\to t+1}\right] = r^{f,*}_t - r^f_t - E_t\left[rx^{FX}_{t+1}\right]$. Summing up over $k$ periods leads to:

$$E_t[\Delta s_{t\to t+k}] \quad = \quad E_t\left[\sum_{j=1}^{k}\left(r^{f,*}_{t+j-1} - r^f_{t+j-1}\right)\right] - E_t\left[\sum_{j=1}^{k} rx^{FX}_{t+j}\right]. \tag{9}$$

From Equations (8) and (9), it follows that the log currency risk premium over $k$ periods is given by:

$$E_t[rx^{FX}_{t\to t+k}] \quad = \quad (y^{(k),*}_t - y^{(k)}_t) + \frac{1}{k}\sum_{j=1}^{k} E_t\left(r^f_{t+j-1} - r^{f,*}_{t+j-1}\right) + \frac{1}{k}\sum_{j=1}^{k} E_t(rx^{FX}_{t+j}). \tag{10}$$

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) in the main text to show that the currency risk premium over $k$ periods depends on conditional SDF entropy:

$$E_t[rx^{FX}_{t\to t+k}] = \frac{1}{k}\left[L_t\left(\frac{\Lambda_{t+k}}{\Lambda_t}\right) - L_t\left(\frac{\Lambda^*_{t+k}}{\Lambda^*_t}\right)\right]. \tag{11}$$

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.[14] If the pricing kernel

[14] 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.

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 $\rho_j$ denotes the $j$-th autocorrelation (Cochrane, 1988).[15] 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:

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

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

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.

## C 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.[16]

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

[16] 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.

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 $\sigma^2$) 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.

# VII 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.

## A 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 (2019); 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 1.** *If the permanent SDF component is unconditionally lognormal, the cross-country covariance of the SDF' permanent components is bounded below by:*

$$cov\left(\log \frac{\Lambda_{t+1}^{\mathbb{P},*}}{\Lambda_t^{\mathbb{P},*}}, \log \frac{\Lambda_{t+1}^{\mathbb{P}}}{\Lambda_t^{\mathbb{P}}}\right) \quad \geq \quad E\left(\log \frac{R_{t+1}^*}{R_{t+1}^{(\infty),*}}\right) + E\left(\log \frac{R_{t+1}}{R_{t+1}^{(\infty)}}\right) - \frac{1}{2} var\left(\log \frac{S_{t+1}^{\mathbb{P}}}{S_t^{\mathbb{P}}}\right), \tag{12}$$

*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 (2018) 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 1 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.

## B 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

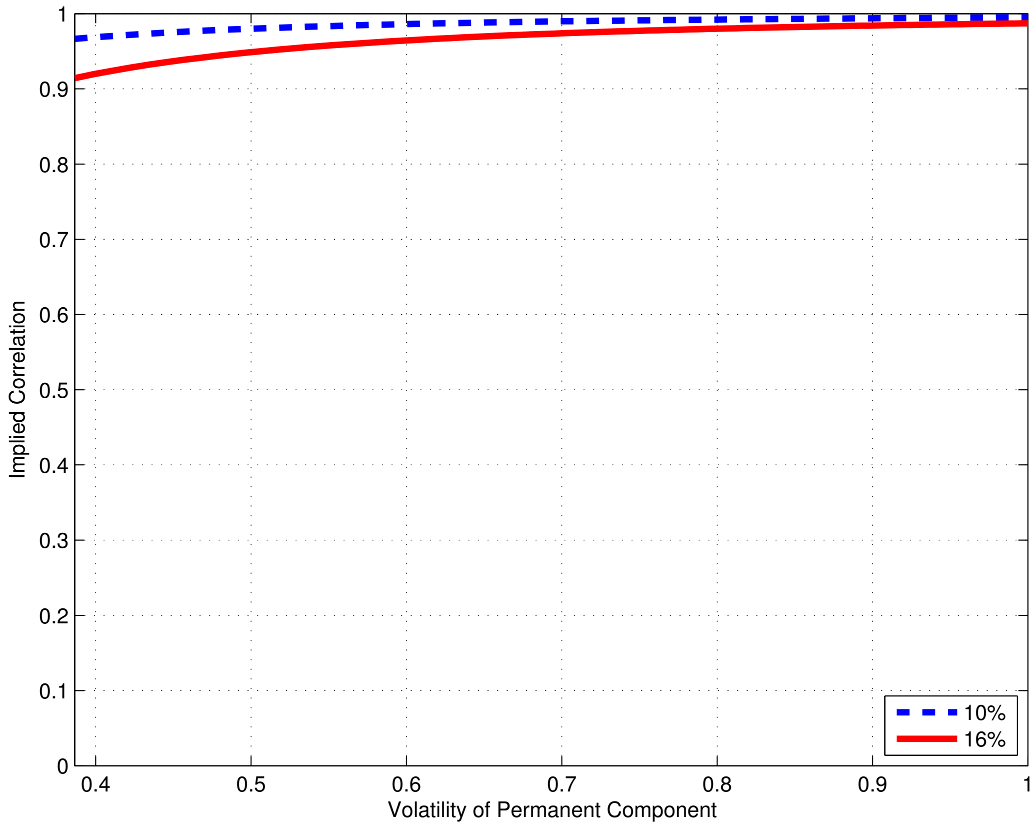


Figure A12: Cross-country Correlation of Permanent SDF Shocks — In this figure, we plot the implied correlation of the domestic and foreign permanent components of the SDF against the standard deviation of the permanent component of the SDF. The dotted line is for $Std\left(\log S_t^{\mathbb{P}}/S_{t+1}^{\mathbb{P}}\right) = 10\%$. 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.

account. When countries share permanent innovations to their SDFs, a simple long bond return parity condition emerges. The proposition below provides the result.

**Proposition 2.** *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:*

$$R_{t+1}^{(\infty),*}\frac{S_t}{S_{t+1}} = R_{t+1}^{(\infty)}, \text{ for all states.} \tag{13}$$

While Proposition 1 in the main text is about expected returns, Proposition 2 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.

## VIII 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.[17] 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 A25 reports the simulated moments.

We first consider the unconditional holding period bond returns across countries. The average (annualized)

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

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.

## Table A25: 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] |
| $\sigma_{m^*}$ | 239.17 | 241.92 | 127.14 | 118.45 | 211.76 | 132.76 | 227.59 | 153.22 |
| $corr(m, m^*)$ | 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.

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