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TAKING THE COCHRANE-PIAZZESI TERM STRUCTURE MODEL OUT OF SAMPLE: MORE DATA, ADDITIONAL CURRENCIES, AND FX IMPLICATIONS

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ABSTRACT

We examine the Cochrane and Piazzesi (2005, 2008) model in several out-of-sample analyzes. The model's one-factor forecasting structure characterizes the term structures of additional currencies in samples ending in 2003. In post-2003 data one-factor structures again characterize each currency's term structure, but we reject equality of the coefficients across the two samples. We derive some implications of the model for the predictability of cross-currency investments, but we find little support for these predictions in either pre-2004 or post-2003 data. The model fails to beat historical average returns in recursive out-or-sample forecasting of excess rates of return for bonds and currencies.

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A data appendix is available at http://www.nber.org/data-appendix/w25092

1 Introduction

In two seminal papers, Cochrane and Piazzesi (2005, 2008) document a strong one-factor structure in the unconstrained predictability of one-year-ahead excess returns on U.S. dollar zero-coupon bonds of several maturities. Cochrane and Piazzesi (2005) note (p. 142), "The same function of forward rates forecasts holding period returns at all maturities. Longer maturities just have greater loadings on this same function." To model this constrained system, they develop a two-step approach in which they first estimate the forecasting factor, which is labeled the 'CP factor' in much of the subsequent literature, by regressing the average future annual excess rates of return on two, three, four, and five year bonds onto a set of forward rates or forward spreads. Then, they regress each excess return on the forecasting factor to get the factor loadings. The constrained model fits the data remarkably well. They also demonstrate that their bond market forecasting factor predicts excess returns in the U.S. stock market, which strengthens the case that it is capturing risk premiums. Cochrane and Piazzesi (2008) reverse engineer an affine term structure model (ATSM) that has the forecasting properties uncovered in the constrained regressions.

This paper examines whether analogous one-factor forecasting structures exist in the predictability of the excess returns on zero-coupon bonds denominated in other currencies, and we find that they do. We initially examine samples that end in 2003, the end of the sample in the original paper. While the factor loadings are quite similar across currencies, the coefficients of the CP factors are not. We then examine data from 2004-2016 and again find a strong one-factor forecasting structure with factor loadings that are quite similar to those of the earlier sample, but the data do not support the hypothesis of equality of the coefficients in the CP factors across the two samples.

Because foreign exchange rates and the term structures of interest rates in the two currencies are closely linked in theory to the stochastic discount factors of the two currencies, we derive predictions from the Cochrane and Piazzesi (2008) ATSM for the excess rate of return on uncovered foreign currency investments. We find that the CP factors from the bond markets of the two currencies and their squared values should forecast the excess rate of return on uncovered foreign currency investments between the two currencies. We investigate this prediction empirically and find that they do not. While this evidence could be viewed as supporting uncovered interest rate parity at the annual horizon, we also find that the standard projection of these excess returns onto the one-year interest differential does show significant forecasting power. In this analysis, though, we also show substantial differences in estimated coefficients across our two sub-samples.

We then explore recursive out-of-sample predictions of the Cochrane and Piazzesi (2005) model and find considerable evidence of instability in the coefficients of the CP factors. Recursive forecasts of excess rates of return from the estimated model are generally unable to beat the recursive forecasts from the historical averages of excess rates of return for both bonds and currencies.

While these findings are perhaps unsurprising given that the out-of-sample period contains the global financial crisis, they demonstrate the necessity of modeling risk premiums while allowing for structural change. We leave this challenging task for future research.

2 Related Literature

The Cochrane and Piazzesi (2005, 2008) papers spawned a vast literature. In this section we briefly review what we consider to be the most important contributions in the literature that are related to our paper.[1]

Dahlquist and Hasseltoft (2013, 2016) and Sekkel (2011) were the first to extend the Cochrane and Piazzesi (2005) model to the bond markets of additional currencies. Dahlquist and Hasseltoft (2013) examine the bond markets of the USD; the Swiss franc, CHF; the euro, EUR; and the British pound, GBP; as well as examining the dollar denominated returns on the foreign bonds. They use a sample period from January 1975 to December 2009, and the CP factor is constructed from projections onto the five forward rates as in the original paper. They estimate local currency CP factors, and they also construct a global CP factor as a GDP-weighted average of the local CP factors. They find that the global CP factor provides some additional explanatory power relative to the local CP factors. In Dahlquist and Hasseltoft (2016) they extend their analysis adding the bond markets of the Australian dollar, AUD; the Canadian dollar, CAD; the Danish kroner, DKK; the Japanese yen, JPY; the Norwegian krone, NOK: and the Swedish krona, SEK; and they employ a sample period from December 1999 to December 2013. They find support for the model in all currencies, but they do not investigate the stability of the coefficients.

Wright (2011) examines the term structures of interest rates for the G-10 countries by estimating ATSMs as in Joslin, Priebsch and Singleton (2014). He studies the implied risk premiums or term premiums, defined as the difference between the long-term yields and expectations of future spot interest rates, finding that these term premiums have generally declined in most countries over the sample period from January 1990 to May 2009. Bauer, Rudebusch and Wu (2014) dispute these conclusions noting that after correcting for small sample bias in the coefficient estimates, the term premiums show a pronounced countercyclical pattern as was found by Cochrane and Piazzesi (2005).

Sekkel (2011) uses the Wright (2011) data to estimate the Cochrane and Piazzesi (2005) model, but he projects the excess returns only onto the one, three, and five year forward rates. He finds that the performance of the model deteriorates during the global financial crisis.

Consistent with the finding of Cochrane and Piazzesi (2005) that the CP factor is not spanned by the first three principal components of bond yields, Duffee (2011) documents that almost half of the variation in U.S. dollar (USD) bond risk premiums cannot be detected using the cross-section of yields. He finds that fluctuations in this hidden component have strong forecasting power for both future short-term interest rates and excess bond returns. The hidden component is negatively correlated with aggregate economic activity, but macroeconomic variables explain only a small fraction of variation in the hidden factor.

Koijen, Lustig and Van Nieuwerburgh (2017) model the stochastic discount factor as depending on the Cochrane and Piazzesi (2005) forecasting factor as well as the return on the stock market and the level of the term structure of interest rates. They demonstrate that such a model does well in simultaneously pricing returns on value and growth stocks in additional to USD zero-coupon bonds.

Kessler and Scherer (2009), Thornton and Valente (2012), Zhu (2015), and Sarno, Schneider and Wagner (2016) perform out-of-sample forecasting analyses with the Cochrane and Piazzesi (2005) model. Kessler and Scherer (2009) assess the performance of trading strategies based on a one-month forecast horizon using data from seven currencies (the AUD, CAD, CHF, EUR, GBP, JPY, and the USD) for the sample period

[1]Because the Cochrane and Piazzesi (2005, 2008) papers have 1,327 and 305 Google Scholar citations, respectively, as of August 2018, our literature review must be highly selective.

February 1997 to July 2007. They use either a 36 or 60 month rolling window to estimate the parameters of the forecasting equation implying that they have either 88 or 64 true out-of-sample forecasts. They find slightly positive but only marginally significant trading profits.

Thornton and Valente (2012) investigate the out-of-sample predictability of USD bond excess returns and assess the economic value of the forecasting ability of empirical models based on Fama and Bliss (1987) and Cochrane and Piazzesi (2005). Their results show that the information content of forward rates does not generate systematic economic value to investors in a dynamic asset allocation exercise. Furthermore, they find that the models do not outperform the no-predictability benchmark, and their relative performance deteriorates over time.

Zhu (2015) explores the forecasting ability of a global CP factor constructed as the forecast of the average returns on the two through five year maturity bonds averaged over four currencies (the EUR, JPY, GBP, and the USD) when regressed on the four individual currency CP factors. The full sample period is January 1980 to December 2011, and the out-of-sample period begins in January 1992. In contrast to our findings, Zhu (2015) finds statistically significant out-of-sample forecasts that beat the historical mean return for all four countries.

Sarno, Schneider and Wagner (2016) find for the USD bond market that the time-varying risk premiums implied by ATSMs do not provide important increases in utility to investors over and above inferences about expected future spot interest rates implied by the expectations hypothesis of the term structure with constant risk premiums.

Turning to the international implications of the modeling, Sarno, Schneider and Wagner (2012) find that separately estimated ATSMs for two currencies, both of which provide very small pricing errors for zero-coupon bonds denominated in those currencies, are not highly correlated with the relative rate of appreciation of those currencies in the foreign exchange market.

Jotikasthira, Le and Lundblad (2015) document that yield curve fluctuations across different currencies are highly correlated. They argue that common macroeconomic shocks influence bond yields both through a monetary policy channel and through a risk compensation channel. Using data from the U.S., the UK, and Germany, they find that world inflation and the level of the U.S. yield curve explain over two-thirds of the covariation of yields at all maturities and that these effects operate largely through the risk compensation channel for long-term bonds.

Pericoli and Taboga (2012) propose a two-country no-arbitrage term-structure model to analyze the joint dynamics of bond yields, macroeconomic variables, and the exchange rate. The model demonstrates how exogenous shocks to the exchange rate affect the yield curves, how bond yields co-move in different countries and how the exchange rate is influenced by interest rates, macroeconomic variables and time-varying bond risk premiums. Upon estimating the model with U.S. and German data, they find that time-varying bond risk premiums account for a significant portion of the variability of the exchange rate.

Our results are also related to the vast literature examining the uncovered interest rate parity (UIRP) hypothesis. Although Chinn and Meredith (2004) provide support for UIRP at the annual horizon, our results are more consistent with the conclusions of Bekaert, Wei and Xing (2007), who argue that UIRP is violated at longer horizons just as is typically the case at the shorter monthly horizon.

The UIRP puzzle concerns the empirical regularity that countries with high nominal interest rates tend to have high expected returns on uncovered short term deposits. Engel (2016) notes that countries with high real interest rates tend to have currencies that are stronger than can be accounted for by the path of expected real interest differentials under UIRP. He observes that these two findings have contradictory implications for the relationship of the foreign-exchange risk premium and interest-rate differentials and shows that existing models appear unable to account for both puzzles. He then introduces a model, in which short-term assets can have liquidity premiums as in Nagel (2016), that potentially reconciles the two sets of findings.

3 The Cochrane-Piazzesi Term Structure Model

In presenting the model, we mostly adopt the notation of Cochrane and Piazzesi (2005). The presentation can be thought of as referring to the term structure of a generic currency. For simplicity, we suppress currency subscripts in laying out the basic term structure model.

The natural logarithm of the price of a pure discount bond at time t that matures in n years and pays one unit of currency at that time is denoted $p_t^{(n)}$. The time subscript t indexes years, in which case months, which are the observation interval of the data, are indicated with (1/12) fractions of a year. The continuously compounded annualized yield on an n-year bond is therefore

$$y_t^{(n)} \equiv -\frac{1}{n} p_t^{(n)}.$$

The natural logarithm of the one-year forward rate at time t for loans between $t+n-1$ and $t+n$ is

$$f_t^{(n)} \equiv p_t^{(n-1)} - p_t^{(n)}.$$

The forward spreads between these forward rates and the one-year yield are

$$fs_t^{(n)} \equiv f_t^{(n)} - y_t^{(1)}.$$

The continuously compounded rate of return from buying an n-year bond at time t and selling it one year later is

$$r_{t+1}^{(n)} \equiv p_{t+1}^{(n-1)} - p_t^{(n)},$$

in which case the excess rate of return is

$$rx_{t+1}^{(n)} \equiv r_{t+1}^{(n)} - y_t^{(1)}.$$

The average of four excess rates of return on bonds with two through five years to maturity is

$$\overline{rx}_{t+1} \equiv (1/4) \sum_{n=2}^{5} rx_{t+1}^{(n)}.$$

Bold symbols without superscripts indicate vectors or matrices. For example, the vector of excess rates of return on bonds with two through five years to maturity is

$$\mathbf{rx}_{t+1} \equiv \left[rx_{t+1}^{(2)},\ rx_{t+1}^{(3)},\ rx_{t+1}^{(4)},\ rx_{t+1}^{(5)} \right]^{\mathsf{T}}.$$

When used as right-hand-side variables in a regression, such vectors include a constant. For example,

$$\mathbf{fs}_t \equiv \left[1,\ fs_t^{(2)},\ fs_t^{(3)}, fs_t^{(4)}, fs_t^{(5)}\right]^\intercal .$$

Whereas Cochrane and Piazzesi (2005) use the levels of the forward rates as forecasting variables for the excess rates of return on bonds, we follow Cochrane and Piazzesi (2008) and use the averages of the three most recent monthly spreads as the forecasting variables:[2]

$$\overline{\mathbf{fs}}_t \equiv (1/3)\sum_{j=0}^{2} \mathbf{fs}_{t-(j/12)}.$$

The unconstrained forecasting system for the excess rates of return in a particular currency's bond market can therefore be written as

$$\mathbf{rx}_{t+1} = \boldsymbol{\beta}\overline{\mathbf{fs}}_t + \boldsymbol{\varepsilon}_{t+1}, \tag{1}$$

where $\boldsymbol{\beta}$ represents the (4×5) matrix of responses of excess returns to the forward spreads. Cochrane and Piazzesi (2005, 2008) motivate their constrained one-factor model of expected bond returns from the finding that the first principal component of the unconstrained expected returns in the system of equations (1) explains over 99% of the variance of these expected returns.

This constrained model of a vector of expected returns was first developed by Hansen and Hodrick (1983) and Gibbons and Ferson (1985) who postulated that a set of expected returns could be proportional to a common unobserved factor, v_t:

$$E_t\left(\mathbf{rx}_{t+1}\right) = \mathbf{b}v_t, \tag{2}$$

where $\mathbf{b} \equiv [b_2,\ b_3,\ b_4,\ b_5]^\intercal$. By projecting the unobserved factor onto some observed information, in this case $\overline{\mathbf{fs}}_t$, one can write

$$v_t = \boldsymbol{\gamma}^\intercal\overline{\mathbf{fs}}_t + \xi_t, \tag{3}$$

where by the properties of linear prediction, the error term, ξ_t, is orthogonal to the right-hand-side variables.

Substituting equation (3) into equation (2) and assuming rational expectations produces a constrained single factor forecasting system that can be written as

$$\mathbf{rx}_{t+1} = \mathbf{b}\boldsymbol{\gamma}^\intercal\overline{\mathbf{fs}}_t + \boldsymbol{\varepsilon}_{t+1}, \tag{4}$$

where $\boldsymbol{\varepsilon}_{t+1}$ now represents both the rational expectations forecast errors for each equation plus $\mathbf{b}\xi_t$. Estimation can be done with the generalized method of moments (GMM) of Hansen (1982) because $\boldsymbol{\varepsilon}_{t+1}$ is orthogonal to $\overline{\mathbf{fs}}_t$ Because $\mathbf{b}$ and $\boldsymbol{\gamma}^\intercal$ are multiplied together, some identifying constraint must be imposed on the estimation, and we follow Cochrane and Piazzesi (2005) in imposing the constraint on $\mathbf{b}$ that the average of the b_n's equals one:

$$(1/4)\sum_{n=2}^{5} b_n = 1.$$

[2] Cochrane and Piazzesi (2008) note that levels of forward rates have near unit root components which are unlikely to match up with rational risk premiums. Forward spreads are more likely to be stationary and hence to capture risk premiums. See also the discussion in Cochrane (2015) who advocates using moving averages of forward spreads to avoid spurious predictability due to measurement error in the yields.

Whereas the unconstrained model in equation (1) has 20 parameters, the constrained model in equation (4) has 8 free parameters, 5 in $\boldsymbol{\gamma}$ and 3 in $\mathbf{b}$.

As Cochrane and Piazzesi (2005) note, estimation of the constrained model can be done in two steps. The first step is an OLS regression of the average excess rate of return on the four long-horizon bonds on the average of the forward spreads as in

$$\overline{rx}_{t+1} = \boldsymbol{\gamma}^{\intercal}\overline{\mathbf{fs}}_t + \bar{\varepsilon}_{t+1}. \tag{5}$$

This imposes the constraint that the average of the b_n's equals one. The second step involves OLS regressions without constant terms of three individual excess rates of return on the fitted value from equation (5):

$$rx_{t+1}^{(n)} = b_n\left(\widehat{\gamma}^{\intercal}\overline{\mathbf{fs}}_t\right) + \varepsilon_{t+1}^{(n)}, \tag{6}$$

and we use the two-year, three-year, and four-year maturities.

3.1 The Affine Model with Restrictions

Before discussing the results of estimating the constrained model, we first introduce the affine term structure model that Cochrane and Piazzesi (2008) reverse engineer to be consistent with the forecasting properties from the constrained regressions of excess returns of the long-term bonds on forward spreads.

In a generic ATSM the continuously compounded short-term interest rate is postulated to be a linear function of a K-dimensional vector of state variables, $\mathbf{X}_t$:

$$r_t = \delta_0 + \boldsymbol{\delta}_1^{\intercal}\mathbf{X}_t.$$

The state variables are assumed to follow a first-order vector autoregression:

$$\mathbf{X}_{t+1} = \boldsymbol{\mu} + \boldsymbol{\Phi}\mathbf{X}_t + \boldsymbol{\Sigma}\boldsymbol{\upsilon}_{t+1}.$$

The vector of innovations, $\boldsymbol{\upsilon}_{t+1}$, is assumed to be $N(0, \boldsymbol{I}_K)$, and the covariance matrix of the state variables is $\boldsymbol{\Sigma}\boldsymbol{\Sigma}^{\intercal}$. The natural logarithm of the stochastic discount factor is specified to be

$$m_{t+1} = -r_t - \frac{1}{2}\boldsymbol{\lambda}_t^{\intercal}\boldsymbol{\lambda}_t - \boldsymbol{\lambda}_t^{\intercal}\boldsymbol{\upsilon}_{t+1}, \tag{7}$$

and the innovations to the state variables are thus potential sources of risks. Finally, the prices of these risks are also postulated to be affine functions of the state variables:

$$\boldsymbol{\lambda}_t = \boldsymbol{\lambda}_0 + \boldsymbol{\lambda}_1\mathbf{X}_t,$$

where $\boldsymbol{\lambda}_0$ is $K \times 1$, and $\boldsymbol{\lambda}_1$ is $K \times K$.

The solution of such an affine term structure model uses the basic no-arbitrage asset pricing model,

$$E_t\left(M_{t+1}R_{t+1}^{(n)}\right) = 1, \tag{8}$$

where $M_{t+1} = \exp(m_{t+1})$ and $R_{t+1}^{(n)} = \exp\left(r_{t+1}^{(n)}\right)$. Substituting for M_{t+1} and $R_{t+1}^{(n)}$ in equation (8) and solving the conditional expectation provides the solution of the ATSM in which the natural logarithms of the bond prices are found to be affine functions of the state variables:

$$p_t^{(n)} = A_n + \boldsymbol{B}_n^{\intercal}\mathbf{X}_t. \tag{9}$$

The recursive formulas for the A_n and $\boldsymbol{B}_n$ coefficients in equation (9) are given in Appendix B.

From the solution of the the ATSM, one finds that the expected excess rates of return on bonds are also affine functions of the state variables:

$$E_t\left(rx_{t+1}^{(n)}\right) = -\,(1/2)\,\boldsymbol{B}_{n-1}^{\intercal}\boldsymbol{\Sigma}\boldsymbol{\Sigma}^{\intercal}\boldsymbol{B}_{n-1} + \boldsymbol{B}_{n-1}^{\intercal}\boldsymbol{\Sigma}\boldsymbol{\lambda}_0 + \boldsymbol{B}_{n-1}^{\intercal}\boldsymbol{\Sigma}\boldsymbol{\lambda}_1\mathbf{X}_t. \tag{10}$$

The three terms on the right-hand side of equation (10) are a Jensen's inequality term related to the variance of the rate of return, a constant risk premium, and a time-varying risk premium. In the general ATSM without constraints on the parameters, time-varying expected excess rates of returns on bonds would be driven by the K state variables. This would be inconsistent with the empirical finding that only one state variable is required to forecast expected excess returns.

To reconcile the theoretical analysis with the empirical findings, Cochrane and Piazzesi (2008) postulate that the term structure of interest rates depends on four state variables, but they constrain the prices of risks such that only one of these variables drives expected excess rates of return. At least since Litterman and Scheinkman (1991) it has been known that time variation in zero-coupon bond yields can be effectively modeled with the first three principal components of the yields, which are a level effect, l_t, a slope effect, s_t, and a curvature effect, c_t. Hence, these three variables are present as state variables. The fourth state variable is the "return forecasting factor", that is, the CP factor:

$$x_t \equiv \widehat{\boldsymbol{\gamma}}^{\intercal}\overline{\mathbf{fs}}_t. \tag{11}$$

The state vector can therefore be written as $\mathbf{X}_t = (x_t,\ l_t,\ s_t,\ c_t)^{\intercal}$.[3] Because Cochrane and Piazzesi (2005) empirically find a very strong one-factor structure in the unconstrained model in equation (1), Cochrane and Piazzesi (2008) place a set of restrictions on the prices of risks, $\boldsymbol{\lambda}_t$, such that a one-factor structure emerges in equation (10). The restrictions on $\boldsymbol{\lambda}_t$ are the following:

$$\boldsymbol{\lambda}_t = \begin{bmatrix} 0 \\ \lambda_{0l} \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} 0 & 0 & 0 & 0 \\ \lambda_{1l} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} x_t \\ l_t \\ s_t \\ c_t \end{bmatrix}. \tag{12}$$

Thus, although innovations in the four state variables drive the zero-coupon yields and bond prices at all maturities, the only innovation that affects the bond market's stochastic discount factor and hence affects expected rates of return on bonds is the innovation in the level of the term structure, denoted $v_{l,t+1}$, and the time varying price of this risk is driven by the return forecasting factor. That is,

[3] Cieslak and Povala (2015) develop a similar ATSM with three state variables: the expected or 'trend' rate of inflation, a real factor orthogonal to expected inflation, and a forecasting variable that only affects the prices of the two risks.

$$\boldsymbol{\lambda}_t^{\intercal}\boldsymbol{\upsilon}_{t+1} = \begin{bmatrix} 0 \\ (\lambda_{0l} + \lambda_{1l}x_t)\,\upsilon_{l,t+1} \\ 0 \\ 0 \end{bmatrix} \tag{13}$$

Substituting from equation (12) into equation (10) gives

$$E_t\left(rx_{t+1}^{(n)}\right) = -\,(1/2)\,\boldsymbol{B}_{n-1}^{\intercal}\boldsymbol{\Sigma}\boldsymbol{\Sigma}^{\intercal}\boldsymbol{B}_{n-1} + \boldsymbol{B}_{n-1}^{\intercal}\boldsymbol{\Sigma}\begin{bmatrix} 0 \\ (\lambda_{0l} + \lambda_{1l}x_t) \\ 0 \\ 0 \end{bmatrix}. \tag{14}$$

While equation (14) is quite close to the constrained econometric model in equation (4) in that each expected return loads with a different coefficient onto the common forecasting factor, the constrained model makes the additional assumption that the constant terms in the equations share the same proportionality as the slope coefficients. The Jensen's inequality terms do not scale in the same way, which makes this assumption technically incorrect. Since these terms are generally considered to be small, in what follows we ignore this issue and follow the approach of Cochrane and Piazzesi (2008).[4]

4 Estimation Results for Nine Term Structures

In this section we estimate the Cochrane and Piazzesi (2005) model for the zero-coupon government bond yields of nine of the G-10 currencies: the AUD, CAD, CHF, EUR, GBP, JPY, NOK, SEK, and the USD. After reviewing the available term structure data for the New Zealand dollar, we viewed it as unreliable and therefore did not include it in our analysis. Sources of data are described in Appendix A.

We present the results in two sections corresponding to data that would have been available when the original model was first estimated and to data that subsequently became available. Because the last observation on the dependent variable in the the first data set is December 2003, we refer to these data as the pre-2004 sample. We begin observations on the dependent variable in the second data set in December 2004 to avoid overlap with the first data set, and we refer to these data as the post-2003 sample. To allow for samples that coincide with the exchange rate data, the dependent variables for the first sample begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The first sample is particularly short for the NOK, so we do not think those results are particularly informative, but we choose to include the results simply because the NOK is included in the post-2003 analysis.

4.1 Results with Pre-2004 Data

Table 1 reports the estimation of the constrained model in equation (4) with the two-step OLS procedure described above. We report asymptotic GMM standard errors that account for the overlapping forecasts

[4] The Online Appendix presents some results of a model that relaxes this restrictive assumption by allowing for separate constants at each maturity. We find that the relevant parameters associated with the time-varying forecasts of the two models are quite close and inference is quite similar.

and the fact that the second step in the estimation uses estimated coefficients from the first step.[5]

Although the unconstrained results are not reported because of the large number of parameters, the first thing to notice in Table 1 is the strong support for the single factor forecasting structure of expected excess returns in each of the nine term structures in the unconstrained estimations. The far right column labelled $\%PC1$ presents the proportion of the variance of the four unconstrained estimates of the excess rates of return, denominated in the particular currency of that row, that is explained by their first principal component. For all the currencies, the first principal component explains at least 98.8% of the variance of these expected excess returns. This evidence represents strong support for the one-factor forecasting model of expected excess bond returns in each of the currencies.

The second noteworthy aspect of Table 1 is the remarkable similarity in the coefficient estimates of b_2, b_3, and b_4. The estimated values of b_2 range from 0.37 for the JPY to 0.47 for the CHF. The estimated values of b_3 range from 0.80 for the SEK to 0.87 for the AUD. The estimated values of b_4 range from 1.19 for the CHF to 1.23 for the USD and the JPY. From equation (14) we see that the estimated values of the b_n's in an ATSM differ because of the different values of $\mathbf{B}_{\mathbf{n-1}}^{\intercal}\boldsymbol{\Sigma}$ associated with the CP factor. The recursive solution for the $\mathbf{B}_{\mathbf{n}}^{\intercal}$ in equation (B.3) indicates that values of $\mathbf{B}_{\mathbf{n}}^{\intercal}$ change as $\boldsymbol{\Phi}^*$, the risk neutral autocorrelation matrix of the state variables, is raised to higher powers. Thus, the finding of similar values of the b_n's across countries indicates that if we were to estimate an ATSM for each currency, the resulting $\boldsymbol{\Phi}^*$ estimates would be quite similar across currencies. At this point, we leave this as a conjecture for future research.

While there is considerable variety in the estimates of the γ_j's across the different currencies, the $\chi^2(4)$ statistics for all currencies except the SEK provide strong rejections of the currency-by-currency null hypothesis that the time-varying, right-hand-side variables have no collective forecasting power. Particularly large values of coefficients for the AUD, SEK, and NOK are an indication of multicollinearity. Although Cochrane and Piazzesi (2005) found a clear "tent" pattern in their projection of average returns onto the levels of the five forward rates, we only see this pattern in projections onto the four forward spreads for the USD and JPY data.

There are at least two reasons why the estimates of the $\boldsymbol{\gamma}$'s might differ across currencies. The first explanation takes a rational expectations econometrics view and recognizes that the forward spreads capture the risk exposures of a country as represented by the reduced form coefficients from an ATSM. Underlying structural differences in the nature of risks would consequently manifest themselves in different $\boldsymbol{\gamma}$'s. Monetary and fiscal policies certainly differ across countries, and we do not attempt to relate the underlying coefficients of the ATSM to more structural coefficients in equations such as the Taylor (1993) rule.

Alternatively, one can take the perspective of Bekaert, Hodrick and Marshall (2001) who argue that the rational expectations econometrics perspective may be too strong. Developed countries, such as those studied here, may actually be following the same time series rule, but the realizations of the shocks hitting the economies may have differed across countries. It may take a very long sample for a particular economy to experience all of the possible realizations from the policy rule with their ex ante frequencies that investors anticipated during the sample. It is certainly true that ex post experiences with inflation have differed across the countries, although at a casual level, all countries now seem to be converging to relatively low

[5]The standard errors could be constructed as in Hansen and Hodrick (1980), by equally weighting the 11 lagged covariances that are non-zero by construction when forecasting annual excess returns with overlapping monthly data. These standard errors are not guaranteed to be positive definite, and in fact in some cases they were not. Consequently, we rely on Newey and West (1987) standard errors using 18 lags as in Cochrane and Piazzesi (2005).

rates of expected inflation.

As an example of this last perspective, it is notable in Table 1 that the R^2's from the first-step regression of the average return on the forward spreads are the highest for the USD and JPY. Bekaert, Hodrick and Marshall (2001) argue that the decline in U.S. inflation under Federal Reserve Chairmen Volcker and Greenspan represents a one-sided realization that made the ex post returns on investments in long-term bonds better than was anticipated.[6] Inflation in Japan during much of the sample was also surprisingly low. Thus, the Japanese situation could be similar to the U.S. in that the stagnation in the Japanese economy and its ultimate experiences with deflation resulted in surprisingly good ex post returns on long-term Japanese bonds even though bond yields were quite low to start.

4.2 Results with Post-2003 Data

Table 2 presents analogous results to those of Table 1 but for the sample period from 2004 to 2016. While the one-factor structure of expected excess returns, estimated from unconstrained regressions, is not quite as strong in this sample, we still see that the first principal component of the expected returns explains between 86.6% of the variance for the NOK and 99.4% for the JPY. The remarkable similarity in the coefficient estimates of b_2, b_3, and b_4 is maintained. The estimated values of b_2 range from 0.28 for the CHF to 0.38 for the CAD; the estimated values of b_3 range from 0.73 for the JPY to 0.82 for the GBP and the CAD; and the estimated values of b_4 range from 1.22 for the EUR, CAD, AUD, SEK, and NOK to 1.29 for the JPY.

As a first step in analyzing the out-of-sample performance of the Cochrane and Piazzesi (2005) model, Table 3 presents tests of the equality of the vectors of b_n's and $\boldsymbol{\gamma}$'s across the two samples on a currency by currency basis. For the vector of b_n's, even though the coefficient estimates are quite similar across the two samples, the small standard errors lead to rejections of equality of the three coefficients for the EUR at the 1% marginal level of significance, for the CHF at the 3% level, and for the JPY at smaller than the 1% level. The tests of the vector of $\boldsymbol{\gamma}$'s rejects equality across the two periods for the USD, the JPY, and the NOK at less than the 1% level, for the GBP at the 9% level, and for the AUD at the 10% level. These findings provide the first evidence of instability in the forecasting relations.

4.3 Correlation Matrix and Variance Decomposition of Country CP Factors

Since one-factor forecasting structures characterize each of the term structures quite well, a natural question to ask is how correlated are the various CP factors. Table 4 provides a correlation matrix for the respective currency-specific CP factors for the pre-2004 sample period.

Of the 36 correlations, 26 are positive, but only the GBP-CHF correlation of 0.63 is larger than 0.50. Of the nine negative ones, the JPY-NOK correlation is the most negative at -0.30. The last column in Table 4 labelled $\%PC(i)$ reports the percent of the variance of the nine CP factors that is explained by the respective principal components. The first three principal components explain 82% of the total variance. While this evidence is suggestive that global risk factors may be at work in explaining the ability of the CP factors to forecast excess bond returns, it is certainly not definitive.[7]

[6] See Bauer and Rudebusch (2017) for an analysis of the U.S. term structure that allows for declining stochastic trends in both the long-run expected rate of inflation and the equilibrium real interest rate.

[7] Jotikasthira, Le and Lundblad (2015) investigate the determinants of the correlations across several major currency term structures.

When we examine the post-2003 samples in Table 5, we find that six of the 36 correlations are negative, and the largest positive correlation is now the GBP-NOK correlaiton of 0.54, which is the only correlation greater than 0.50. Twelve of the correlations change sign, and the largest switch is the GBP-EUR correlation which increased from -0.19 to 0.39. The share of the variance explained by the first three principal components falls to 69%. These changes in correlations are another indication of instability in the model.

We will examine out-of-sample forecasting of bond returns below, but first, we examine some international implications of the model.

5 International Implications

This section derives some implications of the Cochrane and Piazzesi (2005, 2008) model for foreign exchange markets. Doing so requires the introduction of subscripts for the currencies, and we subscript the USD variables with a one and variables denominated in an arbitrary foreign currency with a j. We define exchange rates as $S_{ij,t}$, which represents the currency j price of base currency i at time t. The continuously compounded rate of appreciation of base currency i relative to currency j between times t and $t+1$ is denoted $\Delta s_{ij,t+1}$.

We first argue that tight restrictions between the term structure models of the two currency markets and the relative rate of currency appreciation are not supported empirically.[8] Then, we consider some less constrained empirical predictions.

To understand this argument, consider the basic no arbitrage asset pricing equation for a particular currency that must price all returns denominated in that currency as in equation (8); but now, let Q_{t+1} represent the SDF that prices these generic returns, R_{t+1}, which include other assets and not just the bond market returns of equation (8). Thus, we have

$$E_t\left(Q_{t+1}R_{t+1}\right) = 1. \tag{15}$$

The difference between the SDF in equation (15), Q_{t+1}, and the SDF in equation (8), M_{t+1}, is that Q_{t+1} can contain risks that are orthogonal to the risks that are priced in the term structure of interest rates through M_{t+1}. Analytically, we can decompose Q_{t+1} as

$$Q_{t+1} = M_{t+1}Z_{t+1}. \tag{16}$$

Consistency of the two no arbitrage conditions requires that $E_t\left(Z_{t+1}\right) = 1$, because the risk free rate is correctly priced by M_{t+1}; $E_t\left(M_{t+1}Z_{t+1}\right) = E_t\left(M_{t+1}\right)E_t\left(Z_{t+1}\right)$, because Z_{t+1} and M_{t+1} are orthogonal; and for bond returns, $R_{t+1}^{(n)}$, $E_t\left(Z_{t+1}R_{t+1}^{(n)}\right) = E_t\left(Z_{t+1}\right)E_t\left(R_{t+1}^{(n)}\right)$ because M_{t+1} contains all risks priced in the bond market making Z_{t+1} orthogonal to $R_{t+1}^{(n)}$.

[8]See Backus, Foresi and Telmer (2001) for a discussion of the links between fully specified SDF's and the rate of currency appreciation when financial markets are complete, and see Brandt and Santa-Clara (2002) for a discussion of the effects of incomplete markets.

5.1 Implications for the innovation in currency appreciation

If markets are complete, it is well known that there is a tight relation between the rate of appreciation of currency i relative to currency j and the difference between the natural logarithms of the stochastic discount factor of currency i, $q_{i,t+1}$, and the stochastic discount factor of currency j, $q_{j,t+1}$:

$$\Delta s_{ij,t+1} = q_{i,t+1} - q_{j,t+1}. \tag{17}$$

Substituting for the q's gives

$$\Delta s_{ij,t+1} = m_{i,t+1} + z_{i,t+1} - m_{j,t+1} - z_{j,t+1}, \tag{18}$$

where $z_{i,t} \equiv \ln(Z_{i,t})$. Notice that if the Cochrane and Piazzesi (2005, 2008) ATSM correctly characterized the term structure in each currency, if asset markets were complete, and if the term structure SDF's contained all the sources of risks, then the z's could be eliminated from equation (18). After substituting for the innovations in the m's from equation (13), the innovation in the rate of appreciation of currency i relative to currency j would be

$$\Delta s_{ij,t+1} - E_t\left(\Delta s_{ij,t+1}\right) = \left(\lambda_{j,0l} + \lambda_{j,0l} x_{j,t}\right) \upsilon_{j,l,t+1} - \left(\lambda_{i,0l} + \lambda_{i,0l} x_{i,t}\right) \upsilon_{i,l,t+1}. \tag{19}$$

Thus, the innovation in $\Delta s_{ij,t+1}$ would be fully explained by the innovations in $m_{j,t+1}$ and $m_{i,t+1}$. In the Cochrane and Piazzesi (2008) ATSM, the innovations in the SDF's are innovations in the level factors interacted with a constant and the predetermined CP factors. We investigate this issue for rates of appreciation of the USD versus the other eight currencies in Table 6. Because the exact fit of equation (19) would be unlikely to hold, we run regressions with the expectation that if the model were true, we would have quite significant explanatory power. We proxy the innovation in the rate of appreciation of the USD with respect to currency j with the excess rate of return on a USD investment in the currency j money market, $-\Delta s_{1j,t+1} + r_{j,t} - r_{1,t}$. We proxy the innovations in the level factors with the changes in the levels, as represented by the first principal components of the term structures, because these first principal components are highly serially correlated. For simplicity, we also just report results for the full sample periods associated with each currency. In the regressions in Table 6 the R^2's range from 2% for the CAD and the CHF to 23% for the JPY. This represents strong evidence that the constrained Cochrane and Piazzesi (2005) term structure models do not span the spaces of risks that characterize the rates of currency depreciation, which we interpret as evidence for the presence of additional risks in the SDF's that price all assets.

These results are consistent with the analysis of Sarno, Schneider and Wagner (2012) who estimate four-factor, latent variable ATSM's for the bond markets of two currencies and find that while the bonds are priced very well, the variation of the rate of currency appreciation from the implied ATSM stochastic discount factors does not match well with the actual rate of currency appreciation. Of course, the results could also indicate that financial markets are incomplete as in the analysis of Brandt and Santa-Clara (2002).

5.2 Implications for expected cross-currency investments

To investigate expected rates of return on cross-currency investments that are implied by the model with $Z_{i,t+1}$ present, let $Z_{i,t+1}$ be log-normally distributed. Then, we can assume that stochastic process for $z_{i,t+1}$ is given by

$$z_{i,t+1} = -\frac{1}{2}\boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{\lambda}_{z_i,t} - \boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{v}_{z_i,t+1}, \tag{20}$$

where $\boldsymbol{v}_{z_i,t+1}$ is a vector of risks that are distributed $N(0, I)$ and that are orthogonal to the vector of risks, $\boldsymbol{v}_{i,t+1}$, that drive the term structure of interest rates in that currency.

Substituting for the SDF's from equations (7) and (20) and rearranging terms gives the excess rate of return in currency i on a one-year investment in the money market of currency j:

$$\begin{aligned} -\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = {} & \frac{1}{2}\left(\boldsymbol{\lambda}_{i,t}^{\intercal}\boldsymbol{\lambda}_{i,t} - \boldsymbol{\lambda}_{j,t}^{\intercal}\boldsymbol{\lambda}_{j,t}\right) + \frac{1}{2}\left(\boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{\lambda}_{z_i,t} - \boldsymbol{\lambda}_{z_j,t}^{\intercal}\boldsymbol{\lambda}_{z_j,t}\right) \\ & + \boldsymbol{\lambda}_{i,t}^{\intercal}\boldsymbol{v}_{i,t+1} - \boldsymbol{\lambda}_{j,t}^{\intercal}\boldsymbol{v}_{j,t+1} + \boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{v}_{z_i,t+1} - \boldsymbol{\lambda}_{z_j,t}^{\intercal}\boldsymbol{v}_{z_j,t+1}. \end{aligned} \tag{21}$$

Taking the conditional expectation of equation (21) gives

$$\mathrm{E}_t\left(-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t}\right) = \frac{1}{2}\left(\boldsymbol{\lambda}_{i,t}^{\intercal}\boldsymbol{\lambda}_{i,t} - \boldsymbol{\lambda}_{j,t}^{\intercal}\boldsymbol{\lambda}_{j,t}\right) + \frac{1}{2}\left(\boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{\lambda}_{z_i,t} - \boldsymbol{\lambda}_{z_j,t}^{\intercal}\boldsymbol{\lambda}_{z_j,t}\right). \tag{22}$$

The right-hand side of equation (22) is the expected excess rate of return to borrowing one unit of currency i, investing that amount in the currency j money market, and bearing the foreign exchange risk.[9] By imposing the constraints of the one-factor forecasting model for the two bond markets in equation (12), we find

$$\boldsymbol{\lambda}_{j,t}^{\intercal}\boldsymbol{\lambda}_{j,t} = \left(\lambda_{j,0l} + \lambda_{j,1l}x_{j,t}\right)^2 = \lambda_{j,0l}^2 + 2\lambda_{j,0l}\lambda_{j,1l}x_{j,t} + \lambda_{j,1l}^2 x_{j,t}^2. \tag{23}$$

Substituting from equation (23) into equation (22) implies that the return forecasting CP factors, $x_{i,t}$ and $x_{j,t}$, from the bond markets of the two currencies and their squared values should forecast the excess rate of return to investing a unit of currency i in the currency j money market while bearing the foreign exchange risk:

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \psi_0 + \psi_1 x_{i,t} + \psi_2 x_{i,t}^2 + \psi_3 x_{j,t} + \psi_4 x_{j,t}^2 + \epsilon_{ij,t+1}^s. \tag{24}$$

We leave the regression coefficients in equation (24) unconstrained because we do not observe $\boldsymbol{\lambda}_{z_i,t}^{\intercal}\boldsymbol{\lambda}_{z_i,t} - \boldsymbol{\lambda}_{z_j,t}^{\intercal}\boldsymbol{\lambda}_{z_j,t}$. Although equation (23) demonstrates that the return forecasting factors and their squared values should forecast the excess rate of return in the currency j money market with tight restrictions related to the prices of risks, the return forecasting variables may also enter the determination of the prices of risks, $\boldsymbol{\lambda}_{z_i,t}$ and $\boldsymbol{\lambda}_{z_j,t}$, or they may simply be correlated with the variables that drive these prices of risks, in which case OLS regression of the excess rate of return on $x_{i,t}$ and $x_{j,t}$ and their squared values does not isolate the pure effect of these variables that arises strictly from the fact that they are the determinants of the prices of the term structure risks, $\boldsymbol{\lambda}_{i,t}$ and $\boldsymbol{\lambda}_{j,t}$. Any restrictions arising from an ATSM specification of $\boldsymbol{\lambda}_{i,t}$ and $\boldsymbol{\lambda}_{j,t}$ are lost in the general regression specification in equation (24) because the determinants of $\boldsymbol{\lambda}_{z_i,t}$ and $\boldsymbol{\lambda}_{z_j,t}$

[9] As in equation (10), one can also express this time-varying expected excess rate of return in terms of a Jensen's inequality term and a logarithmic risk premium term.

are not included in the regression.

Table 7 presents the estimated coefficients for equation (24) with their asymptotic standard errors in parenthesis for the pre-2004 sample.[10] Panel A presents the forecasts of the USD excess rates of return from investments in the eight different currencies. Then, Panels B-G present the remaining 28 non-redundant cross-currency forecasts of excess rates of return that do not involve the USD. Each Panel is labeled with its base currency. The statistical significance of the estimates of the ψ_k coefficients is quantified with three different tests. The $\chi^2(2)_i$ statistic tests the null hypothesis that ψ_1 and ψ_2 equal zero, which tests whether the CP factor associated with base currency i and its squared value have forecasting power for the excess rate of return on an investment of base currency i in the currency j money market; the $\chi^2(2)_j$ statistic tests the null hypothesis that ψ_3 and ψ_4 equal zero, which tests whether the CP factor associated with currency j and its squared value have forecasting power; and the $\chi^2(4)$ statistic tests the null hypothesis that ψ_1 through ψ_4 equal zero. Failure to reject these hypotheses would be consistent with the absence of time varying foreign exchange risk premiums as specified, for example, in the uncovered interest rate parity hypothesis.

In Panel A, only for the tests associated with the JPY do we find sufficiently large test statistics to reject the three null hypotheses that the USD CP factor and its squared value as well as the foreign CP factor and its squared value are not significant determinants of the expected annual excess rates of return on investments in the foreign money markets. The adjusted R^2 in the JPY regression is also a substantial 0.38. For the other currencies, two of the $\chi^2(2)$ statistics have p-values less than the 0.1 marginal level of significance, but such a finding would be expected by chance when examining 21 statistics. The other adjusted R^2's in Panel A range from -0.00 for the SEK to 0.19 for the CAD.

For the cross-currency results that do not involve the USD in Panels B-G, only 7 of the 63 test statistics not involving the NOK are larger than the critical value of a χ^2 statistic associated the 0.1 marginal level of significance. Thus, the overall results do not support the ability of the CP factors to predict the excess rates of return in foreign money markets in the sample of data that would have been available in 2005.

How does the model do in the post-2003 sample? The answer is not particularly well. These results are presented in Table 8, which has the same format at Table 7. Overall, the statistical significance of the CP factors and their squared values is little better than chance as only 17% (12 of the 72) of the $\chi^2(2)$ statistics have a $p-$value smaller than 0.1. For example, with the USD as the base currency in Panel A, only two of the USD CP factor tests, in the CAD and NOK regressions, and none of the non-USD CP factor tests have $p-$values smaller than 0.1. The adjusted R^2's range from 0.11 for the AUD to 0.36 for the NOK.

The results are similar for the other non-USD panels. For the GBP results in Panel B, the CHF CP factor is the only non-GBP CP factor to have a $p-$value less than 0.1, and the GBP CP factor is only statistically significant in the CAD and JPY forecasts. The adjusted R^2's range from 0.07 for the EUR and the SEK to 0.30 for the JPY. In the remaining Panels, only for the CHF in Panel D is there much in the way of statistical significance as four out of the 10 tests for the CHF factors have $p-$values smaller than 0.1.

Because the CP factors are correlated, it could be the case that multicollinearity leads to insignificant tests of individual country CP factors but joint significance across the CP factors. Upon examining the $\chi^2(4)$ statistics, we see that 11 of the 36 statistics have $p-$values smaller than 0.1, but seven of these are associated with the CHF. Thus, we are left with the overall impression from these data that annual excess rates of return in foreign exchange markets are essentially unpredictable.

[10] Appendix C derives the standard errors of the parameters in equation (24). These standard errors allow for the fact that the forecasting variables are estimated in first stage regressions.

5.3 Uncovered Interest Rate Parity

Although the CP factors and their squared values are unable to forecast excess rates of return on international money market investments, this finding does not arise because uncovered interest rate parity is supported by the data in which case the excess rates of return would be completely unpredictable. To examine this issue, we run traditional regressions of these annual excess rates of return on the corresponding one-year interest differential as in the following:

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \phi_0 + \phi_1 \left(r_{j,t} - r_{i,t}\right) + \epsilon^s_{ij,t+1} \tag{25}$$

In these regressions, the null hypothesis of no predictability of excess returns is $\phi_1 = 0$. These regressions are analogous to the widely replicated regressions of Fama (1984) in which the rate of appreciation of base currency i relative to currency j is regressed on the interest differential between currencies j and i as in the following:

$$\Delta s_{ij,t+1} = \alpha + \beta \left(r_{j,t} - r_{i,t}\right) + \epsilon^s_{ij,t+1} \tag{26}$$

The relation between the two slope coefficients is $\beta = 1 - \phi_1$. Thus, the historical finding that estimated β's are negative in equation (26) translates into $\phi_1 > 1$ in equation (25).

Tables 9 and 10 present the results of estimating equation (25) for the two samples used above for all of the non-redundant currency pairs in our analysis. Each of the Tables contains two tests: one examining the null hypothesis $\phi_1 = 0$, and one examining the null hypothesis $\phi_1 = 1$.

Panel A of Table 9 presents the results with the USD as the base currency for the pre-2004 sample. For all currencies other than the SEK, we see estimates of ϕ_1 ranging from 1.68 for the EUR to 4.54 for the JPY.[11] These estimates provide strong rejections of the null hypothesis $\phi_1 = 0$ associated with unpredictability of the excess rates of return with $\chi^2(1)$ statistics ranging from 5.10 with a $p-$value of 0.02 for the EUR to 19.59 for the JPY with a $p-$value of 0.00. The point estimate for the SEK of 0.91 is insignificantly different from zero. For the currencies other than the SEK and the NOK, the R^2's range in value from 0.09 for the EUR to 0.39 for the JPY. Although all of the point estimates of ϕ_1 parameters that are significantly different from zero are larger than one, we are only able to reject the hypothesis that $\phi_1 = 1$ for the JPY.

The results in Panels B through F of Table 9 are broadly consistent with those in Panel A. With the GBP, the EUR, or the CHF as the base currency in Panels B-D, respectively, all of the estimates of ϕ_1's are greater than one. We see strong rejections of the hypothesis that $\phi_1 = 0$ for the following currency pairs: the GBP vs. the EUR, the CHF, and the JPY; the EUR vs. the CAD and the SEK; and the CHF vs. the CAD and the SEK. The results for the JPY vs. the GBP, the CAD, and the AUD are particularly strong and also allow rejection of the hypothesis that $\phi_1 = 1$.

These results are completely consistent with the literature on the FX carry trade, which is a strategy that borrows low interest rate currencies and lends high interest rate currencies. The dependent variable is the return to the carry trade when $r_{j,t} > r_{i,t}$, and the highly positive values of the slope coefficients indicate that expected carry trade profits are conditionally high when $r_{j,t} - r_{i,t}$ is conditionally high.[12]

[11] Once again, for completeness we present the results for the NOK, but because the NOK sample ending in 2003 is particularly short, we do not interpret them.

[12] See Daniel, Hodrick and Lu (2017) for a recent review of the literature on the risks of the carry trade at the monthly holding period horizon. Lustig, Stathopoulos and Verdelhan (2017) find that investing in the carry trade with longer term bonds while maintaining the one-month holding period is unattractive as the term premiums offset the currency premiums.

Now, consider the results from estimating equation (25) for the sample in which the dependent variables start in 2004:12 and end in 2016:12 in Table 10. These results are presented in the same format as Table 9. The estimates of ϕ_1 in Panel A with the USD as the base currency are now uniformly negative (except for the CHF) and are generally as large in absolute value as the positive values from Table 9. The standard errors are larger though, and only the test statistics for the EUR and the CAD are sufficiently large to allow rejection of the hypothesis that $\phi_1 = 0$.

Only 9 of the 28 estimates in Panels B through H are positive, and only 4 test statistics indicate rejection of the hypothesis that $\phi_1 = 0$ at the 0.1 marginal level of significance. Clearly, the conclusion should be that the later sample containing the financial crisis is quite different from the pre-2004 sample. These results are also consistent with the post-2007 deterioration in the returns to the carry trade.

6 Out-of-Sample Results

The previous sections examined the predictability of excess returns in bond and foreign exchange markets with classical asymptotic distribution theory. Inoue and Kilian (2005) argue that such an approach is actually more powerful than out-of-sample experiments, yet such experiments are routinely done and are considered to be a good indicator of instability in the underlying forecasting model. This section consequently examines whether the Cochrane and Piazzesi (2005, 2008) model can forecast the excess rates of return in bond and foreign currency markets out of sample. As above, we use the sample period that would have been available when the original paper was written as the in-sample period and treat post-2003 beginning in January 2004 through December 2016 as the out-of-sample period.

We follow Welch and Goyal (2007) and Campbell and Thompson (2007) in assessing the models' out-of-sample forecasts by examining two statistics. The first is the R^2 that compares the mean squared error of the conditional forecasts of excess returns from the term structure model to the mean squared error from assuming that the conditional forecasts of the excess returns are the conditional sample means using data up to that point in time. Analytically, if $\hat{r}_t$ represents the $t-$th out-of-sample forecast from the Cochrane and Piazzesi (2005, 2008) model using parameters estimated with all the historical data available at that time, and if $\overline{r}_t$ represents the analogous forecast from the historical sample mean, using the same sample period, then with T_{os} total out-of-sample observations, the mean squared error from the CP forecasts is

$$MSE_{CP} = \frac{1}{T_{os}} \sum_{t=1}^{T_{os}} \left(r_t - \hat{r}_t \right)^2 , \tag{27}$$

and the mean squared error from the historical mean forecasts is

$$MSE_{HM} = \frac{1}{T_{os}} \sum_{t=1}^{T_{os}} \left(r_t - \overline{r}_t \right)^2 . \tag{28}$$

The R^2 is then defined as

$$R^2 = 1 - \frac{MSE_{CP}}{MSE_{HM}} \tag{29}$$

The second closely related statistic is the Clark and McCracken (2005) $MSE-F$ which tests for the equality of the two forecasts:

$$MSE - F = T_{os} \frac{MSE_{HM} - MSE_{CP}}{MSE_{CP}}. \tag{30}$$

Table 11 presents the out-of-sample forecasts of excess bond returns for the Cochrane and Piazzesi (2005) model estimated separately for each currency. The first forecast is 2004:01, and the last is 2016:12.

The results are quite mixed. The model's forecasts are worse than the forecast based on the historical mean at all maturities for the USD, the EUR, the JPY, the AUD, and the NOK. Only for the GBP do the model forecasts beat historical mean forecast for all maturities. For the CHF, the CAD, and the SEK, the results are mixed across maturities.

6.1 An Alternative Model with Free Constants

In discussing the relation of the Cochrane and Piazzesi (2008) ATSM to the empirical model in Cochrane and Piazzesi (2005), we noted that the former does not constrain the constant terms to have the same factor of proportionality across maturities as is imposed by the latter. To see whether relaxing this constraint which formally nests the historical mean model as a constrained version of the larger model, we recursively estimated the model with free constant terms for each maturity.

The results of the out-of-sample forecasts are presented in Table 12. All of the R^2's except for maturities 3, 4, and 5 for the CHF are negative. The forward spreads apparently provide no useful out-of-sample forecasting power for the excess bond returns.

6.2 Constraining Parameters Across Currencies

In out-of-sample forecasting situations, it is often advised to limit the number of free parameters that are estimated. We experimented with this Occam's razor intuition and recursively estimated a model that constrains the slope coefficients in the forecasting equations to be the same across countries while freeing the constants of the model from the constraint that they are proportional to the slope coefficients. These out-of-sample forecasting results show that all of the model forecasts except for three maturities of the CHF are worse than the historical mean, and these results are consequently presented in the Online Appendix.

6.3 Out-of-Sample Forecasts of Currency Returns

Given the inability of the CP forecasting factors and their squared values to forecast excess rates of return in currency markets in in-sample regressions, the reader should expect that the model will not be useful in out-of-sample experiments. For completeness, we present these results in Table 13. The results are indeed as anticipated as the out-of-sample forecasts from the model are unable to beat the historical mean excess returns of all currencies versus the USD.

6.4 Evolution of the USD Parameters

The failure of the model in the out-of-sample forecasting experiments and the rejection of equality of coefficients across sub-periods suggests substantial parameter instability. While a full analysis of this issue is not something we have space to accomplish in this article, Figure 1 presents the recursive estimates of the parameters of the Cochrane and Piazzesi (2005) model for the USD term structure as they evolve in the out-of-sample estimation period.[13] The estimates of the b_n parameters remain incredibly stable as the four lines are virtually horizontal. It is also clear that beginning in 2008 with the advent of the financial crisis, the estimated $\gamma(2)$ changes over the course of two years from positive to negative, the estimated $\gamma(3)$ begins a slow decline, and the estimated $\gamma(5)$ experiences a steady increase. The estimated $\gamma(4)$ is reasonably constant after a blip in 2009. Because these are recursive estimates that use all of the sample to that point in time, they are more stable than would be recursive rolling estimates that use the same sample size at each point in time. In that sense, the slow evolution masks more dramatic changes.

7 Conclusions

In this paper we document substantive instabilities in the empirical analysis of risk premiums in bond and foreign exchange markets. One puzzle appears to be the observation that there is a strong one-factor structure to the forecasts of expected returns in the bond markets in a particular sample of data, but a different one-factor structure in another sample. Modeling the sources of the structural changes should be high on the research agenda. It is also puzzling that the CP factors in two currencies have such strong predictability in their respective bond markets but not in the foreign exchange market between the two currencies.

There are many directions that research on time varying risk premiums in bond and foreign exchange markets could go. Here we review some recent approaches.

In extending the Cochrane and Piazzesi (2005) model to additional currencies and considering its international implications, we have not addressed the term structure literature arguing that macroeconomic variables, such as inflation and employment, have additional forecasting power over and above that available in bond yields. In this regard we cite two recent critiques of this literature. First, Ghysels, Horan and Moench (2018) find that several studies touting the significantly improved forecasting performance of macroeconomic variables above that provided by yields overstate their importance because the studies use revised data. Ghysels, Horan and Moench (2018) find that use of real time U.S. data substantially reduces the implied predictive power. Second, Bauer and Hamilton (2017) argue that after taking account of small sample distortions in the test statistics induced by the use of macro variables with trends, the evidence for additional predictability from macro variables is much weaker. Because addressing these issues in our multiple currency context is beyond what can be accomplished in a given article, we leave these issues to future research.

We have also focused exclusively on the annual forecasting horizon. Most bond market ATSMs are estimated at the monthly horizon and typically find that monthly risk premiums are driven by more than one state variable. In contrast, we find the strong one-factor structure originally documented by Cochrane and Piazzesi (2005) at the annual horizon. Examining the dynamics of the state variables in monthly models and seeing whether they imply a single state variable at the annual horizon would be an interesting project. Recent papers that examine multiple horizons include Bacchetta and Van Wincoop (2010), Engel (2016), Lustig, Stathopoulos and Verdelhan (2017), and Chernov and Creal (2018) who find interesting patterns in expected returns at different horizons.

While ATSMs are typically developed under the assumption of rational expectations, it may be the

[13] Figures for the other currencies are available in the Online Appendix.

case that behavioral finance with its time varying sentiments could be responsible for our findings of model instability. One recent empirical analysis with a behavioral slant is Brooks and Moskowitz (2017), who examine quarterly returns on bonds from Australia, Germany, Canada, Japan, Sweden, the UK, and the U.S. Using panel data methods with time fixed effects, they argue that measures of value, carry, and momentum dominate the CP factor in forecasting excess returns. While we are forecasting actual excess returns on a currency by currency basis, the panel data approach of Brooks and Moskowitz (2017) implies that they are forecasting deviations from cross-sectional average returns.

Another alternative approach to these issues would rely on the analysis of Krishnamurthy and Vissing-Jorgensen (2011) who argue that the supply of U.S. Treasury securities affects the level and slope of the yield curve. Do such changes in quantities also affect the risk premiums in the other bond markets and in the foreign exchange markets? Valchev (2017) answers this question affirmatively. It is natural to think that major changes in monetary and fiscal policies, including the quantitative easing done by major central banks during the international financial crisis, could induce the changes in the parameters of the CP factors and the resulting changes in forecasting power that we observe. Actually demonstrating this empirically is a challenging task. Evidence of substantive structural change in the international financial markets can also be found in the deviations from covered interest rate parity documented by Du, Tepper and Verdelhan (2018) and Rime, Schrimpf and Syrstad (2017).[14]

Rather than focusing on time varying risk premiums, Valchev (2017) and Jiang, Krishnamurthy and Lustig (2018) are two recent papers that empirically explore differential time-varying liquidity premiums, or non-pecuniary returns, on government bonds as explanations of rates of currency depreciation. Can these models explain the instabilities that we document?

Our econometric analysis also is conducted under the standard assumption that investors have rational expectations and that the data are stationary and ergodic. It has long been recognized that changes in monetary policy regimes can cause problems with econometric analysis of the term structure. Fuhrer (1996) argues that investors are aware of changes in regimes but do not anticipate future changes, which he views as a compromise between full rationality and learning. Bekaert, Hodrick and Marshall (2001) argue that so-called peso problems, caused by differences between the frequency of realizations of the data and the conditional distributions investors had at the time that they set bond prices, could be responsible for the anomalies observed in the term structure literature.

The necessity for investors to learn about changes in monetary policy, the rate of inflation, or the real interest rate are also important areas of recent research that relaxes the rational expectations assumption. Piazzesi, Salomao and Schneider (2015) note that professional forecasts of interest rates differ from those based on regressions. They build on the insights of Froot (1989) who argued that evidence against the expectations hypothesis of the term structure was plausibly due to the failure of the rational expectations assumption imposed in the tests rather than to failures of the expectations hypothesis itself.

Giacoletti, Laursen and Singleton (2016) argue that marginal investors in the bond market act as Bayesian learners to form prospective real-time views about bond market risks. While the sources of risks are the first three principal components of the yield curve, knowledge of the extent of disagreement among professionals is informative about how today's yield curve will impact its future shape and thus the prices of risks.

The studies cited here provide some interesting directions in which research can go. Most of these papers

[14] Andersen, Duffie and Song (2018) provide a theoretical explanation for deviations from covered interest rate parity in a world with highly levered, risky financial market makers.

do not investigate time variation in the parameters of their empirical models. Our paper provides a set of challenging empirical results demonstrating more attention should be devoted to this type of analysis. The paper also provides interesting empirical evidence showing an absence of links that should theoretically be present between the term structures of interest rates in two currencies and the currency market between them.

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A Data

Data on the term structures of interest rates for the different currencies were obtained from several sources. The USD data are from the CRSP Fama-Bliss database. This is the same source as Cochrane and Piazzesi (2005). For yields from the non-USD term structures, we obtained data from Jonathan Wright's web site. These data were used in Wright (2011). We updated the data from the web sites of the respective central banks. The monthly term structure data all end in December 2016. The data begin in June 1952 for the USD, in January 1970 for the GBP, in January 1973 for the EUR spliced with the Deutsche mark prior to 1999; in March 1988 for the CHF; in January 1986 for the CAD; in January 1985 for the JPY; in February 1987 for the AUD; in January 1987 for the SEK; and in January 1998 for the NOK. The exchange rate data are from the International Monetary Fund, and the sample period is January 1973 to December 2016 for all exchange rates.

B The Affine Model Solutions

The solutions to the coefficients of the natural logarithms of the bond prices in the affine model given in equation (9) are the following difference equations:

$$\begin{aligned} A_n &= A_{n-1} - \delta_0 + \boldsymbol{B}_{n-1}^{\intercal} \left(\boldsymbol{\mu} - \boldsymbol{\Sigma}\boldsymbol{\lambda}_0\right) + (1/2)\, \boldsymbol{B}_{n-1}^{\intercal} \boldsymbol{\Sigma}\boldsymbol{\Sigma}^{\intercal} \boldsymbol{B}_{n-1} && \text{(B.1)} \\ \boldsymbol{B}_n^{\intercal} &= -\boldsymbol{\delta}_1^{\intercal} + \boldsymbol{B}_{n-1}^{\intercal} \left(\boldsymbol{\Phi} - \boldsymbol{\Sigma}\boldsymbol{\lambda}_1\right) && \text{(B.2)} \end{aligned}$$

with initial conditions $A_0 = 0$ and $\boldsymbol{B}_0 = \mathbf{0}$. By defining $\boldsymbol{\Phi}^* = (\boldsymbol{\Phi} - \boldsymbol{\Sigma}\boldsymbol{\lambda}_1)$, we can write equation (B.2) as

$$\boldsymbol{B}_n^{\intercal} = -\boldsymbol{\delta}_1^{\intercal} \left(\mathbf{I} - \boldsymbol{\Phi}^*\right)^{-1} \left(\mathbf{I} - \boldsymbol{\Phi}^{*n}\right) \tag{B.3}$$

C The Standard Errors

This appendix derives the standard errors for the two-step estimation of the term structure models that generate the CP forecasting factors and the corresponding forecasting equation for the excess return on base currency i relative to currency j. Let $\bar{\varepsilon}_{i,t+1}$ and $\bar{\varepsilon}_{j,t+1}$ be the error terms in equation (5) for the term structure regressions associated with the currencies i and j, respectively. The error term, $\varepsilon_{j,t+1}^{s}$, is defined in equation (24). Let $\mathbf{h}_{ij,t} \equiv \left[1,\ x_{i,t},\ x_{i,t}^2,\ x_{j,t},\ x_{j,t}^2\right]^{\intercal}$, where $x_{j,t} = \widehat{\boldsymbol{\gamma}}_j^{\intercal} \overline{\mathbf{fs}}_{j,t}$ is the return forecasting variable from the estimation of equation (5) for currency j, and let $\boldsymbol{\psi}$ represent the vector of parameters in equation (24). Then, the orthogonality conditions associated with the forecasts of the average excess returns in the two bond markets and the excess rate of return in the currency market are the following:

$$\begin{aligned} E\left[\overline{\mathbf{fs}}_{1,t} \times \bar{\varepsilon}_{1,t+1}\right] &= \mathbf{0} \\ E\left[\overline{\mathbf{fs}}_{j,t} \times \bar{\varepsilon}_{j,t+1}\right] &= \mathbf{0} \\ E\left[\mathbf{h}_{ij,t} \times \varepsilon_{j,t+1}^{s}\right] &= \mathbf{0}. \end{aligned} \tag{C.1}$$

The parameter vector is $\boldsymbol{\theta} = \left[\boldsymbol{\gamma}_1^{\intercal},\ \boldsymbol{\gamma}_j^{\intercal},\ \boldsymbol{\psi}^{\intercal}\right]^{\intercal}$. Let $\mathbf{g}_T(\boldsymbol{\theta})$ denote the sample mean of the orthogonality conditions in the system of equations given in (C.1). Because the system is just identified, these sample orthogonality conditions can be set to zero, and the asymptotic variance of the parameter estimates can be estimated as

$$\mathbf{V}(\boldsymbol{\theta}) = \frac{1}{T} \mathbf{D}_T^{-1} \mathbf{S}_T \mathbf{D}_T^{-1\intercal} \tag{C.2}$$

where

$$\mathbf{D}_T = \frac{\partial \mathbf{g}_T(\boldsymbol{\theta})}{\partial \boldsymbol{\theta}^{\intercal}} \tag{C.3}$$

is the sample estimate of the Jacobian of the orthogonality conditions, $\mathbf{D}$, which is defined below, and

$$\mathbf{S}_{\mathbf{T}} \equiv \mathbf{C}_0 + \sum_{k=1}^{K} \frac{K-k}{K} \left(\mathbf{C}_k + \mathbf{C}_k^{\intercal}\right), \tag{C.4}$$

is the sample estimate of the variance of the orthogonality conditions. The autocovariances are estimated with

$$\mathbf{C}_k \equiv \frac{1}{T} \sum_{t=k+1}^{T} \mathbf{g}_t \mathbf{g}_{t-k}^{\intercal} \tag{C.5}$$

where $\mathbf{g}_t$ is the vector of observations on the orthogonality conditions are time t, and we use $K = 18$.

The derivatives in equation (C.3) are sample estimates of

$$\mathbf{D} = \begin{bmatrix} -E\left(\overline{\mathbf{fs}}_{1,t}\overline{\mathbf{fs}}_{1,t}^{\intercal}\right) & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & -E\left(\overline{\mathbf{fs}}_{j,t}\overline{\mathbf{fs}}_{j,t}^{\intercal}\right) & \mathbf{0} \\ \mathbf{D}_1 & \mathbf{D}_2 & \mathbf{D}_3 \end{bmatrix}$$

where $\mathbf{D}_1 \equiv \nabla_{\boldsymbol{\gamma}_1^{\intercal}} E\left(\varepsilon_{j,t+1}^{s}\mathbf{h}_{ij,t}\right)$, $\mathbf{D}_2 \equiv \nabla_{\boldsymbol{\gamma}_j^{\intercal}} E\left(\varepsilon_{j,t+1}^{s}\mathbf{h}_{ij,t}\right)$, and $\mathbf{D}_3 \equiv \nabla_{\psi^{\intercal}} E\left(\varepsilon_{j,t+1}^{s}\mathbf{h}_{ij,t}\right)$, respectively. We estimate $\mathbf{D}_T$ using the jacobianest function from the Matlab DERIVEST suite of D'Ericco (2011).

From the structure of the $\mathbf{D}$ matrix and the partitioned inverse formula, one sees that the variances of the estimates of $\boldsymbol{\gamma}_1$ and $\boldsymbol{\gamma}_j$ are not affected by the estimation of $\boldsymbol{\psi}$ whereas the variances of the latter parameters are affected by the estimation of the former.

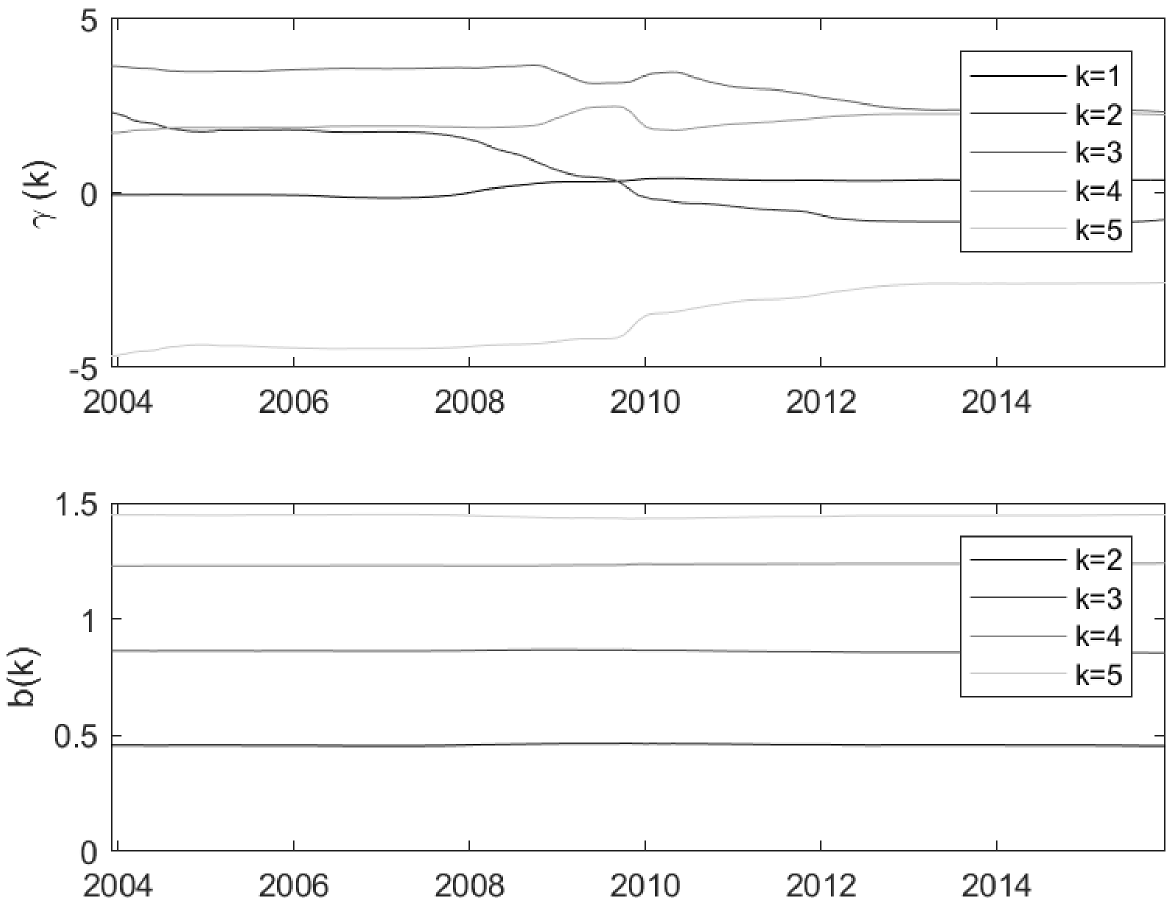

Figure 1: Evolution of parameters (USD)

Table 1: The Single Factor Model (Pre-2004 Data)

The Table reports coefficient estimates for the two-step estimation of the constrained single factor model. The first step involves OLS estimation of the average one-year excess rates of return on bonds with two through five years to maturity, $\overline{rx}_{t+1}$, on a constant and the average of the current value and two monthly lags of the four forward spreads, $\overline{\mathbf{fs}}_t$:

$$\overline{rx}_{t+1} = \gamma^{\intercal}\overline{\mathbf{fs}}_t + \overline{\varepsilon}_{t+1}.$$

The second step involves OLS regressions of the individual excess rates of return on bonds with two through four years to maturity on the fitted value from the first step:

$$rx_{t+1}^{(n)} = b_n\left(\widehat{\gamma}^{\intercal}\overline{\mathbf{fs}}_t\right) + \varepsilon_{t+1}^{(n)}$$

Standard errors in the first step are based on the usual GMM versions from OLS orthogonality conditions, and the standard errors in the second step allow for the estimation error in the first step. All standard errors are constructed with 18 Newey-West (1987) lags and are in parentheses. The $\chi^2(4)$ statistic tests the hypothesis that γ_2 through γ_5 equal zero with p-values in angled brackets. The R^2 is from the first step regression. The column labeled $\%PC1$ presents the percentage of the variance of the unconstrained estimates of the four excess rates of return explained by their first principal component. The sample periods for the dependent variables all end in 2003:12. The samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK.

CUR	b_2	b_3	b_4	γ_1	γ_2	γ_3	γ_4	γ_5	$\chi^2(4)$	R^2	%PC1
USD	0.46	0.86	1.23	-0.04	2.32	3.65	1.73	-4.67	100.04	0.39	0.997
	(0.02)	(0.02)	(0.01)	(0.63)	(1.63)	(1.33)	(0.95)	(0.58)	⟨0.00⟩		
GBP	0.44	0.86	1.21	1.04	-5.40	11.24	-6.54	0.39	9.46	0.09	0.996
	(0.05)	(0.03)	(0.02)	(0.53)	(4.31)	(11.72)	(14.23)	(5.97)	⟨0.05⟩		
EUR	0.44	0.86	1.21	1.02	-1.77	-4.19	16.47	-10.50	9.66	0.12	0.992
	(0.03)	(0.03)	(0.01)	(0.90)	(5.07)	(14.33)	(19.38)	(9.17)	⟨0.05⟩		
CHF	0.47	0.86	1.19	0.78	-1.45	-9.62	26.03	-14.60	22.25	0.24	0.994
	(0.08)	(0.04)	(0.03)	(1.45)	(11.88)	(30.09)	(38.64)	(18.01)	⟨0.00⟩		
CAD	0.46	0.86	1.20	1.64	0.53	-5.16	19.68	-13.59	34.30	0.23	0.993
	(0.03)	(0.03)	(0.01)	(0.84)	(2.34)	(5.25)	(6.95)	(4.03)	⟨0.00⟩		
JPY	0.37	0.81	1.23	1.00	-14.99	17.34	-0.26	-5.02	26.00	0.25	0.997
	(0.02)	(0.02)	(0.00)	(1.23)	(3.75)	(4.15)	(2.07)	(1.88)	⟨0.00⟩		
AUD	0.44	0.87	1.21	0.82	-32.92	107.79	-138.85	61.21	15.97	0.11	0.996
	(0.02)	(0.02)	(0.01)	(1.07)	(12.15)	(42.03)	(61.52)	(29.98)	⟨0.00⟩		
SEK	0.40	0.80	1.20	1.11	25.03	-52.71	54.15	-21.11	4.60	0.09	0.998
	(0.05)	(0.03)	(0.02)	(1.44)	(13.90)	(39.27)	(51.85)	(24.07)	⟨0.33⟩		
NOK	0.44	0.86	1.21	1.99	33.21	-128.76	202.25	-102.78	116.53	0.42	0.988
	(0.03)	(0.02)	(0.01)	(1.24)	(19.50)	(70.19)	(99.33)	(47.02)	⟨0.00⟩		

Table 2: The Single Factor Model (Post-2003 Data)

The Table reports coefficient estimates for the two-step estimation of the constrained single factor model. The first step involves OLS estimation of the average one-year excess rates of return on bonds with two through five years to maturity, $\overline{rx}_{t+1}$, on a constant and the average of the current value and two monthly lags of the four forward spreads, $\overline{\mathbf{fs}}_t$:

$$\overline{rx}_{t+1} = \gamma^{\intercal}\overline{\mathbf{fs}}_t + \bar{\varepsilon}_{t+1}.$$

The second step involves OLS regressions of the individual excess rates of return on bonds with two through four years to maturity on the fitted value from the first step:

$$rx_{t+1}^{(n)} = b_n \left(\widehat{\gamma}^{\intercal}\overline{\mathbf{fs}}_t\right) + \varepsilon_{t+1}^{(n)}$$

Standard errors in the first step are based on the usual GMM versions from OLS orthogonality conditions, and the standard errors in the second step allow for the estimation error in the first step. All standard errors are constructed with 18 Newey-West (1987) lags and are in parentheses. The $\chi^2(4)$ statistic tests the hypothesis that γ_2 through γ_5 equal zero with p-values in angled brackets. The R^2 is from the first step regression. The column labeled $\%PC1$ presents the percentage of the variance of the unconstrained estimates of the four excess rates of return explained by their first principal component. The sample periods for the dependent variables for all currencies begin in 2004:12 and end in 2016:12.

CUR	b_2	b_3	b_4	γ_1	γ_2	γ_3	γ_4	γ_5	$\chi^2(4)$	R^2	%PC1
USD	0.35	0.78	1.24	0.07	-7.57	2.17	2.56	-0.80	25.20	0.37	0.974
	(0.06)	(0.06)	(0.01)	(0.69)	(2.62)	(1.75)	(0.60)	(0.61)	⟨0.00⟩		
GBP	0.37	0.82	1.23	0.17	-8.33	-5.16	26.84	-15.82	64.67	0.29	0.963
	(0.08)	(0.06)	(0.02)	(0.76)	(7.58)	(19.98)	(24.83)	(10.47)	⟨0.00⟩		
EUR	0.34	0.77	1.22	1.61	-21.00	35.22	-24.91	6.39	6.46	0.08	0.972
	(0.06)	(0.05)	(0.01)	(1.49)	(14.68)	(33.30)	(35.71)	(14.38)	⟨0.17⟩		
CHF	0.28	0.74	1.25	0.12	-10.41	16.56	-14.21	5.73	166.37	0.39	0.983
	(0.06)	(0.06)	(0.01)	(0.47)	(2.88)	(8.60)	(11.62)	(5.46)	⟨0.00⟩		
CAD	0.38	0.82	1.22	1.60	4.22	-14.46	22.67	-11.20	19.45	0.17	0.949
	(0.08)	(0.06)	(0.02)	(1.04)	(3.61)	(6.34)	(6.34)	(2.70)	⟨0.00⟩		
JPY	0.30	0.73	1.29	0.40	3.33	-1.00	4.03	-3.35	37.00	0.34	0.994
	(0.03)	(0.04)	(0.02)	(0.22)	(1.73)	(2.69)	(1.72)	(1.13)	⟨0.00⟩		
AUD	0.33	0.81	1.22	1.50	3.06	2.67	-4.15	0.90	17.52	0.12	0.970
	(0.11)	(0.08)	(0.02)	(1.00)	(2.61)	(2.08)	(1.46)	(1.11)	⟨0.00⟩		
SEK	0.37	0.80	1.22	1.90	15.17	-41.67	49.77	-21.08	11.37	0.13	0.973
	(0.08)	(0.07)	(0.02)	(1.08)	(8.55)	(17.71)	(19.88)	(8.85)	⟨0.02⟩		
NOK	0.36	0.79	1.22	0.97	-4.14	-1.65	5.95	-2.72	2.84	0.08	0.866
	(0.14)	(0.13)	(0.02)	(0.98)	(10.55)	(22.99)	(22.88)	(8.73)	⟨0.59⟩		

Table 3: Tests of Equality of Coefficients

The Table reports two statistics that test the equality of the coefficient estimates in the Cochrane and Piazzesi (2005) models for the two samples estimated in Tables 1 and 2. The first test examines the **b** coefficients and is distributed as a $\chi^2(3)$. The second test examines the $\boldsymbol{\gamma}$ coefficients and is distributed as a $\chi^2(5)$. The *p*-values are in angled brackets. The first sample periods for the dependent variables all end in 2003:12. These samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The second sample period for the dependent variables is 2004:12 to 2016:12 for all currencies.

CUR	$\chi^2(3)$ for **b**'s $\langle$ p-val $\rangle$	$\chi^2(5)$ for $\boldsymbol{\gamma}$'s $\langle$ p-val $\rangle$
USD	5.15	81.76
	$\langle$ 0.16 $\rangle$	$\langle$ 0.00 $\rangle$
GBP	4.77	9.52
	$\langle$ 0.19 $\rangle$	$\langle$ 0.09 $\rangle$
EUR	11.64	4.39
	$\langle$ 0.01 $\rangle$	$\langle$ 0.49 $\rangle$
CHF	8.69	5.47
	$\langle$ 0.03 $\rangle$	$\langle$ 0.36 $\rangle$
CAD	5.46	7.16
	$\langle$ 0.14 $\rangle$	$\langle$ 0.21 $\rangle$
JPY	26.58	25.98
	$\langle$ 0.00 $\rangle$	$\langle$ 0.00 $\rangle$
AUD	2.87	9.37
	$\langle$ 0.41 $\rangle$	$\langle$ 0.10 $\rangle$
SEK	1.55	1.33
	$\langle$ 0.67 $\rangle$	$\langle$ 0.93 $\rangle$
NOK	3.10	34.09
	$\langle$ 0.38 $\rangle$	$\langle$ 0.00 $\rangle$

Table 4: Correlation Matrix and Variance Decomposition of Country CP Factors (Pre-2004 Data)

The Table presents the correlation matrix of the CP factors, the fitted return forecasting variables from the term structure regressions for the different currencies in Table 1. Because the samples are different lengths, the correlations are estimated over the shorter of the two samples. The sample periods for the dependent variables all end in 2003:12. The samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The last column labelled %PC reports the percent of variance explained by the i-th principal component.

	Correlations									%PC	
	USD	GBP	EUR	CHF	CAD	JPY	AUD	SEK	NOK	i	%PC(i)
USD	1.00	0.06	0.23	0.26	0.37	0.22	0.35	0.11	-0.02	1	0.40
GBP	0.06	1.00	-0.19	0.63	0.49	0.45	0.10	-0.22	0.27	2	0.29
EUR	0.23	-0.19	1.00	0.12	-0.11	0.18	-0.08	0.46	0.44	3	0.13
CHF	0.26	0.63	0.12	1.00	0.19	0.34	-0.16	-0.17	0.34	4	0.09
CAD	0.37	0.49	-0.11	0.19	1.00	0.24	0.12	0.09	-0.08	5	0.04
JPY	0.22	0.45	0.18	0.34	0.24	1.00	0.02	0.02	-0.30	6	0.03
AUD	0.35	0.10	-0.08	-0.16	0.12	0.02	1.00	-0.07	0.18	7	0.01
SEK	0.11	-0.22	0.46	-0.17	0.09	0.02	-0.07	1.00	0.39	8	0.01
NOK	-0.02	0.27	0.44	0.34	-0.08	-0.30	0.18	0.39	1.00	9	0.00

Table 5: Correlation Matrix and Variance Decomposition of Country CP Factors (Post-2003 Data)

The Table presents the correlation matrix of the CP factors, the fitted return forecasting variables from the term structure regressions for the different currencies in Table 2. The sample period for the dependent variables is 2004:12 to 2016:12. The last column labelled %PC reports the percent of variance explained by the i-th principal component.

	Correlations									%PC	
	USD	GBP	EUR	CHF	CAD	JPY	AUD	SEK	NOK	i	%PC(i)
USD	1.00	0.42	-0.07	0.12	0.34	0.35	0.12	0.08	0.15	1	0.37
GBP	0.42	1.00	0.39	0.34	0.40	-0.11	0.08	0.16	0.54	2	0.18
EUR	-0.07	0.39	1.00	0.33	0.19	-0.32	0.08	0.13	0.35	3	0.14
CHF	0.12	0.34	0.33	1.00	0.17	0.08	-0.01	0.18	0.44	4	0.10
CAD	0.34	0.40	0.19	0.17	1.00	0.37	0.16	0.11	0.04	5	0.08
JPY	0.35	-0.11	-0.32	0.08	0.37	1.00	0.06	0.01	-0.29	6	0.06
AUD	0.12	0.08	0.08	-0.01	0.16	0.06	1.00	0.15	-0.30	7	0.04
SEK	0.08	0.16	0.13	0.18	0.11	0.01	0.15	1.00	0.20	8	0.02
NOK	0.15	0.54	0.35	0.44	0.04	-0.29	-0.30	0.20	1.00	9	0.01

Table 6: Explaining Currency Market Excess Returns with Changes in Level Factors

The Table presents estimation results for the regression

$$-\Delta s_{1j,t+1} + r_{j,t} - r_{1,t} = \beta_0 + \beta_1 \Delta l_{1,t+1} + \beta_2 \Delta l_{1,t+1} \times x_{1,t} + \beta_3 \Delta l_{j,t+1} + \beta_4 \Delta l_{j,t+1} \times x_{j,t} + \varepsilon_{1j,t+1}$$

where the dependent variable is the excess rate of return in USD on an annual investment in the money market of currency j, which is our proxy for the innovation in the rate of dollar appreciation. The regressors are the contemporaneous changes in the first principal components of the yields for the USD, $\Delta l_{1,t+1}$, and currency j, $\Delta l_{j,t+1}$, and the interaction of these variables with their respective currency specific CP factors, which are the term structure excess return forecasting variables, $x_{1,t}$ and $x_{j,t}$. Standard errors are in parentheses. The sample periods for the dependent variables all end in 2016:12. The samples begin in 1974:12 for the GBP and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK.

CUR	β_0	β_1	β_2	β_3	β_4	R^2
GBP	0.60	67.19	-17.42	-69.35	48.99	0.05
	(1.45)	(48.22)	(16.73)	(58.53)	(22.87)	
EUR	-0.43	-53.75	-5.55	42.14	-42.59	0.04
	(1.71)	(52.77)	(23.15)	(79.43)	(43.78)	
CHF	0.05	-5.23	-27.99	-29.78	1.04	0.02
	(2.01)	(102.34)	(62.41)	(75.36)	(73.78)	
CAD	1.59	24.99	0.56	12.57	14.58	0.02
	(1.42)	(79.51)	(24.25)	(69.07)	(20.94)	
JPY	-0.95	159.31	-79.08	-302.53	83.38	0.23
	(2.14)	(158.57)	(44.50)	(116.41)	(50.06)	
AUD	3.92	21.78	-27.40	98.66	36.56	0.13
	(1.97)	(99.81)	(42.01)	(187.34)	(99.49)	
SEK	0.60	30.17	-61.84	258.79	-65.46	0.20
	(1.78)	(101.23)	(51.72)	(72.19)	(43.40)	
NOK	0.98	-7.12	-110.05	128.38	-26.44	0.10
	(2.51)	(204.10)	(113.57)	(309.66)	(308.21)	

Table 7: Forecasts of Excess Rates of Return in Currency Markets (Pre-2004 Data)

The Table presents estimation results for the regression

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \psi_0 + \psi_1 x_{i,t} + \psi_2 x_{i,t}^2 + \psi_3 x_{j,t} + \psi_4 x_{j,t}^2 + \epsilon^s_{ij,t+1}$$

where the dependent variable is the excess rate of return in base currency i on an annual investment in the money market of currency j. The regressors are the CP factors, the fitted return forecasting variables from the term structure regressions for currencies i and j, and their squared values. Standard errors are in parentheses, and p-values are in angled brackets. The $\chi^2(2)_i$ statistic tests the null hypothesis that ψ_1 and ψ_2 equal zero, the $\chi^2(2)_j$ statistic tests the null hypothesis that ψ_3 and ψ_4 equal zero, and the $\chi^2(4)$ statistic tests the null hypothesis that ψ_1 through ψ_4 equal zero. The sample periods for the dependent variables all end in 2003:12. The samples begin in 1974:12 for the GBP and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK.

CUR j	ψ_0	ψ_1	ψ_2	ψ_3	ψ_4	$\chi^2(2)_i$	$\chi^2(2)_j$	$\chi^2(4)$	R^2
Panel A: i = USD									
GBP	1.63	0.48	-0.01	0.73	-0.76	0.55	1.12	1.37	0.04
	(2.56)	(0.67)	(0.13)	(1.74)	(0.73)	⟨0.76⟩	⟨0.57⟩	⟨0.85⟩	
EUR	-0.57	1.44	-0.01	-2.15	0.56	4.79	1.10	5.07	0.08
	(3.16)	(0.66)	(0.11)	(3.00)	(0.59)	⟨0.09⟩	⟨0.58⟩	⟨0.28⟩	
CHF	-2.12	-2.80	1.08	-1.83	0.63	2.15	0.86	3.76	0.13
	(5.41)	(4.48)	(0.97)	(2.18)	(1.04)	⟨0.34⟩	⟨0.65⟩	⟨0.44⟩	
CAD	3.60	-1.13	0.40	-1.13	-0.17	1.13	5.22	7.26	0.19
	(2.44)	(2.20)	(0.47)	(0.85)	(0.26)	⟨0.57⟩	⟨0.07⟩	⟨0.12⟩	
JPY	-2.25	1.21	0.39	1.77	-1.11	11.32	6.96	22.22	0.38
	(5.79)	(3.23)	(0.36)	(3.26)	(0.81)	⟨0.00⟩	⟨0.03⟩	⟨0.00⟩	
AUD	6.71	-3.16	0.80	-3.45	0.52	0.72	1.75	3.80	0.05
	(8.69)	(6.03)	(1.19)	(2.70)	(0.55)	⟨0.70⟩	⟨0.42⟩	⟨0.43⟩	
SEK	-0.16	-1.69	0.33	1.39	-0.10	0.11	0.95	1.51	-0.00
	(4.78)	(5.11)	(1.02)	(2.27)	(0.59)	⟨0.95⟩	⟨0.62⟩	⟨0.82⟩	
NOK	-0.23	-0.81	1.23	0.33	-0.26	1.58	1.85	4.61	0.36
	(6.03)	(9.06)	(1.97)	(1.29)	(0.25)	⟨0.45⟩	⟨0.40⟩	⟨0.33⟩	
Panel B: i = GBP									
EUR	-2.65	0.77	0.33	-1.96	0.78	2.31	3.70	12.21	0.11
	(3.10)	(1.05)	(0.49)	(2.62)	(0.47)	⟨0.32⟩	⟨0.16⟩	⟨0.02⟩	
CHF	-4.35	0.95	1.96	-2.18	-1.01	1.66	2.28	3.00	0.43
	(5.05)	(6.42)	(2.63)	(2.03)	(0.90)	⟨0.44⟩	⟨0.32⟩	⟨0.56⟩	
CAD	-1.89	0.86	-0.15	1.68	-0.57	0.10	0.93	2.17	0.04
	(3.41)	(4.01)	(2.24)	(1.82)	(0.62)	⟨0.95⟩	⟨0.63⟩	⟨0.70⟩	
JPY	-7.10	2.51	0.72	4.34	-1.55	0.96	3.51	4.79	0.09
	(9.44)	(5.64)	(3.99)	(3.07)	(0.83)	⟨0.62⟩	⟨0.17⟩	⟨0.31⟩	
AUD	4.18	0.54	0.11	-3.74	0.37	0.27	2.48	2.70	0.06
	(4.64)	(4.45)	(1.99)	(3.50)	(0.63)	⟨0.87⟩	⟨0.29⟩	⟨0.61⟩	
SEK	-7.58	1.46	0.42	2.50	-0.11	1.71	1.10	4.39	0.14
	(5.03)	(2.85)	(1.27)	(2.53)	(0.43)	⟨0.43⟩	⟨0.58⟩	⟨0.36⟩	
NOK	-10.62	23.26	-8.24	-0.21	-0.21	0.25	3.57	5.76	0.21
	(65.22)	(74.87)	(21.61)	(1.29)	(0.14)	⟨0.88⟩	⟨0.17⟩	⟨0.22⟩	
Panel C: i = EUR									
CHF	0.76	-2.23	0.62	-0.50	0.02	0.30	0.56	2.16	0.07
	(4.67)	(8.65)	(3.49)	(0.67)	(0.24)	⟨0.86⟩	⟨0.76⟩	⟨0.71⟩	
CAD	-0.12	4.71	-1.63	2.31	-1.02	0.07	4.65	4.73	0.12
	(10.26)	(18.01)	(6.03)	(1.76)	(0.49)	⟨0.96⟩	⟨0.10⟩	⟨0.32⟩	
JPY	-0.68	-4.18	1.13	4.74	-1.25	0.30	3.86	5.50	0.06

CUR j	ψ_0	ψ_1	ψ_2	ψ_3	ψ_4	$\chi^2(2)_i$	$\chi^2(2)_j$	$\chi^2(4)$	R^2
	(7.57)	(11.09)	(3.77)	(3.77)	(0.72)	⟨0.86⟩	⟨0.15⟩	⟨0.24⟩	
AUD	5.30	0.58	0.34	-2.04	-0.16	0.21	7.07	7.10	0.11
	(4.52)	(18.12)	(9.50)	(4.54)	(0.91)	⟨0.90⟩	⟨0.03⟩	⟨0.13⟩	
SEK	-4.93	3.49	0.21	0.14	0.06	0.79	0.10	2.28	0.14
	(7.25)	(12.41)	(4.39)	(2.76)	(0.55)	⟨0.67⟩	⟨0.95⟩	⟨0.68⟩	
NOK	12.53	-14.11	5.30	-1.47	-0.29	0.10	2.27	4.94	0.49
	(33.93)	(55.20)	(19.19)	(1.32)	(0.37)	⟨0.95⟩	⟨0.32⟩	⟨0.29⟩	
Panel D: i = CHF									
CAD	2.21	0.67	0.06	3.29	-1.34	0.15	3.15	3.98	0.13
	(5.17)	(1.86)	(0.92)	(2.96)	(0.76)	⟨0.93⟩	⟨0.21⟩	⟨0.41⟩	
JPY	-2.39	0.38	-0.88	5.40	-1.18	1.01	1.55	1.92	0.07
	(7.89)	(2.05)	(0.91)	(5.57)	(1.05)	⟨0.60⟩	⟨0.46⟩	⟨0.75⟩	
AUD	1.76	1.36	0.63	-0.31	-0.39	2.22	3.13	6.66	0.16
	(6.59)	(1.07)	(0.83)	(5.63)	(1.00)	⟨0.33⟩	⟨0.21⟩	⟨0.15⟩	
SEK	-5.25	1.95	0.68	2.98	-0.35	5.03	1.56	16.35	0.20
	(3.91)	(1.22)	(0.40)	(2.57)	(0.51)	⟨0.08⟩	⟨0.46⟩	⟨0.00⟩	
NOK	4.00	-0.90	0.77	-1.41	-0.13	0.79	12.36	15.10	0.30
	(1.55)	(2.55)	(1.20)	(0.46)	(0.12)	⟨0.68⟩	⟨0.00⟩	⟨0.00⟩	
Panel E: i = CAD									
JPY	-4.93	-1.13	0.65	4.48	-1.28	0.84	0.97	3.91	0.10
	(4.49)	(2.14)	(0.74)	(4.90)	(1.32)	⟨0.66⟩	⟨0.61⟩	⟨0.42⟩	
AUD	3.45	-2.85	0.42	-0.17	0.04	1.79	0.02	2.07	0.14
	(4.28)	(2.15)	(0.47)	(4.43)	(0.73)	⟨0.41⟩	⟨0.99⟩	⟨0.72⟩	
SEK	0.20	-2.79	0.72	0.94	-0.34	2.26	0.95	3.45	0.04
	(3.01)	(1.95)	(0.62)	(1.82)	(0.36)	⟨0.32⟩	⟨0.62⟩	⟨0.49⟩	
NOK	-5.57	-1.17	1.57	-0.06	-0.16	10.07	1.71	22.94	0.46
	(7.70)	(9.87)	(2.09)	(1.34)	(0.16)	⟨0.01⟩	⟨0.42⟩	⟨0.00⟩	
Panel F: i = JPY									
AUD	14.87	-7.38	1.79	-4.27	0.11	3.89	5.38	8.82	0.27
	(11.36)	(3.79)	(0.99)	(6.48)	(1.28)	⟨0.14⟩	⟨0.07⟩	⟨0.07⟩	
SEK	2.71	-7.26	1.93	4.31	-1.08	4.23	1.80	6.58	0.11
	(8.88)	(3.67)	(0.99)	(3.85)	(0.87)	⟨0.12⟩	⟨0.41⟩	⟨0.16⟩	
NOK	15.20	-9.81	-0.87	-0.46	-0.52	3.08	3.60	15.44	0.25
	(3.99)	(8.68)	(5.60)	(1.89)	(0.29)	⟨0.21⟩	⟨0.16⟩	⟨0.00⟩	
Panel G: i = AUD									
SEK	-6.49	2.65	-0.33	3.44	-0.83	3.54	1.91	5.67	0.07
	(5.59)	(1.65)	(0.32)	(3.40)	(0.60)	⟨0.17⟩	⟨0.39⟩	⟨0.23⟩	
NOK	5.05	-3.83	1.25	-1.42	-0.35	0.28	9.72	11.20	0.11
	(20.84)	(22.98)	(4.67)	(1.15)	(0.11)	⟨0.87⟩	⟨0.01⟩	⟨0.02⟩	
Panel H: i = SEK									
NOK	-5.16	12.78	-2.85	-1.68	-0.37	1.21	3.13	13.62	0.26
	(24.81)	(33.47)	(10.57)	(1.42)	(0.22)	⟨0.55⟩	⟨0.21⟩	⟨0.01⟩	

Table 8: Forecasts of Excess Rates of Return in Currency Markets (Post-2003 Data)

The Table presents estimation results for the regression

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \psi_0 + \psi_1 x_{i,t} + \psi_2 x_{i,t}^2 + \psi_3 x_{j,t} + \psi_4 x_{j,t}^2 + \epsilon_{ij,t+1}^s$$

where the dependent variable is the excess rate of return in base currency i on an annual investment in the money market of currency j. The regressors are the CP factors, the fitted return forecasting variables from the term structure regressions for currencies i and j, and their squared values. Standard errors are in parentheses, and p-values are in angled brackets. The $\chi^2(2)_i$ statistic tests the null hypothesis that ψ_1 and ψ_2 equal zero, the $\chi^2(2)_j$ statistic tests the null hypothesis that ψ_3 and ψ_4 equal zero, and the $\chi^2(4)$ statistic tests the null hypothesis that ψ_1 through ψ_4 equal zero. The sample periods for the dependent variables all begin in 2004:12 and end in 2016:12.

CUR j	ψ_0	ψ_1	ψ_2	ψ_3	ψ_4	$\chi^2(2)_i$	$\chi^2(2)_j$	$\chi^2(4)$	R^2
				Panel A: i = USD					
GBP	4.47	1.11	-1.05	-2.30	0.08	1.15	2.98	7.39	0.26
	(2.46)	(1.93)	(0.98)	(2.66)	(1.19)	⟨0.56⟩	⟨0.23⟩	⟨0.12⟩	
EUR	10.16	1.05	-1.03	-8.27	1.61	3.23	2.63	4.10	0.21
	(7.49)	(1.95)	(0.76)	(9.59)	(3.57)	⟨0.20⟩	⟨0.27⟩	⟨0.39⟩	
CHF	-0.27	3.99	-1.11	0.47	-0.38	3.82	3.76	6.69	0.08
	(3.86)	(2.13)	(0.57)	(4.68)	(0.83)	⟨0.15⟩	⟨0.15⟩	⟨0.15⟩	
CAD	-0.14	-3.31	-0.38	6.10	-0.72	5.69	1.47	11.41	0.29
	(11.69)	(1.67)	(0.57)	(15.36)	(3.55)	⟨0.06⟩	⟨0.48⟩	⟨0.02⟩	
JPY	-11.21	-0.46	0.30	25.10	-10.42	0.09	1.29	4.85	0.18
	(10.92)	(6.67)	(1.75)	(39.14)	(25.15)	⟨0.95⟩	⟨0.52⟩	⟨0.30⟩	
AUD	6.37	2.17	-1.62	-1.76	0.66	2.36	0.14	2.47	0.11
	(4.30)	(2.64)	(1.22)	(4.69)	(2.04)	⟨0.31⟩	⟨0.93⟩	⟨0.65⟩	
SEK	9.92	4.49	-2.47	-7.75	1.44	4.29	1.09	6.50	0.32
	(10.31)	(3.93)	(1.34)	(13.39)	(3.32)	⟨0.12⟩	⟨0.58⟩	⟨0.16⟩	
NOK	9.98	1.63	-1.34	-18.87	7.97	7.35	0.66	7.51	0.36
	(16.96)	(3.17)	(0.90)	(25.97)	(9.95)	⟨0.03⟩	⟨0.72⟩	⟨0.11⟩	
				Panel B: i = GBP					
EUR	9.69	-1.55	0.42	-8.23	1.74	0.11	3.60	3.63	0.07
	(7.09)	(4.88)	(1.27)	(8.56)	(3.54)	⟨0.95⟩	⟨0.17⟩	⟨0.46⟩	
CHF	-0.43	-0.99	0.78	0.23	0.61	2.68	16.37	49.82	0.18
	(3.26)	(3.34)	(0.88)	(3.10)	(0.48)	⟨0.26⟩	⟨0.00⟩	⟨0.00⟩	
CAD	18.75	-3.38	1.50	-20.67	4.73	6.07	4.18	10.11	0.26
	(12.65)	(4.60)	(0.86)	(11.52)	(3.07)	⟨0.05⟩	⟨0.12⟩	⟨0.04⟩	
JPY	-20.52	8.89	-0.67	13.34	1.85	6.36	1.20	14.25	0.30
	(12.70)	(10.69)	(3.10)	(30.54)	(30.61)	⟨0.04⟩	⟨0.55⟩	⟨0.01⟩	
AUD	2.71	-1.91	1.17	-1.15	0.42	2.07	0.26	5.31	0.11
	(4.41)	(5.54)	(1.32)	(2.31)	(1.07)	⟨0.36⟩	⟨0.88⟩	⟨0.26⟩	
SEK	5.03	-4.30	1.03	-4.83	1.87	1.96	0.44	2.52	0.07
	(9.79)	(3.43)	(0.75)	(11.82)	(3.63)	⟨0.38⟩	⟨0.80⟩	⟨0.64⟩	
NOK	4.35	-3.71	0.89	-5.98	3.48	1.09	0.85	2.78	0.10
	(8.06)	(4.08)	(0.86)	(12.64)	(5.09)	⟨0.58⟩	⟨0.65⟩	⟨0.59⟩	
				Panel C: i = EUR					
CHF	-4.75	4.20	-0.39	1.20	-0.06	1.30	5.17	15.12	0.14
	(7.88)	(9.27)	(3.17)	(5.16)	(0.92)	⟨0.52⟩	⟨0.08⟩	⟨0.00⟩	
CAD	14.18	-4.51	2.47	-16.98	4.32	0.40	1.37	5.99	0.27
	(20.07)	(12.62)	(4.76)	(14.90)	(3.70)	⟨0.82⟩	⟨0.50⟩	⟨0.20⟩	
JPY	-14.89	-0.48	1.81	29.75	-16.12	1.16	3.57	4.63	0.16
	(10.12)	(14.93)	(5.35)	(26.74)	(23.34)	⟨0.56⟩	⟨0.17⟩	⟨0.33⟩	
AUD	7.77	-10.75	4.51	-1.03	0.14	0.22	0.07	0.67	0.08

CUR j	ψ_0	ψ_1	ψ_2	ψ_3	ψ_4	$\chi^2(2)_i$	$\chi^2(2)_j$	$\chi^2(4)$	R^2
	(14.66)	(24.11)	(9.64)	(5.39)	(3.23)	⟨0.90⟩	⟨0.97⟩	⟨0.95⟩	
SEK	5.67	0.53	0.53	-11.47	3.12	1.67	2.55	6.19	0.26
	(7.41)	(13.64)	(5.29)	(7.22)	(1.99)	⟨0.43⟩	⟨0.28⟩	⟨0.19⟩	
NOK	1.65	-2.98	2.34	-9.95	4.20	1.02	0.68	4.60	0.32
	(12.27)	(12.94)	(5.10)	(13.17)	(5.09)	⟨0.60⟩	⟨0.71⟩	⟨0.33⟩	
Panel D: i = CHF									
CAD	19.09	2.11	-0.55	-22.77	4.63	0.38	5.39	9.19	0.28
	(15.29)	(7.85)	(1.56)	(12.70)	(3.26)	⟨0.83⟩	⟨0.07⟩	⟨0.06⟩	
JPY	-14.94	7.87	-0.81	15.35	-6.27	13.48	1.94	26.27	0.22
	(8.93)	(3.83)	(0.63)	(21.39)	(16.59)	⟨0.00⟩	⟨0.38⟩	⟨0.00⟩	
AUD	1.91	3.97	-1.26	0.15	-1.11	6.78	0.78	7.93	0.05
	(6.19)	(8.35)	(1.48)	(2.05)	(1.78)	⟨0.03⟩	⟨0.68⟩	⟨0.09⟩	
SEK	4.19	-0.94	-0.57	-5.76	1.52	35.32	0.96	37.03	0.16
	(9.30)	(4.68)	(0.79)	(6.25)	(1.56)	⟨0.00⟩	⟨0.62⟩	⟨0.00⟩	
NOK	4.34	4.28	-1.24	-16.72	6.16	8.33	0.55	11.43	0.26
	(16.39)	(7.34)	(1.23)	(24.76)	(8.39)	⟨0.02⟩	⟨0.76⟩	⟨0.02⟩	
Panel E: i = CAD									
JPY	-19.66	8.40	-1.61	33.80	-19.89	0.46	1.51	3.47	0.20
	(15.09)	(14.50)	(3.34)	(27.90)	(18.16)	⟨0.80⟩	⟨0.47⟩	⟨0.48⟩	
AUD	-8.93	11.72	-1.58	-0.70	-0.45	6.77	1.33	11.57	0.37
	(7.85)	(8.49)	(1.70)	(1.09)	(0.65)	⟨0.03⟩	⟨0.51⟩	⟨0.02⟩	
SEK	-6.43	8.25	-1.44	-1.55	-0.17	0.57	0.59	2.01	0.11
	(19.38)	(17.61)	(4.21)	(8.88)	(2.70)	⟨0.75⟩	⟨0.75⟩	⟨0.73⟩	
NOK	-7.38	9.89	-2.29	-4.66	1.79	1.83	0.45	2.01	0.15
	(8.47)	(8.16)	(1.74)	(7.38)	(2.67)	⟨0.40⟩	⟨0.80⟩	⟨0.73⟩	
Panel F: i = JPY									
AUD	15.30	-19.91	12.63	-4.99	-0.30	2.50	3.34	4.54	0.16
	(6.04)	(12.82)	(10.13)	(4.66)	(3.16)	⟨0.29⟩	⟨0.19⟩	⟨0.34⟩	
SEK	15.34	-14.32	5.55	-6.16	0.23	0.67	3.43	5.63	0.12
	(18.53)	(23.37)	(19.84)	(20.36)	(5.88)	⟨0.72⟩	⟨0.18⟩	⟨0.23⟩	
NOK	24.79	-28.23	11.58	-26.13	9.28	1.67	0.46	2.95	0.29
	(33.01)	(27.53)	(18.72)	(38.52)	(13.98)	⟨0.43⟩	⟨0.79⟩	⟨0.57⟩	
Panel G: i = AUD									
SEK	-6.08	-0.96	0.70	6.29	-2.50	0.22	1.85	1.89	0.10
	(7.38)	(4.02)	(1.89)	(11.62)	(3.13)	⟨0.90⟩	⟨0.40⟩	⟨0.76⟩	
NOK	-1.52	-0.61	0.04	0.63	-1.52	0.08	0.90	1.10	0.09
	(5.59)	(3.50)	(2.09)	(5.63)	(1.83)	⟨0.96⟩	⟨0.64⟩	⟨0.89⟩	
Panel H: i = SEK									
NOK	-3.97	6.07	-1.11	-5.67	2.54	2.57	0.63	3.25	0.19
	(6.40)	(6.65)	(1.78)	(7.87)	(3.24)	⟨0.28⟩	⟨0.73⟩	⟨0.52⟩	

Table 9: Testing Uncovered Interest Rate Parity (Pre-2004 Data)

The Table presents estimation results for the regression

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \phi_0 + \phi_1(r_{j,t} - r_{i,t}) + \epsilon_{ij,t+1}$$

where the dependent variable is the excess rate of return in base currency i on an annual investment in the money market of currency j. The regressor is the difference in the one-year yields between country j and country i. Standard errors are in parentheses, and p-values are in angled brackets. The sample periods for the dependent variables all end in 2003:12. The samples begin in 1974:12 for the GBP and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK.

CUR j	ϕ_0	ϕ_1	$\chi^2(1)_{\phi_1=0}$	$\chi^2(1)_{\phi_1=1}$	R^2
Panel A: i = USD					
GBP	-3.16	2.00	5.18	1.29	0.10
	(2.38)	(0.88)	⟨0.02⟩	⟨0.26⟩	
EUR	2.16	1.68	5.10	0.83	0.09
	(2.05)	(0.74)	⟨0.02⟩	⟨0.36⟩	
CHF	3.13	2.67	6.60	2.58	0.22
	(3.08)	(1.04)	⟨0.01⟩	⟨0.11⟩	
CAD	-1.17	1.89	8.83	1.96	0.22
	(1.05)	(0.64)	⟨0.00⟩	⟨0.16⟩	
JPY	14.09	4.54	19.59	11.91	0.39
	(3.82)	(1.03)	⟨0.00⟩	⟨0.00⟩	
AUD	-3.98	2.46	8.83	3.12	0.23
	(2.88)	(0.83)	⟨0.00⟩	⟨0.08⟩	
SEK	-1.43	0.91	0.41	0.00	0.03
	(3.11)	(1.41)	⟨0.52⟩	⟨0.95⟩	
NOK	-7.57	5.45	28.86	19.24	0.62
	(2.67)	(1.01)	⟨0.00⟩	⟨0.00⟩	
Panel B: i = GBP					
EUR	4.06	1.47	4.83	0.49	0.09
	(2.05)	(0.67)	⟨0.03⟩	⟨0.48⟩	
CHF	7.91	2.93	9.38	4.07	0.22
	(2.89)	(0.96)	⟨0.00⟩	⟨0.04⟩	
CAD	0.59	2.43	2.15	0.74	0.09
	(1.90)	(1.66)	⟨0.14⟩	⟨0.39⟩	
JPY	19.10	4.24	17.96	10.49	0.14
	(4.97)	(1.00)	⟨0.00⟩	⟨0.00⟩	
AUD	-0.39	1.22	0.96	0.03	0.03
	(2.03)	(1.24)	⟨0.33⟩	⟨0.86⟩	
SEK	-1.39	1.50	1.62	0.18	0.08
	(1.72)	(1.18)	⟨0.20⟩	⟨0.67⟩	
NOK	0.05	2.52	10.07	3.65	0.27
	(1.57)	(0.79)	⟨0.00⟩	⟨0.06⟩	

CUR j	ϕ_0	ϕ_1	$\chi^2(1)_{\phi_1=0}$	$\chi^2(1)_{\phi_1=1}$	R^2
			Panel C: i = EUR		
CHF	1.52	1.71	2.02	0.35	0.05
	(1.62)	(1.20)	$\langle 0.16 \rangle$	$\langle 0.55 \rangle$	
CAD	-3.41	2.47	9.98	3.55	0.19
	(1.90)	(0.78)	$\langle 0.00 \rangle$	$\langle 0.06 \rangle$	
JPY	2.63	1.63	1.38	0.20	0.03
	(3.01)	(1.39)	$\langle 0.24 \rangle$	$\langle 0.65 \rangle$	
AUD	-1.67	1.39	2.52	0.20	0.11
	(2.17)	(0.88)	$\langle 0.11 \rangle$	$\langle 0.65 \rangle$	
SEK	-3.24	1.69	7.66	1.29	0.13
	(1.72)	(0.61)	$\langle 0.01 \rangle$	$\langle 0.26 \rangle$	
NOK	2.63	0.45	0.03	0.05	-0.01
	(5.39)	(2.52)	$\langle 0.86 \rangle$	$\langle 0.83 \rangle$	
			Panel D: i = CHF		
CAD	-9.07	3.91	15.44	8.55	0.25
	(2.77)	(0.99)	$\langle 0.00 \rangle$	$\langle 0.00 \rangle$	
JPY	5.37	4.39	1.91	1.14	0.08
	(4.92)	(3.17)	$\langle 0.17 \rangle$	$\langle 0.29 \rangle$	
AUD	-3.71	1.42	2.18	0.19	0.07
	(2.92)	(0.96)	$\langle 0.14 \rangle$	$\langle 0.66 \rangle$	
SEK	-8.40	2.76	7.42	3.02	0.18
	(2.98)	(1.02)	$\langle 0.01 \rangle$	$\langle 0.08 \rangle$	
NOK	-0.26	1.06	0.35	0.00	0.02
	(7.03)	(1.79)	$\langle 0.55 \rangle$	$\langle 0.97 \rangle$	
			Panel E: i = CAD		
JPY	22.23	5.80	24.46	16.75	0.38
	(5.49)	(1.17)	$\langle 0.00 \rangle$	$\langle 0.00 \rangle$	
AUD	0.49	0.45	0.31	0.45	0.00
	(2.22)	(0.82)	$\langle 0.58 \rangle$	$\langle 0.50 \rangle$	
SEK	-0.96	0.67	0.13	0.03	0.01
	(2.24)	(1.89)	$\langle 0.72 \rangle$	$\langle 0.86 \rangle$	
NOK	-6.65	4.72	6.33	3.93	0.28
	(3.12)	(1.87)	$\langle 0.01 \rangle$	$\langle 0.05 \rangle$	
			Panel F: i = JPY		
AUD	-14.91	3.45	15.15	7.64	0.30
	(5.35)	(0.89)	$\langle 0.00 \rangle$	$\langle 0.01 \rangle$	
SEK	-3.74	1.19	0.09	0.00	0.00
	(17.93)	(4.00)	$\langle 0.77 \rangle$	$\langle 0.96 \rangle$	
NOK	-69.96	12.42	23.25	19.66	0.45
	(15.32)	(2.58)	$\langle 0.00 \rangle$	$\langle 0.00 \rangle$	
			Panel G: i = AUD		
SEK	-1.86	-0.87	1.15	5.34	0.03
	(1.74)	(0.81)	$\langle 0.28 \rangle$	$\langle 0.02 \rangle$	
NOK	-0.11	1.50	0.76	0.08	0.00
	(3.36)	(1.72)	$\langle 0.38 \rangle$	$\langle 0.77 \rangle$	
			Panel H: i = SEK		
NOK	0.54	1.53	0.98	0.12	0.01
	(2.00)	(1.55)	$\langle 0.32 \rangle$	$\langle 0.73 \rangle$	

Table 10: Testing Uncovered Interest Rate Parity (Post-2003)

The Table presents estimation results for the regression

$$-\Delta s_{ij,t+1} + r_{j,t} - r_{i,t} = \phi_0 + \phi_1(r_{j,t} - r_{i,t}) + \epsilon_{ij,t+1}$$

where the dependent variable is the excess rate of return in base currency i on an annual investment in the money market of currency j. The regressor is the difference in the one-year yields between country j and country i. Standard errors are in parentheses, and p-values are in angled brackets. The sample periods for the dependent variables all begin in 2004:12 and end in 2016:12.

CUR j	ϕ_0	ϕ_1	$\chi^2(1)_{\phi_1=0}$	$\chi^2(1)_{\phi_1=1}$	R^2
		Panel A: i = USD			
GBP	0.34	-3.11	0.90	1.57	0.09
	(2.35)	(3.28)	$\langle 0.34 \rangle$	$\langle 0.21 \rangle$	
EUR	-1.19	-3.06	5.49	9.67	0.09
	(1.69)	(1.30)	$\langle 0.02 \rangle$	$\langle 0.00 \rangle$	
CHF	1.21	0.06	0.00	0.51	-0.01
	(2.12)	(1.32)	$\langle 0.96 \rangle$	$\langle 0.48 \rangle$	
CAD	2.44	-6.41	9.41	12.57	0.20
	(1.60)	(2.09)	$\langle 0.00 \rangle$	$\langle 0.00 \rangle$	
JPY	-1.54	-0.15	0.01	0.58	-0.01
	(4.10)	(1.51)	$\langle 0.92 \rangle$	$\langle 0.44 \rangle$	
AUD	6.76	-1.56	0.33	0.89	0.01
	(7.01)	(2.71)	$\langle 0.57 \rangle$	$\langle 0.35 \rangle$	
SEK	-0.93	-3.02	3.51	6.23	0.09
	(2.39)	(1.61)	$\langle 0.06 \rangle$	$\langle 0.01 \rangle$	
NOK	1.00	-2.40	2.55	5.11	0.07
	(2.46)	(1.51)	$\langle 0.11 \rangle$	$\langle 0.02 \rangle$	
		Panel B: i = GBP			
EUR	-0.32	-1.49	0.73	2.04	0.02
	(2.02)	(1.75)	$\langle 0.39 \rangle$	$\langle 0.15 \rangle$	
CHF	4.11	0.78	0.28	0.02	0.01
	(2.43)	(1.48)	$\langle 0.60 \rangle$	$\langle 0.88 \rangle$	
CAD	0.69	-3.96	13.38	21.00	0.22
	(1.79)	(1.08)	$\langle 0.00 \rangle$	$\langle 0.00 \rangle$	
JPY	-0.95	-0.69	0.10	0.58	0.00
	(5.44)	(2.22)	$\langle 0.76 \rangle$	$\langle 0.45 \rangle$	
AUD	4.25	0.12	0.01	0.30	-0.01
	(3.19)	(1.62)	$\langle 0.94 \rangle$	$\langle 0.59 \rangle$	
SEK	0.42	-0.35	0.07	0.98	-0.00
	(2.31)	(1.37)	$\langle 0.80 \rangle$	$\langle 0.32 \rangle$	
NOK	0.90	-0.59	0.27	1.98	0.00
	(2.26)	(1.13)	$\langle 0.60 \rangle$	$\langle 0.16 \rangle$	

CUR j	ϕ_0	ϕ_1	$\chi^2(1)_{\phi_1=0}$	$\chi^2(1)_{\phi_1=1}$	R^2
			Panel C: i = EUR		
CHF	5.15	3.96	3.42	1.91	0.17
	(2.55)	(2.14)	$\langle 0.06 \rangle$	$\langle 0.17 \rangle$	
CAD	2.54	-3.21	1.04	1.79	0.05
	(2.60)	(3.14)	$\langle 0.31 \rangle$	$\langle 0.18 \rangle$	
JPY	-0.81	-0.35	0.02	0.28	-0.01
	(4.76)	(2.57)	$\langle 0.89 \rangle$	$\langle 0.60 \rangle$	
AUD	0.87	1.05	0.11	0.00	-0.00
	(9.02)	(3.17)	$\langle 0.74 \rangle$	$\langle 0.99 \rangle$	
SEK	-0.09	-0.49	0.05	0.45	-0.01
	(1.61)	(2.21)	$\langle 0.83 \rangle$	$\langle 0.50 \rangle$	
NOK	1.51	-1.68	1.09	2.78	0.02
	(1.09)	(1.61)	$\langle 0.30 \rangle$	$\langle 0.10 \rangle$	
			Panel D: i = CHF		
CAD	-5.69	4.06	2.89	1.64	0.07
	(3.38)	(2.39)	$\langle 0.09 \rangle$	$\langle 0.20 \rangle$	
JPY	-3.01	-1.04	0.07	0.27	-0.00
	(3.53)	(3.94)	$\langle 0.79 \rangle$	$\langle 0.61 \rangle$	
AUD	-4.56	1.78	0.26	0.05	0.01
	(12.73)	(3.47)	$\langle 0.61 \rangle$	$\langle 0.82 \rangle$	
SEK	-2.23	-0.06	0.00	0.06	-0.01
	(3.90)	(4.27)	$\langle 0.99 \rangle$	$\langle 0.80 \rangle$	
NOK	4.06	-3.75	0.93	1.49	0.03
	(7.82)	(3.89)	$\langle 0.33 \rangle$	$\langle 0.22 \rangle$	
			Panel E: i = CAD		
JPY	0.52	1.26	0.22	0.01	0.01
	(4.34)	(2.67)	$\langle 0.64 \rangle$	$\langle 0.92 \rangle$	
AUD	2.25	0.10	0.00	0.22	-0.01
	(4.02)	(1.94)	$\langle 0.96 \rangle$	$\langle 0.64 \rangle$	
SEK	-2.00	-2.29	3.38	6.99	0.04
	(1.55)	(1.24)	$\langle 0.07 \rangle$	$\langle 0.01 \rangle$	
NOK	-1.16	-0.06	0.00	1.17	-0.01
	(1.38)	(0.98)	$\langle 0.95 \rangle$	$\langle 0.28 \rangle$	
			Panel F: i = JPY		
AUD	8.91	-1.24	0.12	0.40	0.01
	(13.59)	(3.55)	$\langle 0.73 \rangle$	$\langle 0.53 \rangle$	
SEK	3.37	-2.26	0.32	0.67	0.02
	(6.14)	(3.99)	$\langle 0.57 \rangle$	$\langle 0.41 \rangle$	
NOK	6.65	-2.92	0.56	1.01	0.05
	(9.29)	(3.90)	$\langle 0.45 \rangle$	$\langle 0.31 \rangle$	
			Panel G: i = AUD		
SEK	-1.95	0.75	0.08	0.01	-0.00
	(7.89)	(2.60)	$\langle 0.77 \rangle$	$\langle 0.92 \rangle$	
NOK	-6.15	-1.39	1.83	5.42	0.03
	(2.93)	(1.03)	$\langle 0.18 \rangle$	$\langle 0.02 \rangle$	
			Panel H: i = SEK		
NOK	1.86	-2.25	1.92	4.00	0.05
	(1.41)	(1.63)	$\langle 0.17 \rangle$	$\langle 0.05 \rangle$	

Table 11: Out-of-Sample Forecasts for Excess Bond Returns: Cochrane-Piazzesi Models vs. Historical Means

The Table reports two statistics that compare the out-of-sample forecasts from recursive estimations of the Cochrane and Piazzesi (2005) model for the excess rates of returns on bonds denominated in different currencies to the forecasts based only on the historical mean excess rates of return. The first statistic is the R^2, which is calculated as one minus the ratio of the mean squared error of the CP forecasts to the mean squared error of the historical mean. The second statistic tests the equality of the forecasts and is the Clark and McCracken (2005) $MSE-F$ statistic. The sample periods for the dependent variables during the initial in-sample estimation all end in 2003:12, which is the end of the Cochrane and Piazzesi (2005) sample. The samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The out-of-sample periods are all 2004:01 to 2016:12.

	R^2				MSE-F			
CUR	$rx_{t+1}^{(2)}$	$rx_{t+1}^{(3)}$	$rx_{t+1}^{(4)}$	$rx_{t+1}^{(5)}$	$rx_{t+1}^{(2)}$	$rx_{t+1}^{(3)}$	$rx_{t+1}^{(4)}$	$rx_{t+1}^{(5)}$
USD	-1.51	-1.95	-1.90	-1.68	-87.26	-95.86	-94.97	-90.85
GBP	0.06	0.10	0.11	0.09	9.29	16.50	18.14	15.01
EUR	-0.10	-0.17	-0.25	-0.31	-13.20	-20.65	-28.77	-34.07
CHF	-0.16	0.05	0.05	0.01	-19.75	7.64	7.99	0.73
CAD	0.01	0.00	-0.06	-0.15	1.36	0.62	-7.59	-19.13
JPY	-0.23	-0.13	-0.17	-0.21	-26.63	-16.26	-20.54	-24.88
AUD	-2.28	-2.26	-2.21	-2.08	-100.78	-100.48	-99.86	-97.98
SEK	-0.14	-0.08	-0.01	0.03	-17.69	-11.16	-2.11	4.86
NOK	-1.07	-1.18	-1.17	-1.07	-74.81	-78.62	-78.09	-74.91

Table 12: Out-of-Sample Forecasts of Excess Bond Returns: Cochrane-Piazzesi Models with Free Constants vs. Historical Means

The Table reports two statistics that compare the out-of-sample forecasts from recursive estimations of an alternative version of the Cochrane and Piazzesi (2005) model, which allows for free constant terms, for the excess rates of returns on bonds denominated in different currencies to the forecasts based only on the historical mean excess rates of return. The first statistic is the R^2, which is calculated as one minus the ratio of the mean squared error of the CP forecasts to the mean squared error of the historical mean. The second statistic tests the equality of the forecasts and is the Clark and McCracken (2005) $MSE-F$ statistic. The sample periods for the dependent variables during the initial in-sample estimation all end in 2003:12, which is the end of the Cochrane and Piazzesi (2005) sample. The samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The out-of-sample periods are all 2004:01 to 2016:12.

	R^2				MSE-F			
CUR	$rx_{t+1}^{(2)}$	$rx_{t+1}^{(3)}$	$rx_{t+1}^{(4)}$	$rx_{t+1}^{(5)}$	$rx_{t+1}^{(2)}$	$rx_{t+1}^{(3)}$	$rx_{t+1}^{(4)}$	$rx_{t+1}^{(5)}$
USD	-1.36	-2.14	-2.23	-2.21	-83.51	-98.89	-100.06	-99.87
GBP	-0.07	-0.08	-0.10	-0.13	-9.01	-10.57	-13.42	-16.34
EUR	-0.08	-0.11	-0.20	-0.28	-10.21	-14.39	-23.80	-31.88
CHF	-0.36	0.04	0.10	0.07	-38.24	6.13	16.57	11.38
CAD	-0.21	-0.13	-0.16	-0.27	-25.50	-16.85	-20.13	-31.15
JPY	-0.17	-0.09	-0.14	-0.21	-21.34	-11.59	-18.35	-24.71
AUD	-1.89	-2.08	-2.34	-2.42	-94.90	-97.91	-101.57	-102.66
SEK	-0.15	-0.12	-0.08	-0.06	-18.49	-15.37	-11.31	-7.84
NOK	-0.60	-0.82	-0.98	-1.07	-54.17	-65.24	-71.82	-75.12

Table 13: Out-of-Sample Forecasts for Excess Foreign Exchange Returns: Cochrane-Piazzesi Factors vs. Historical Means

The Table reports two statistics that compare the out-of-sample forecasts from recursive estimation of equation (24) for the excess return in USD on one-year investments in the money markets of different currencies to the forecasts based on the historical mean excess rates of return on those currencies. The first statistic is the R^2, which is calculated as one minus the ratio of the mean squared error of the CP forecasts to the mean squared error of the historical mean. The second statistic tests the equality of the forecasts and is the Clark and McCracken (2005) $MSE-F$ statistic. The sample periods for the dependent variables during the initial in-sample estimation all end in 2003:12, which is the end of the Cochrane and Piazzesi (2005) sample. The samples begin in 1974:12 for the USD, the GBP, and the EUR; in 1989:03 for the CHF; in 1987:03 for the CAD; in 1986:03 for the JPY; in 1988:04 for the AUD; in 1988:03 for the SEK; and in 1999:03 for the NOK. The out-of-sample periods are all 2004:01 to 2016:12.

CUR	R^2	MSE-F
GBP	-0.04	-6.08
EUR	-0.37	-38.86
CHF	-0.57	-52.63
CAD	-0.27	-31.09
JPY	-0.26	-29.91
AUD	-0.84	-66.07
SEK	-0.56	-52.33
NOK	-0.56	-52.02