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International evidence on bond risk premia

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ABSTRACT

This paper revisits the study of time-varying excess bond returns in international bond markets. Using newly available yield curve data from 10 different countries with independent monetary policy, I test the robustness of Cochrane and Piazzesi (2005). For most countries in my sample, I find more modest predictive power for forward rates than originally found by Cochrane and Piazzesi (2005) for the US. Their single-factor model captures well the predictability in international data, and this factor also tends to have a tent-shape in most countries of my sample. CP factors are more idiosyncratic across countries than yields or forward rates. Finally, I show that the recent financial crisis has significantly affected the predictability of excess bond returns.

1. Introduction

The study of time-varying risk premia in bond markets has generated an immense literature in economics and finance. Early important contributions by Fama and Bliss (1987), Campbell and Shiller (1991) presented strong evidence against the expectations hypothesis in the US bond market. More recently, Cochrane and Piazzesi (2005) show that a single factor, a linear combination of forward rates (CP factor, henceforth), is a stronger predictor of one-year ahead excess bond returns of all maturities than traditional instruments, such as yield spreads and the spread between forward and one-year yield rates. Furthermore, Cochrane and Piazzesi (2005, 2008) show that the CP factor cannot be captured by the popular yield curve factors of level, slope and curvature. Recently, some authors[1] have incorporated the CP factor as a stylized fact to be matched by term structure models.

Data from different countries might provide valuable information on the robustness of the CP factor. The vast majority of papers on bond risk premia still use yield curve data from only one country at a time, ignoring an important source of additional information on the dynamics of yield curves that could be provided by a cross-section of different countries. In particular, I use bond yield curves estimated for 10 countries with independent monetary policies to shed light on the following questions: Does a linear combination of forward rates predict excess return in different countries? How similar are the bond return forecasting factors for the different countries? Has the recent financial crisis impacted the predictability of excess bond returns?

This paper brings new evidence on international bond risk premia using a panel of nominal zero-coupon government bond yields constructed by Wright (forthcoming). The dataset includes data for Australia, Canada, Japan, Germany, New Zealand, Norway, Switzerland, Sweden, the United Kingdom and the United States with rates from 3 months to 10 years. I show that the main results of Cochrane and Piazzesi (2005) hold for this diversified set of developed countries, namely that the same combination of forward rates forecast monthly excess returns of different maturities. Moreover, for most of countries, this linear combination also has a tent-shape. On average, I find smaller R^2s than Cochrane and Piazzesi (2005). CP factors rise sharply at the start of recessions and are considerably volatile. Finally, CP factors are found to be more country idiosyncratic than traditional yield curve factors.

This paper is related to a growing literature analyzing yield curves in multiple international markets. Ilmanen (1995) used monthly data on long term government bonds from six different developed countries[2] in order to study the predictable variation of

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[1] See Campbell et al. (2009) and Xiong and Yan (2010).

[2] United States, Canada, Japan, Germany, France and the United Kingdom.

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their respective monthly excess returns from January 1978 to June 1993. Diebold et al. (2008) fitted yield curve data for Germany, Japan, the United Kingdom, and the United States to study a global yield curve model,[3] where the dynamics of yields depend on local and global factors. They find that global yield curve factors explain a significant portion of a country's yield curve level and slope. Wright (forthcoming) constructs the yield curve dataset used in this paper and studies the macroeconomic determinants of risk premia. He focuses on inflation uncertainty as a prominent candidate for positive and time-varying risk premia in government bonds. Using different measures of inflation uncertainty, he shows a significant positive relationship between the two. Consequently, he argues that a significant share of the reduction in nominal interest rates observed in the last 15 years is due to reduced inflation uncertainty brought about by improved monetary policy. Finally, Kessler and Scherer (2009) use forward rates and estimate CP factors to study time-varying risk premia in a set of seven countries.[4] The authors use swap rates to study the predictability of the 1-year zero coupon rate using forward rates. They find strong evidence of predictability, with R^2 reaching numbers in excess of 60% for the 1-year ahead forecast in countries like Australia, Canada and the US. A potential drawback of the paper is the short sample used, February 1997 to July 2007.[5]

The remainder of the paper is as follows: the next section briefly discusses the related literature. Section 3 presents the econometric framework. Section 4 goes over data issues. The next Section 5, presents the results. Finally, Section 6 concludes.

2. Econometric framework

2.1. Bond notation

Before describing the model to be estimated, I briefly detail the notation. Let $P_{n,t}$ denote the price of an n-month zero coupon bond in month t. The yield on this bond is given by

$$y_{n,t} = -\frac{1}{n} log(P_{n,t})$$

The 12-month forward rate ending n months hence is

$$f_{n,t} \equiv log(P_{n-12,t}) - log(P_{n,t})$$

The return from buying an n-month bond in month $t-12$ and selling it as an $n-12$-month bond in month t is

$$r^{(n)}_{t-12,t} \equiv log(P_{n-12,t}) - log(P_{n,t-12})$$

Thus, the excess return from holding an n-month bond for 12 months over holding a 12-month bond for that same holding period is

$$rx^{(n)}_{t-12,t} \equiv log(P_{n-12,t}) - log(P_{n,t-12}) - y_{1,t-12}$$

2.2. Bond forecasting

As in Cochrane and Piazzesi (2005), I regress excess bond returns on initial forward rates. In order to minimize the almost perfect collinearity problem, I follow Bansal et al. (2004), Cochrane and Piazzesi (2008), Wright and Zhou (2009), Singleton (2005) and use three forward rates, instead of five, and estimate the following predictive regressions

$$rx^{(n)}_{t-12,t} = \beta_{0,n} + \beta_{1,n} y_{12,t-12} + \beta_{2,n} f_{36,t-12} + \beta_{3,n} f_{60,t-12} \varepsilon^{(n)}_{t-12,t} \quad (1)$$

for n = 24, 36, 48, 60. The fact that a similar pattern of coefficients was found for every bond maturity led Cochrane and Piazzesi (2005) to estimate a restricted version of the forecasting regression above, imposing that the coefficients on the forward rates are equal for every maturity n, up to a constant.

$$rx^{(n)}_{t-12,t} = b_n(\gamma_0 + \gamma_1 y_{12,t-12} + \gamma_2 f_{36,t-12} + \gamma_3 f_{60,t-12}) + \varepsilon^{(n)}_{t-12,t} \quad (2)$$

In order to estimate the model, I follow their two-step procedure. First, I estimate γ by regressing the average excess return across all maturities on all forward rates,

$$rx^{*}_{t-12,t} = \gamma_0 + \gamma_1 y_{12,t-12} + \gamma_2 f_{36,t-12} + \gamma_5 f_{60,t-12} + \varepsilon^{*}_{t-12,t} \quad (3)$$

where $rx^{*}_{t-12,t} = \frac{1}{4}\sum_{n=2}^{5} rx^{12n}_{t-12,t}$, to identify the forecasting factor CP_t

$$CP_{t-12} = \hat{\gamma}_0 + \hat{\gamma}_1 y_{12,t-12} + \hat{\gamma}_2 f_{36,t-12} + \hat{\gamma}_5 f_{60,t-12}$$

The b_n are estimated by regressing the excess returns for every maturity on the fitted values of the above regression.

$$rx^{(n)}_{t-12,t} = b_n(CP_{t-12}) + \varepsilon^{(n)}_{t-12,t} \quad (4)$$

In order to identify the values of γ_j ($j = 0,1,3,5$) and b_n ($n = 2,3,4,5$), I follow Cochrane and Piazzesi (2005) and impose that $\frac{1}{4}\sum_{n=2}^{5} b_{(n)} = 1$. The 2-stage regression imposes that the coefficients of the unrestricted regression (1) and the restricted regression (2) are related by $\beta_{j,n} = \gamma_j b_n$.

In the next section, I further examine the model above using term structure data from 10 different countries. Before presenting the empirical results, I provide more details and a brief discussion about my sample.

3. Data

I use end-of-month yield curve data for 10 different countries assembled and constructed by Wright (forthcoming): Australia, Canada, Germany, Japan, New Zealand, Norway, Sweden, Switzerland, UK and United States.[6] Table 1 gives more detail on the

Table 1
Yield curve data sources.

Country	Source	Start Date	Methodology
Australia	Datastream and Wright (forthcoming)	Feb 1987	Nelson–Siegel
Canada	Bank of Canada and BIS database	Jan 1986	Spline
Germany	Bundesbank and BIS database	Jan 1973	Svensson
Japan	Datastream and Wright (forthcoming)	Jan 1985	Svensson
Norway	Norges Bank and BIS database	Jan 1998	Svensson
New Zealand	Datastream and Wright (forthcoming)	Jan 1990	Nelson–Siegel
Switzerland	Swiss National Bank and BIS database	Jan 1988	Svensson
Sweden	Riksbank and BIS database	Dec 1992	Svensson
UK	Anderson and Sleath (1999)	Jan 1975	Spline
US.	Gürkaynak, Sack and Wright (2007)	Nov 1971	Svensson

Note: This table shows the primary source, as well as the starting date and methodology of yield curve estimation for every country in the sample. For further details, see Wright (forthcoming).

[3] See also Brennan and Xia (2006), Dungey et al. (2000) for other papers estimating term structure models for a panel of countries.

[4] Australia, Canada, Germany, Japan, Switzerland, United Kingdom and United States.

[5] Using a different dataset (the Fama-Bliss CRSP data) and a longer sample (January 1964 to December 2003), Cochrane and Piazzesi (2005) find, generally, R^2s in the neighborhood of 30% for the US.

[6] Unlike Cochrane and Piazzesi (2005), I do not use the Fama-Bliss dataset for the US., but the Gürkaynak et al. (2007) dataset, available at the Federal Reserve Board website.

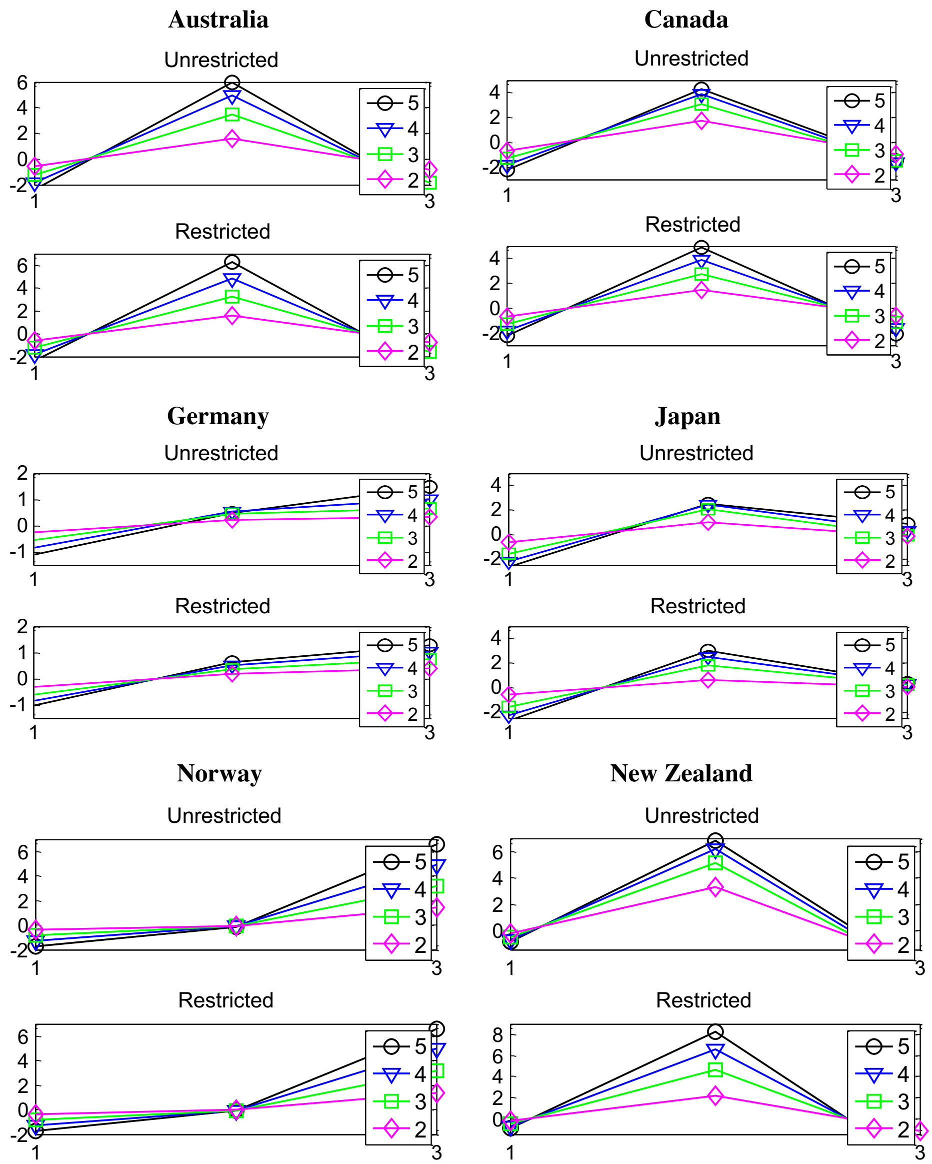

Fig. 1. Regression coefficients of 1-year excess return on forward rates. *Notes*: For every country, the first graph presents the $\beta_{j,n}$ from the unrestricted regressions on three forward rates. The second graph shows $\gamma_j b_n$, the coefficients from the restricted regression with the single factor. The legend represents the maturity of the bond being forecasted, and the *x*-axis shows the forward rates (first, third and fifth forward rates).

sources and starting date of each series, as well as on the methodology used in the estimation of the yield curve for each country. The reader is referred to Wright (forthcoming) for a more detailed description about the construction of the dataset.

In contrast to Cochrane and Piazzesi (2005), who use unsmoothed Fama-Bliss (FB) yield curve data from the Center for Research on Securities Prices (CRSP), most of the data in this paper come from smoothed curves. The yield curve for each country was constructed by their respective Central Banks. The advantage of using data created by local sources lies in their comparative informational advantage at constructing their respective yield curves. The Svensson (1994) estimation method is the most common technique employed. The New Zealand and Australian yield curves are constructed with the closely related Nelson and Siegel (1987) method, and the British and Canadian ones, with Spline methods.

As discussed in Cochrane and Piazzesi (2008), the choice of whether or not to smooth the yield curve may have important implications when using forward rates to forecast excess returns. Smoothing has the benefit of removing measurement errors from the data, at the cost of potentially taking away some of its forecasting signal, especially if the information not summarized by the traditional factors of level, slope and curvature provide important information for forecasting purposes. Recently, Cochrane and Piazzesi (2008), Duffee (2008) have argued that the fourth and fifth principal components of bond yields have forecasting power for the dynamics of the yield curve, even though they only explain a negligible share of the cross-section of yields. Yield curves estimated with the Nelson–Siegel–Svensson (NSS) methods will not keep these principal components, whereas yield curves estimated by Spline methods might. In the next section, I conduct some robustness tests on the impact of yield curve estimation technique on the predictability evidence by means of forward rates.

4. Empirical results and discussion

Similar to Cochrane and Piazzesi (2005), I find that a single linear combination of forward rates forecasts excess bond returns of

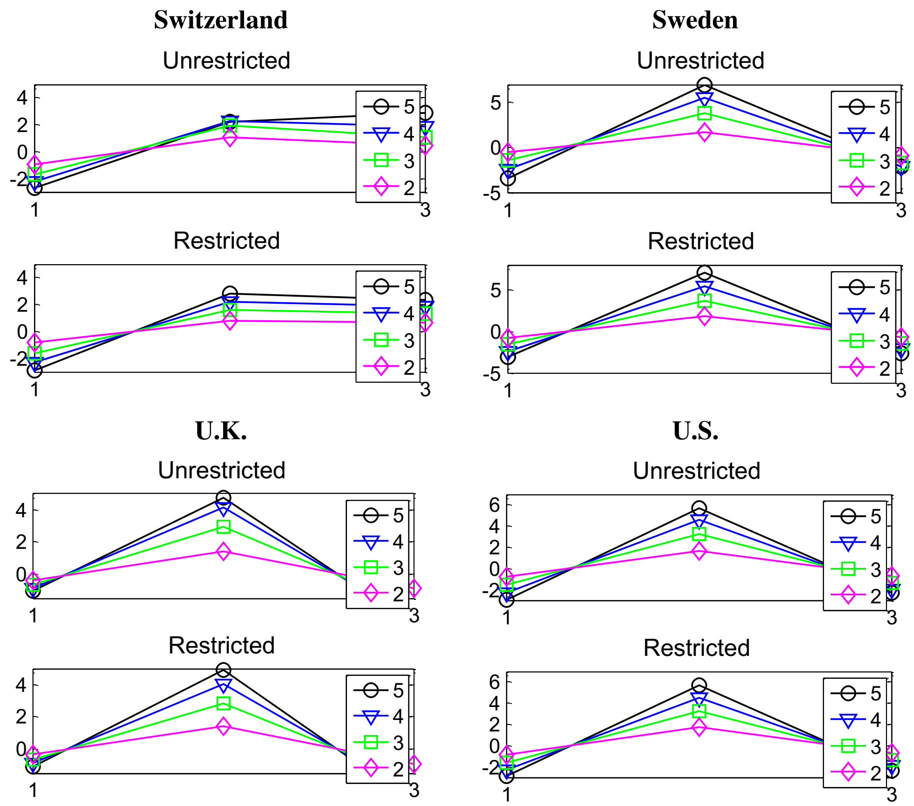

Fig. 2. Regression coefficients of 1-year excess return on forward rates. *Notes*: For every country, the first graph presents the $\beta_{j,n}$ from the unrestricted regressions on three forward rates. The second graph shows $\gamma_j b_n$, the coefficients from the single-factor restricted regression. The legend represents the maturity of the bond being forecasted, and the *x*-axis shows the forward rates (first, third and fifth forward rates).

all maturities for all countries in my sample. With few exceptions,[7] this single factor also has a tent-shape. Figs. 1 and 2 show the pattern of coefficients of the unrestricted as well as the restricted regressions. It is visible that the single-factor model captures the unrestricted parameter estimates very well, even for the countries where this factor does not present a clear tent-shape. Table 2 presents the R^2s for the unrestricted forecasting regression (1), whereas Table 3 shows the results for the single-factor regression given by Eq. (4). For every country in my sample, the fit of the regressions is only very marginally affected by the restriction of a single forecasting factor. Unrestricted and restricted regressions have almost identical R^2s. Individual hypothesis tests of the equality between the unrestricted and restricted coefficients are not able to reject the null of coefficient equality and produce average t-statistics around 0.10. More formally, we can test the restriction that the parameters of the unrestricted return forecasting regressions (1), $\beta_{j,n}$ and the ones from the restricted single-factor regression (2), $\gamma_j b_n$ are *jointly* equal.

I cannot reject the restrictions that $\beta_{j,n} = \gamma_j b_n$ for six, out of the 10 countries in my sample, as shown by Table 4. Cochrane and Piazzesi (2005) reject the restrictions implied by the model using unsmoothed Fama-Bliss bond data for the US and credit this rejection to the presence of measurement errors. Under the assumption that measurement errors are serially uncorrelated, I lag the right-hand side variables by 1 and 2 months, respectively. As a result, the evidence against the single-factor model weakens considerably, and Sweden and the UK are the only countries where the single-factor restrictions are still rejected. In an online appendix, I show the return forecasting factors estimated with lagged forward rates. These are very similar to those obtained using time-t forward rates.

There is some evidence of heterogeneity in the forecastability of bond yields across the different countries. Even though one has to be careful with rigid comparisons, since my panel is unbalanced, I find significant differences even among countries that have samples that span similar time periods. For example, in New Zealand, the combination of forward rates is able to forecast about 50% of the returns on 5-year bonds, while in Sweden forward rates forecast around 27% of the return on a bond of the same maturity. In Japan and Switzerland, the 5-year bond is predicted with R^2s of 27–33%, whereas in Canada, a country whose data are from about the same period, this number is slightly smaller than 15%.

For most countries with longer samples (Australia, Canada, Germany, UK and US.), a linear combination of forward rates predicts around 10% of the two-year bond, with that number rising to about 15% for the five-year bond. Those numbers are below the ones found by Cochrane and Piazzesi (2005) for the US. with the FB dataset. As shown by Cochrane and Piazzesi (2008), smoothed yield curves explains some of the difference for the US. One might conjecture that the use of smoothed yield curves for most of countries also explains the lower predictability of their respective excess bond returns. Comparing the forecasting performance of unrestricted regressions (see Table 2) does not reveal any particular forecasting power gain in regressions using Spline yield curve data. The R^2s of the forecasting regressions for Canada and the UK, countries with a Spline estimated yield curve, are not greater than the ones from countries with NSS yield curves and samples of comparative size, as Australia and Japan for Canada, and

[7] Norway presents a clear inverse tent-shape pattern of coefficients, whereas in Germany and Switzerland, the third and fifth forward rates have similar coefficient values.

Table 2
Estimates of the return forecasting factor.

	γ_0	γ_1	γ_2	γ_3	R^2
Australia					
ols	−2.26	−1.10	3.26	−1.70	14.47
se	2.13	0.36	0.87	1.01	
Canada					
ols	0.18	−1.32	3.26	−1.77	13.43
se	1.79	0.44	1.23	1.11	
Germany					
ols	−2.69	−0.60	0.40	0.70	12.63
se	1.48	0.55	1.27	0.95	
Japan					
ols	−1.28	−1.85	2.11	0.22	26.70
se	0.73	0.74	1.86	1.28	
Norway					
ols	−13.25	−1.07	0.56	3.05	28.15
se	3.68	0.54	1.91	1.79	
New Zealand					
OLS	−9.88	−0.31	2.69	−0.85	48.27
se	1.87	0.43	1.26	0.96	
Switzerland					
ols	−5.03	−1.63	1.05	1.86	35.12
se	1.51	0.42	1.64	1.26	
Sweden					
ols	−2.64	−1.41	4.09	−1.92	25.48
se	1.67	0.97	1.91	1.29	
UK					
ols	−1.33	−1.11	3.67	−2.26	18.02
se	1.39	0.40	0.99	0.78	
US					
ols	−1.45	−1.50	2.62	−0.88	14.48
se	2.01	0.57	1.56	1.22	
US FB					
ols	−1.66	−1.10	2.96	−1.40	22.01
se	1.72	0.61	1.20	0.84	

Note: This table shows the estimates of the unrestricted return forecasting factor in Eq. (3) $rx^*_{t-12,t} = \gamma_0 + \gamma_1 y_{12,t-12} + \gamma_2 f_{36,t-12} + \gamma_5 f_{60,t-12} + \varepsilon^*_{t-12,t}$, with Newey–West standard errors with 18 lags, where $rx^*_{t-12,t}$ is the average (across maturity) excess returns. US FB reports the estimates using the unsmoothed Fama-Bliss dataset (with data ending in December of 2008), whereas US reports the estimates using the Gürkaynak et al. (2007) dataset.

Germany and the US. for the UK.[8] Nevertheless, the UK and Canada, countries with relatively long samples and Spline yield curves, also present more modest evidence of predictability.

In order to evaluate directly the impact of using NSS yield curves on predictability, I estimate NSS yield curves for Canada and the UK and make their yield curves totally comparable with the remaining international yield curves from my dataset.[9] Predictability is smaller when one uses yield curves estimated with the NSS technique, but this effect is small, as shown by Table 5. For Canada, the R^2 decreases from 13.43% to 11.10%. In the UK, predictability falls slightly more, from 18.02% to 15.77%. For both countries, the general pattern of coefficients remain the same. Consequently, the choice of yield curve estimation does not seem to explain the differences in predictability between countries. The sample period seems to be a far more important factor. The last row of Table 2 presents estimates of the return forecasting factor with Fama-Bliss dataset for November 1971 to December 2008.[10]

Table 3
Restricted model–domestic factor.

	AUS	CAN	GER	JAP	NOR	NZ	CH	SE	UK	US
$n = 2$										
b_2	0.39	0.47	0.45	0.36	0.34	0.43	0.44	0.41	0.41	0.46
se	0.07	0.04	0.05	0.03	0.05	0.03	0.06	0.04	0.04	0.05
R^2	0.08	0.12	0.11	0.20	0.13	0.41	0.28	0.18	0.14	0.12
$n = 3$										
b_3	0.84	0.86	0.87	0.81	0.79	0.84	0.84	0.83	0.84	0.85
se	0.03	0.04	0.04	0.03	0.04	0.02	0.04	0.03	0.03	0.04
R^2	0.13	0.13	0.12	0.25	0.22	0.46	0.33	0.24	0.18	0.13
$n = 4$										
b_4	1.22	1.19	1.20	1.23	1.23	1.21	1.20	1.21	1.22	1.19
se	0.02	0.01	0.01	0.01	0.01	0.01	0.03	0.01	0.01	0.02
R^2	0.15	0.13	0.13	0.27	0.30	0.49	0.35	0.26	0.19	0.15
$n = 5$										
b_5	1.55	1.49	1.47	1.60	1.64	1.53	1.53	1.55	1.53	1.50
se	0.08	0.07	0.07	0.06	0.08	0.04	0.08	0.05	0.07	0.08
R^2	0.16	0.14	0.13	0.28	0.36	0.49	0.37	0.27	0.18	0.16

Note: This table displays estimates of the restricted regression $rx^{(n)}_{t-12,t} = b_n(CP_{t-12}) + \varepsilon^{(n)}_{t-12,t}$ for each country in the sample for n = 2, 3, 4, and 5.

The R^2 of the regression is approximately 22%, significantly lower than the original Cochrane and Piazzesi (2005) estimates. I show below that the ongoing financial crisis has hindered the predictability of excess bond returns.

Countries that display the highest evidence in favor of predictability are among the ones with the smallest sample, Norway and New Zealand. This might be reflecting the fact that a regression with persistent data tends to fit well when estimated over a short period, or a structural break in the parameters of the model. In either case, there is a clear tendency for less predictability for countries whose data go back to the early 1970s.

In a recent paper, Kessler and Scherer (2009) examine the predictability of excess bond returns on seven countries which are also present in my sample, using swap data on zero coupon bonds for the period of February 1997 to July 2007. The authors find much stronger evidence of predictability for all countries in their sample. I examine the evidence of predictability for this same sample period using my dataset. As Table 6 shows, results for this restricted sample are broadly in line with Kessler and Scherer (2009). Nevertheless, there is a significant drop in predictability as I extend the sample. Moreover, the decrease in predictability seems to be higher, the longer the available sample. This could be reflecting a structural break, as the late 1990s and the 2000s (before the financial crisis) were a period of relatively moderate volatility. There is considerable evidence across all countries that predictability decreases with a longer sample period.

Overall, Cochrane and Piazzesi (2005) main finding that a single linear combination of forward rates forecasts excess returns on bonds of one- to five-years of maturity, is also verified in a diversified set of developed countries. Moreover, for most of these countries, the pattern of coefficients of the return forecasting regression also displays a tent-shape. However, using smoothed yield curve data and a sufficiently long sample, the evidence in favor of predictability is not as strong as the one found in the US. bond market with the FB dataset, and is considerably smaller than the ones provided by Kessler and Scherer (2009) using swap rates for six different countries with a more limited sample size. One might conjecture what are the consequences of smaller predictability evidence for "trading" rules profits based on the estimated models. The online appendix includes results on profitability of a real-time "trading" rule. Although these "trading" rule profits are smaller than the ones reported in the online appendix by Cochrane and Piazzesi (2005), cumulative profits are still positive for the majority of countries in my sample.

[8] Admittedly, different countries, even with an exact same sample length, do not provide perfect counter-factual evidence on the impacts of Spline or NSS yield curve for forecasting purposes. The suggestion made above refers simply to the absence of any strong observable impact.

[9] It is not possible, though, to recover Spline yield curves from NSS yield curves, because the latter are already very smoothed and the Spline would basically fit the same yield curve.

[10] December 2008 is the last available observation in the Fama-Bliss dataset as of this date.

Table 4
GMM tests of single-factor.

Lag	Test	AUS	CAN	GER	JAP	NOR	NZ	CH	SE	UK	US
0	Wald	19.31	10.75	9.76	8.75	3.40	19.51	16.42	23.24	32.72	9.14
	p-value	2.27	29.36	36.99	46.05	94.62	2.12	5.85	0.57	0.01	42.47
1	Wald	9.33	6.98	8.93	6.63	2.68	12.18	13.07	22.40	26.25	9.24
	p-value	40.74	63.93	44.33	67.51	97.58	20.35	15.93	0.77	0.19	41.56
2	Wald	5.10	6.59	9.11	5.49	3.53	9.60	13.03	22.95	24.77	9.35
	p-value	82.53	67.96	42.70	78.97	93.96	38.36	16.12	0.63	0.32	40.59

Note: Wald tests the joint parameter restriction that the unrestricted coefficients, $\beta_{j,n}$, and the single-factor restricted coefficients, $\gamma_j b_n$, are statistically equal. The 5% critical value for all tests is 16.9. Lag 1 and 2 refers to tests of equality of coefficients when the regressions are estimated with the forward rates lagged by 1 and 2 months, respectively.

Table 5
Contrasting predictability with Spline and NSS yield curves.

	γ_0	γ_1	γ_2	γ_3	R^2
Canada Spline	0.18	−1.32	3.26	−1.77	13.43
Canada NSS	−0.15	−1.05	2.11	−0.89	11.10
UK Spline	−1.33	−1.11	3.67	−2.26	18.02
UK NSS	−1.22	−1.01	3.41	−2.13	15.77

Note: This table presents estimates of unrestricted regressions for Canada and the UK using both Spline and Nelson–Siegel–Svensson (NSS) yield curves. NSS yields were constructing by fitting each country's Spline yield curves with the parametric technique of Svensson (1994).

Table 6
Comparison of predictability with Kessler and Scherer (2009).

	AUS	CAN	GER	JAP	NZ	CH	SE	UK	US	US FB
Feb/97-Jul/07 R^2	72.85	56.23	47.70	68.55	61.11	56.29	56.75	47.95	42.00	54.96
Full Sample R^2	16.50	13.96	13.06	28.55	49.33	37.47	27.25	17.98	15.70	24.37

Note: This table presents the R^2s from unrestricted regressions (5) for the 5-year zero coupon bond, but with their respective sample as in Kessler and Scherer (2009) beginning in February 1997 and ending in July 2007, and their R^2s with the full samples beginning at their respective dates and ending May 2009. The last column, US FB, reports the same regressions using the Fama-Bliss dataset. The R^2 values for the subsample (February 1997 to July 2007) are broadly in line with those obtained by Kessler and Scherer (2009).

Table 7
% of Cross-country variance due to first principal component.

	y_{12}	f_{36}	f_{60}	*CP*
(i) % Variance	64.4	86.5	90.5	57.5
(ii) % Variance	86.0	93.3	94.5	59.9

Note: Principal components decomposition of yields, forward rates and CP factors for every country in the sample. The first row (i), includes every country in the sample, but Norway, which is excluded because of its limited sample. The common sample starts in December/1993. In the second row (ii), I restrict the sample to the US., UK, and Germany. In this case, the common sample starts in January/1980.

4.1. How common are CP factors?

Several papers suggest the existence of strong contemporaneous cross-linkages in yield curves.[11] Nevertheless, the existence of global yield curve factors should not necessarily be taken as evidence for the existence of a global CP factor, since the latter is independent of traditional yield curve factors and cannot be summarized by them, as discussed in Cochrane and Piazzesi (2005).

For all countries in the sample, Table 7 shows the proportion of cross-country variance of y_{12}, f_{36} and f_{60} that is explained by their first principal component. The first principal component explain about 64% of the variance of y_{12}. This number rises to 86.5 and 90.5% for the f_{36} and f_{60}, respectively. It also shows the proportion of the variance of CP factors explained by the first principal component is the smallest, at 57.5%. In a sample with countries for which the time series are the longest (US., UK, Germany and Japan), we see that the proportion of variance explained by the first principal component rises considerably for y_{12}, f_{36} and f_{60}, but only marginally for CP. Excess return forecasting factors seem therefore to be considerably more idiosyncratic than yield curves or forward rates.

4.2. The impact of the recent financial crisis

Cochrane and Piazzesi (2005) report that the return forecasting factor for the US is countercylical. In my sample, the return forecasting factors appear to rise at the beginning of recessions, as in Sweden and Germany in the early 1990's, Japan throughout the 1990 decade, and the US. in the Volcker disinflation and the 1990–1991 and 2001 recessions. However, the current recession seems to be an exception. Fig. 3 shows the estimated CP factors. For most countries in the sample, the return forecasting factors have not experienced sharp rises during the current financial crisis period, as would be expected given the size of the current global recession. As conjectured by Wright (forthcoming), who estimates a similar drop in risk premiums for most countries in the same sample,[12] this may be the result of a series of events brought about by the financial crisis, as quantitative easing by a number of Central Banks, flight-to-quality by investors, and the zero bound of interest rates, a situation where overnight rates can only move in one direction.

[11] See, for example, Diebold et al. (2008), Perignon et al. (2007), among others.

[12] For a complete account of the effects of the current financial crisis in the US fixed income markets, see Dwyer and Tkac (2009). Nippani and Smith (2010) present and analysis of the crisis impact on default risk on US long term bonds, and Baba and Packer (2009) show evidence of malfunction of the exchange rate market around this same period.

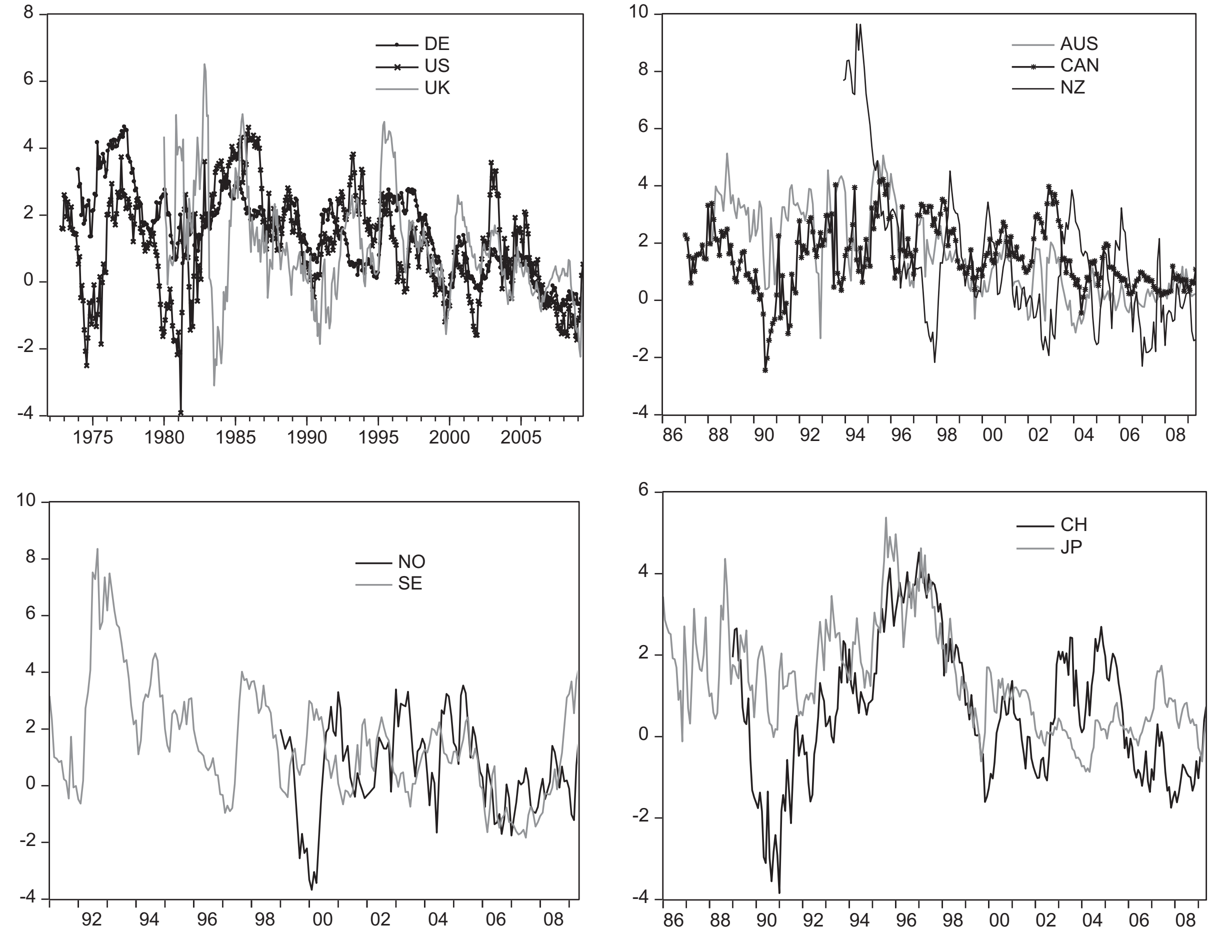

Fig. 3. CP factors of individual countries. *Notes*: This figures display the bond excess return forecasting factors (CP factors), given by $(CP_{t-12,t} = \hat{\gamma}_0 + \hat{\gamma}_1 y_{12,t-12} + \hat{\gamma}_2 f_{36,t-12} + \hat{\gamma}_5 f_{60,t-12})$ as in Eq. (3) for every country in the sample.

Table 8
Predictability of excess bond returns and the financial crisis.

	AUS	CAN	GER	JAP	NOR	NZ	CH	SE	UK	US
Restricted Sample R^2	17.45	18.69	16.37	27.55	41.65	57.32	42.71	33.33	24.30	22.49
Full Sample R^2	14.47	13.43	12.63	26.70	28.15	48.27	35.12	25.48	18.02	14.48
Chow Forecast Test	1.68	2.41	1.59	0.22	3.23	3.91	2.68	1.82	2.50	4.19
p-value	0.03	0.00	0.04	0.99	0.00	0.00	0.00	0.02	0.00	0.00

Note: This table presents the R^2s from unrestricted regressions (5), but with their respective sample ending in December/2007, before the financial crisis, and their R^2s with the full sample ending May/2009. The Chow Forecast Test for structural break in the parameters, and their respective *p*-values are also shown, where the break date is August of 2007.

In this section, I analyze the extent to which the recent financial crisis has impacted the predictability of excess bond returns in my sample. I estimate the unrestricted predictive regressions of Eq. (1) with a sample ending in August 2007, before the current crisis, and compare the results with the ones estimated with the full sample ending in May 2009.

I find a higher degree of predictability before the crisis for all countries in my sample, except Japan. Table 8 shows the R^2s estimates of Eq. (1) before and after the crisis. On average, R^2s are about one third higher before the crisis. Understandably, the change in R^2s are bigger for the countries with smaller samples, as Norway and Sweden, but predictability is affected even among the countries with the biggest samples, as the US. and Germany. The average excess bond return is forecasted with a R^2 of 22% in the US. before the crisis, whereas this value falls to 14% after the crisis period is included in the sample. In order to test the stability of the parameters before and after the crisis, I apply Chow's forecast test, with the break date set at August 2007. I reject the hypothesis of parameter stability for all the countries in my sample, except Japan. These results bring further evidence that the current financial crisis has indeed affected the predictability of excess bond returns.

5. Conclusion

This paper shows that the CP excess bond return forecasting factor performs well in international data: a single linear combination of forward rates forecasts excess bond returns of every maturity in all countries studied. I find significant evidence of predictability for every country, with some evidence of heterogeneity across countries. In the majority of cases, the single return forecasting factor also displays a tent-shape form. Nevertheless, the levels of predictability are significantly lower than the ones estimated by Cochrane and Piazzesi (2005) for US. bond yields.

I find that CP factors exhibit considerably less common variation than yields and forward rates. Contrary to the more traditional yield curve factors of slope and curvature, whose dynamics have been shown by Diebold et al. (2008) to be greatly influenced by global factors, CP factors seem to be driven more strongly by country idiosyncratic factors.

Excess return forecasting factors appear to be counter-cyclical, rising in recessions and falling in expansions. One big exception seems to be the current recession, where the factors have actually decreased in a number of countries as a probable consequence of the quantitative easing monetary policies followed by most of central banks in the sample to address the current financial crisis, as well as flight-to-quality by investors and the extremely low or zero overnight interest rates in some of the countries in the sample. As a consequence of those facts, I show that the predictability of excess bond returns by means of a linear combination of forward rates was significantly affected for the vast majority of countries in the sample.

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